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Free Trade Agreements *In* the Americas: Are the Trade Effects Larger than Anticipated?

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1. INTRODUCTION

IN the wake of the collapse of the much-publicised 'Free Trade Agreement *of* the Americas' in 2005, observers typically cite only the North American Free Trade Agreement (NAFTA) and the Mercado Común del Sur (Mercosur) as evidence of regional free trade agreements (FTAs) in the Western Hemisphere. However, the 'Americas' have many more FTAs – both at the bilateral level and the regional level. As of 2000, the Americas can boast at least six regional FTAs in operation currently and as many as 17 bilateral FTAs, although four of the 17 are between Western Hemisphere (WH) economies with non-WH economies (USA-Israel, Canada-Israel, Mexico-Israel and Mexico-European Union). Two more major FTAs have been signed in 2007 (though not yet implemented) – a regional FTA between the United States with five (of six) Central American countries and the Dominican Republic (DR-CAFTA) and the US-Korea FTA. As in Europe, regional FTAs and common markets began around 1960. The two oldest (effective agreements) in the Americas are the Central American Common Market or CACM (which began in 1961 as an FTA) and the Caribbean Community and Common Market or CARICOM (which began in 1966 as an FTA). As in the rest of the world, the vast bulk of the economic integration agreements (EIAs) in the Americas emerged in the past 20 years. In this paper, we examine the *ex post* effects of EIAs in the Americas that were formed between 1960 and 2000.

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This paper has three overarching goals. First, in most cases – certainly for NAFTA and Mercosur – policy makers in the regions estimated *ex ante* the trade and national output effects of forming these trade agreements using computable general-equilibrium (CGE) models. However, we conjecture that policy makers should – and *likely do* – want to have (if feasible) *ex post* estimates of the effects of trade agreements on their members' trade (and, subsequently, output, consumption, etc.). This paper uses panel econometric methods and data to evaluate *ex post* the trade impacts of all 19 bilateral and regional trade agreements within the Americas, the three North American countries' FTAs with Israel, the Mexican-EU FTA, and *all other EIAs* in existence between 1960 and 2000 for our sample of nearly 100 countries (the latter captured in one summary effect). Second, this is not the first paper to provide *ex post* estimates of the trade effects of NAFTA and Mercosur. While both agreements are fairly young (NAFTA phased-in beginning in 1994 and Mercosur in 1991), most *ex post* estimates of these two agreements' effects on their members' trade have been small. However, we will argue that previous approaches have tended to *underestimate* the trade impacts of these two agreements – *and others*, likely owing to previous approaches tending to ignore the 'self-selection' of countries into FTAs due to an emerging environment of 'competitive liberalisation' (creating a downward endogeneity bias) and tending to ignore recent developments in the theoretical foundations for the gravity equation to generate these estimates. Third, based upon these estimates, we will argue that countries' policy makers have self-selected into FTAs increasingly in the past two decades on the expectation that the effects of FTAs would be much larger (*ex post*) than suggested by the (*ex ante*) CGE estimates generated prior to the agreements, partly because *ex post* analyses do not require detailed measurement of the 'trade-policy liberalisations', which are often complicated and not easily quantified.

2. ECONOMIC INTEGRATION AGREEMENTS IN THE AMERICAS

As is the case around the world, not all 'trade agreements' between nations are FTAs. The term 'Economic Integration Agreement' (EIA) will be used in this paper rather than FTA to be inclusive. The term 'Economic Integration' spans integration of goods, services, capital and labour markets. All of the agreements we will examine are at least 'free' trade agreements, in the sense that we will not include preferential trade agreements where liberalisation was not intended to be full (partial liberalisation or one-way liberalisation was intended). However, many of the agreements have deeper integration than FTAs, such as customs unions (with a common external tariff) or common markets (with liberalisation in the movement of capital and labour). For instance, several countries of Central America formed the Central American Economic Integration Agreement (CAEIA),

one of the first post-WWII FTAs in 1961, which became a customs union in 1966 (CACM). The Caribbean Community and Common Market (CARICOM) is actually a common market with mobility of labour and capital. We will also use the term 'Economic Integration' – not 'Regional Economic Integration' – because some of the agreements we will evaluate cross different continents. In this paper, the bilateral EIAs between Israel and Canada, Mexico and the United States and between Mexico and the European Union are 'unnatural' in the sense that they cross different continents.

In the Americas, there are presently six 'regional' (not bilateral) EIAs. The oldest is the Central American Common Market, which began as a free trade agreement (for three of five countries) in 1961, but broadened to five countries and deepened into a customs union (not really a common market) in 1966. The agreement, by most accounts, became moribund between 1976 and 1990 due to the El Salvador-Honduras war that began in 1969 and armed conflicts in the region in the 1980s. However, in 1990, the agreement was revived by the nations' presidents (CACM2). CACM will refer to the customs union that prevailed until the mid-1970s and CACM2 will refer to the present agreement that was revived in 1990. The second oldest agreement is the Caribbean Community and Common Market (CARICOM), which started as the Caribbean Free Trade Association (CARIFTA) in 1966 and then expanded in members over the years and deepened into a common market in 1973. In 1969, a group of countries sharing the borders of the Andes created the Andean Group. The group had a cooperation agreement for many years, with the first full year of an actual customs union in 1995. In 1991, Argentina, Brazil, Paraguay and Uruguay signed the Asunción Treaty, forming the Mercado Común del Sur (Mercosur), which transformed into a customs union in 1995. The year 1992 saw the signing of the North American Free Trade Agreement (NAFTA), which began implementation of free trade among Canada, Mexico and the United States in 1994, on the heels of the 1989 bilateral Canadian-US FTA. The sixth regional trade agreement in the Americas is known as the Group of Three (Colombia, Mexico and Venezuela). This FTA went into effect in 1995. These were the only regional EIAs established between 1960 and 2000, the period of our analysis. The only EIA not mentioned is the Latin America Free Trade Agreement, for which the Treaty of Montevideo was signed in 1960. However, after 18 years the FTA never really surfaced. In 1978, the participants restructured a more flexible sectoral-based preferential trade agreement that became known as the Latin American Integration Agreement (LAIA). However, this agreement never had the scope or depth of (what most observers would consider) an FTA.

The analysis in this paper will also examine the *ex post* effects of numerous bilateral EIAs in the Americas (as well as three between the three North American countries and Israel and the Mexican-EU FTA). These include (with the date of entry into force in parentheses): El Salvador-Panama (02/14/1974), Guatemala-Panama (04/25/1975), Israel-USA (08/19/1985), Colombia-Costa Rica (09/02/1985),

TABLE 1
Economic Integration Agreements in the Americas, 1960–2000

Acronym	*Full Name*	*Date it Entered into Force*
CAEIA	Central American Economic Integration Agreement	06/03/1961
CACM	Central American Common Market (replaced CAEIA)	01/01/1966
CARIFTA	Caribbean Free Trade Agreement	12/10/1966
CARICOM	Cari. Community & Common Market (replaced CARIFTA)	08/01/1973
CACM2	Central America Common Market (re-initiated)	06/30/1990
MERCOSUR	Mercado Común del Sur	03/26/1991
NAFTA	North American Free Trade Agreement	01/01/1994
G3	Group of Three: Colombia-Mexico-Venezuela FTA	01/01/1995
ANDEAN_COM	Andean Community Customs Union	02/01/1995
SAL_PAN	El Salvador-Panama Bilateral Free Trade Agreement	02/14/1974
GUAT_PAN	Guatemala-Panama Bilateral Free Trade Agreement	04/25/1975
USA_ISR	United States-Israel Bilateral Trade Agreement	08/19/1985
COS_COL	Costa Rica-Colombia Bilateral Free Trade Agreement	09/02/1985
DOM_PAN	Dominican Republic-Panama Bilateral FTA	06/08/1987
USA_CAN	United States-Canada Bilateral Free Trade Agreement	01/01/1989
BOL_MEX	Bolivia-Mexico Bilateral Free Trade Agreement	01/01/1995
COS_MEX	Costa Rica-Mexico Bilateral Free Trade Agreement	01/01/1995
MER_CHIL	Mercosur-Chile Trade Agreement	10/01/1996
CAN_ISR	Canada-Israel Bilateral Free Trade Agreement	01/09/1997
MER_BOL	Mercosur-Bolivia Trade Agreement	03/02/1997
CHIL_CAN	Chile-Canada Bilateral Free Trade Agreement	07/05/1997
MEX_NIC	Mexico-Nicaragua Bilateral Trade Agreement	07/01/1998
DOM_CARIC	Dominican Republic-Caribbean Community FTA	08/22/1998
CHIL_MEX	Chile-Mexico Bilateral Free Trade Agreement	08/01/1999
MEX_EU	Mexico-European Union Free Trade Agreement	07/01/2000
MEX_ISR	Mexico-Israel Bilateral Free Trade Agreement	07/01/2000

Dominican Republic-Panama (06/08/1987), Canada-USA (01/01/1989), Bolivia-Mexico (01/01/1995), Costa Rica-Mexico (01/01/1995), Chile-Mercosur (10/01/1996), Canada-Israel (01/09/1997), Bolivia-Mercosur (03/02/1997), Chile-Canada (07/05/1997), Mexico-Nicaragua (07/01/1998), CARICOM-Dominican Republic (08/22/1998), Chile-Mexico (08/01/1999), Mexico-EU (07/01/2000) and Mexico-Israel (07/01/2000). Virtually all of these agreements are FTAs. Table 1 summarises the EIAs in the Americas.

3. DETERMINANTS OF BILATERAL TRADE FLOWS AND THE POTENTIAL TRADE EFFECTS OF EIAs

In this section, we discuss conceptually the overall methodological approach for estimating *ex post* the effects of EIAs on bilateral trade flows of members.

There are three major points we wish to emphasise. First, we distinguish our *ex post* approach from traditional *ex ante* analyses. The international trade literature has long relied upon computable general-equilibrium (CGE) models to evaluate *ex ante* the potential trade, output, consumption and welfare effects of forming EIAs. However, policy makers should – and, we conjecture, do – want *ex post* estimates of the effects of FTAs on trade, etc. That is, they want measures of whether such agreements actually 'work'. The approach long used for *ex post* empirical analysis of EIAs has been the 'gravity equation', which we discuss below. One of the advantages of this approach is that it does not require detailed measurement of the 'liberalisations'; by construction, both 'border' barriers (i.e. quantifiable tariffs) and the difficult-to-quantify 'non-border' barriers (i.e. non-tariff barriers, regulations, etc.) are captured using binary variables. Second, while the gravity equation has been used for decades to try to deliver such *ex post* estimates, a recurring problem is that such estimates are often seemingly implausible and quite fragile. We address econometric methods that suggest more 'sturdy' and plausible estimates. Third, ideally one would want to estimate the full (general-equilibrium comparative-static) effect of an EIA on two countries' trade after accounting for all relative price changes as in typical *ex ante* CGE models, and not just the partial effect on the two countries' trade (one which ignores changes in *other* prices that potentially can have feedback effects on the bilateral trade flow of the two members). Such issues were raised recently in Anderson and van Wincoop (2003) in the context of the gravity equation; we address these issues later in Section 5.

a. Determinants of Bilateral Trade

It will be useful first to discuss the underlying economic context of world trade *in the absence* of policy-oriented barriers to trade. After we establish the fundamental determinants of trade in the presence of only 'natural' barriers to trade (e.g. distance between economic agents), we then introduce (exogenously) policy-oriented – or 'artificial' – trade barriers. This will provide the background to then discuss *endogenous* regionalism behaviour by governments, which likely has biased most previous *ex post* estimates of trade effects of EIAs.

We address briefly determinants of bilateral trade flows in an N-country world ($N > 2$) in the absence (presence) of policy-based (natural) trade barriers. The modern theory of international trade – largely developed in the context of two countries with production of goods in two industries using two factors of production – usually emphasises that the economic rationales for international trade are 'traditional' comparative advantage (or inter-industry trade, driven by Heckscher-Ohlin relative factor endowment differences or Ricardian relative productivity differences) and 'acquired' comparative advantage (or intra-industry trade, due to increasing returns to scale in production of slightly differentiated products), but historically ignoring transport costs and economic geography.

However, motivated by the robust empirical regularity that bilateral trade flows between pairs of countries are explained well by the product of their gross domestic products (GDPs) and their bilateral distance, trade economists have formulated over the past 25 years multi-country (or *N*-country) theoretical foundations for a 'gravity equation' of bilateral international trade, and in a manner consistent with established theories of intra- and inter-industry international trade. For instance, the first formal theoretical foundation for the gravity equation with a one-sector endowment economy, but many countries, was Anderson (1979). Anderson showed that a simple (conditional) general-equilibrium Armington model with products differentiated by country of origin and constant-elasticity-of-substitution preferences yields a basic gravity equation:[1]

$$\mathrm{PX}_{ij} = \beta_0(\mathrm{GDP}_i)^{\beta_1}(\mathrm{GDP}_j)^{\beta_2}(\mathrm{DIST}_{ij})^{\beta_3}\varepsilon_{ij}, \tag{1}$$

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 centres of countries i and j, and ε_{ij} is assumed to be a log-normally distributed error term. The theory suggested that $\beta_1 = \beta_2 = 1$ and $\beta_3 < 0$.

Other papers extended these theoretical foundations in various important directions. Helpman and Krugman (1985) introduced monopolistic competition and increasing returns to scale, motivating a gravity equation (without trade costs) to explain intra-industry trade between countries with similar relative factor endowments and labour productivities. Bergstrand (1985) motivated theoretically and introduced econometrically (crude) proxies for multilateral price terms for importers and exporters as important for determining bilateral trade flows; for instance, the trade flow from i to j is influenced by the prices, transport costs and other trade costs that the consumer in j faces from its $N-2$ other trade partners as well as domestic firms. Bergstrand (1989 and 1990) showed formally that a gravity equation evolved from a traditional Heckscher-Ohlin model with two industries, two factors and N countries with both inter-industry and (Helpman-Krugman) intra-industry trade. Evenett and Keller (2002) provided empirical evidence that a model with both Heckscher-Ohlin inter-industry trade and Helpman-Krugman intra-industry trade with imperfect specialisation fit the data best. Eaton

[1] The theoretical gravity equation (16) with trade costs in Anderson (1979) was more complicated; however, equation (1) could be obtained under certain further assumptions described there. As noted in Anderson and van Wincoop (2004), Anderson (1979) and Anderson and van Wincoop (2003) are 'conditional' general-equilibrium models, employing a 'trade separability' assumption where the allocation of bilateral flows across N countries is separable from production and consumption allocations within countries.

and Kortum (2002) formulated a theoretical foundation for the gravity equation based upon a Ricardian model with a continuum of goods.

Most recently, Anderson and van Wincoop (2003) have shown formally that proper estimation of the gravity equation (to avoid omitted variables bias) should recognise that multilateral price terms for both the exporter and importer countries are endogenous. They showed that estimation of the system of non-linear equations (2)–(4) below using custom non-linear least squares programming could account properly for the endogeneity of prices:

$$\mathrm{PX}_{ij} = \beta_0(\mathrm{GDP}_i)^1(\mathrm{GDP}_j)^1(t_{ij})^{1-\sigma}P_i^{\sigma-1}P_j^{\sigma-1}\varepsilon_{ij}, \tag{2}$$

where $\sigma > 1$, t_{ij} denotes bilateral trade costs (which potentially can be explained by various observable variables) and P_i and P_j are 'endogenous' multilateral price terms that account for trade costs that agents in countries i and j face from all N countries (including at home), where:

$$P_i = \left[\sum_{j=1}^{N} \theta_j (t_{ij}/P_j)^{1-\sigma}\right]^{1/(1-\sigma)} \tag{3}$$

$$P_j = \left[\sum_{i=1}^{N} \theta_i (t_{ij}/P_i)^{1-\sigma}\right]^{1/(1-\sigma)} \tag{4}$$

under an assumption that bilateral trade barriers t_{ij} and t_{ji} are symmetric for all pairs. Letting GDP^T denote total income of all regions, which is constant across region pairs, then θ_i (θ_j) denotes $\mathrm{GDP}_i/\mathrm{GDP}^T$ ($\mathrm{GDP}_j/\mathrm{GDP}^T$). Details of estimating equations (2)–(4) for aggregate trade flows using non-linear least squares or equation (2) using fixed effects for P_i and P_j are addressed in Anderson and van Wincoop (2003), Feenstra (2004), and Baier and Bergstrand (2002, 2006 and 2007).[2] Baier and Bergstrand (2002) extend the Anderson-van Wincoop one-sector, N-country endowment economy to a world with two sectors, two factors and N countries with Heckscher-Ohlin inter-industry trade and Helpman-Krugman intra-industry trade; cf. Carrere (2006).

[2] See Anderson and van Wincoop (2004) for an excellent survey of the literature on theoretical foundations for the gravity model. In Anderson (1979), all prices were normalised to unity. In Bergstrand (1985, 1989 and 1990), a 'small-country' assumption was employed to treat the other $N-1$ countries' price levels as exogenous to the country pair ij. In Anderson and van Wincoop (2003) all countries' price levels are endogenous. Also, see Evenett and Hutchinson (2002) for a volume of papers on gravity equation methodology.

b. Using the Gravity Equation to Estimate Ex Post the Effects of EIAs on Members' Trade

The gravity equation in specification (1) has been used traditionally for about 40 years to explain the variation in bilateral trade flows among pairs of countries for a particular year and more recently for panel variation (especially, within variation using fixed effects; cf. Egger, 2000 and 2002). Typically, several other binary variables are included to capture variation in various trade costs, such as an adjacency dummy and a language dummy. More relevant here, most researchers have included a dummy variable for the presence or absence of an EIA. Quantitative estimates of the coefficients of these EIA dummies have varied dramatically, cf. Frankel (1997), with some estimated average 'treatment' effects seemingly small and others even negative. 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.[3]

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.[4] This paper applies developments in the

[3] 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 and 1989) and Boisso and Ferrantino (1997). However, no systematic explanation is provided.

[4] 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.

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-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 labour economics studies (cf. Wooldridge, 2002), provides rich tools that have not previously been used for analysing 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 US multilateral imports and US multilateral non-tariff 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 US imports increased *tenfold*. Lee and Swagel (1997) also showed using instrumental variables that previous estimates of the impact of trade liberalisation 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 underestimated, 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 and 2004a) 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.

To support our claim that estimates of the impact of EIAs may be biased, we provide coefficient estimates from a typical cross-section gravity equation for multiple years: 1960, 1970, 1980, 1990 and 2000. These coefficient estimates come from a typical log-linear version of equation (1) amended to include dummy variables for common land border (adjacency), common language and common membership in various EIAs, estimated using the (non-zero) nominal trade flows among the 96 countries identified in the Data Appendix. These estimates are derived including separate EIA dummy variables for all 23 regional and bilateral EIAs in the Americas described in Section 1 *and* one separate dummy variable to capture *all 'other' EIAs* (*OtherEIA*). For instance, $Mercosur_{ijt}$ is defined to equal 1 if a country pair *ij* in year *t* were members of Mercosur, and 0 otherwise. $OtherEIA_{ijt}$ is defined as 1 if country pair *ij* in year *t* had any other EIA agreement.

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 centres to calculate the great circle distances. The language and adjacency dummy variables were also compiled from the *CIA Factbook*. The EIA dummy variables were calculated using appendices in Lawrence (1996) and Frankel (1997), various websites, and EIAs notified to the GATT/WTO under GATT Articles XXIV or the Enabling Clause for developing economies; we included only full (no partial) EIAs. Table 2 lists the trade agreements existing between 1960 and 2000 for the 96 countries in our sample.[5]

Table 3 reports cross-sectional estimates of typical gravity equations for various years. The EIAs' coefficient estimates are reported in six 'groups'. The first group includes regional EIAs formed in the 1960s and 1970s. The second group includes regional EIAs formed in the 1990s; none were formed in the 1980s (only bilateral agreements were formed). The third group is *bilateral* EIAs formed in the 1970s. The fourth group includes bilateral EIAs formed in the 1980s. The fifth group includes bilateral EIAs formed in the 1990s. The sixth 'group' has one variable: the dummy representing 'all other EIAs'.

As Table 3 shows, there is both considerable variability and lack of feasibility across the EIA 'average treatment effect' estimates (i.e. the coefficient estimates for the EIA dummies). For instance, common membership in CARICOM had an effect of increasing members' trade by 1980 by 11,230 per cent ($e^{4.73} = 113.3$, implying an 11,230 per cent increase [$= (113.3 - 1) \times 100$]), which seems implausible. On the other hand, by 1980 all other EIAs in the sample of 96 countries (other than in the Americas) had *decreased* trade by 17 per cent ($e^{-0.19}$), which also seems implausible. Moreover, the range of estimates suggests very 'fragile' estimates, in line with the concerns raised in Ghosh and Yamarik (2004).

As discussed earlier, typical gravity equation (1) is likely mis-specified owing to ignoring theoretical foundations that have developed over the past several decades. Table 4 provides estimates of theoretically motivated gravity equation (2) using (as is now common) country-specific fixed effects to account for the variation of multilateral price terms P_i and P_j in equation (2) to avoid omitted variables bias and restricting the coefficient estimates for GDPs to be unity (as suggested by theory). As Table 4 reports, accounting for the theoretically-motivated

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

TABLE 2
Economic Integration Agreements among 96 Countries in the Sample, 1960–2000

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 1945), 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 re-initiated 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

Caribbean 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): Colombia, Mexico, Venezuela

Mercado Común del Sur, or Mercosur (1991): Argentina, Brazil, Paraguay, Uruguay (formed in 1991 and a free trade area in 1995)

Andean Community (1993): Bolivia, Colombia, Ecuador, 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 per cent of merchandise trade in 1998)

TABLE 2 *Continued*

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)
3
Common Market for Eastern and Southern Africa (2000): Egypt, Kenya, Madagascar, Malawi, Mauritius, Sudan, Zimbabwe, Zambia
EU-Israel Agreement (2000)
a
EU–Mexico (2000)
Poland–Turkey (2000)
Mexico–Guatemala (2000)
Mexico–Honduras (2000)
Mexico–Israel (2000)
Mexico–El Salvador (2000), New Zealand–Singapore (2000)
Romania–Turkey (1998)
Mexico–Chile (1999)
Common Market for Eastern and Southern Africa (2000): Egypt, Kenya, Madagascar, Malawi, Mauritius, Sudan, Zimbabwe, Zambia
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)

Notes:
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_esummary_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 (1997).
Lawrence (1996).

TABLE 3
Typical Cross-section (atheoretical) Gravity Equation Coefficient Estimates

Variable	*(1) 1960*	*(2) 1970*	*(3) 1980*	*(4) 1990*	*(5) 2000*
ln GDP_i	0.76 (47.06)[c]	0.89 (57.90)[c]	1.02 (70.26)[c]	1.09 (85.48)[c]	1.19 (104.68)[c]
ln GDP_j	0.76 (50.16)[c]	0.92 (64.28)[c]	1.01 (74.44)[c]	0.97 (78.42)[c]	0.98 (87.97)[c]
ln $DIST_{ij}$	−0.65 (−16.81)[c]	−0.84 (−21.15)[c]	−1.05 (−27.95)[c]	−1.06 (−28.57)[c]	−1.15 (−31.71)[c]
ADJ_{ij}	0.15 (1.00)	0.15 (0.90)	0.43 (2.71)[c]	0.64 (4.06)[c]	0.71 (4.50)[c]
$LANG_{ij}$	0.05 (0.57)	0.25 (2.54)[b]	0.47 (4.95)[c]	0.72 (7.26)[c]	0.53 (5.35)[c]
CAEIA					
CACM		3.10 (7.17)[c]			
CARIFTA		3.32 (2.67)[c]			
CARICOM			4.73 (6.09)[c]	4.49 (5.49)[c]	4.89 (6.06)[c]
CACM2					2.31 (5.06)[c]
MERCOSUR					1.04 (1.80)[a]
NAFTA					−0.01 (−1.01)
G3					−0.60 (−0.74)
ANDEAN_COM					1.09 (2.38)[c]
SAL_PAN			2.40 (1.80)[a]	1.84 (1.31)	2.74 (1.97)[b]
GUAT_PAN			1.71 (1.28)	2.34 (1.66)[a]	2.12 (1.53)
USA_ISR				1.37 (0.98)	1.20 (0.87)
COS_COL				−0.22 (−0.16)	0.31 (0.22)
DOM_PAN				2.07 (1.47)	3.11 (2.24)[b]
USA_CAN				−0.66 (−0.47)	−0.72 (−0.42)
BOL_MEX					0.02 (0.01)
COS_MEX					0.58 (0.42)
MER_CHIL					0.55 (0.79)
CAN_ISR					0.66 (0.47)
MER_BOL					0.14 (0.20)
CHIL_CAN					1.40 (1.01)
MEX_NIC					0.89 (0.64)
DOM_CARIC					2.13 (2.43)[b]
CHIL_MEX					1.14 (0.82)
MEX_EU					−0.55 (−1.46)
MEX_ISR					−0.39 (−0.28)
OtherEIA	0.61[c] (3.40)	0.84[c] (3.63)	−0.19 (−1.12)	−0.18 (−1.22)	0.13 (1.15)
Constant	−10.16[c] (−21.81)	−14.28[c] (−30.60)	−17.35[c] (−38.07)	−18.54 (−43.64)	−20.10 (−51.54)
RMSE	1.4139	1.7567	1.8872	1.9871	1.9569
R^2	0.6035	0.6356	0.6471	0.6662	0.7168
No. observations	2,789	4,030	5,494	6,474	7,302

Notes:
t-Statistics are in parentheses. a, b and c denote statistical significance at the 10 per cent, 5 per cent and 1 per cent levels, respectively, using two-tailed *t*-tests.

TABLE 4
Theory-motivated Cross-section Gravity Equations with Country Fixed Effects

Variable	*(1) 1960*	*(2) 1970*	*(3) 1980*	*(4) 1990*	*(5) 2000*
ln $DIST_{ij}$	−0.70 (−16.86)[c]	−0.87 (−20.88)[c]	−1.27 (−31.02)[c]	−1.29 (−30.10)[c]	−1.46 (−36.59)[c]
ADJ_{ij}	0.31 (2.33)[b]	0.37 (2.46)[b]	0.51 (3.31)[c]	0.63 (4.17)[c]	0.50 (2.84)[c]
$LANG_{ij}$	0.36 (3.68)[c]	0.76 (7.31)[c]	0.73 (6.94)[c]	0.91 (8.95)[c]	0.80 (8.12)[c]
CAEIA					
CACM		3.50 (14.39)[c]			
CARIFTA		2.51 (8.24)[c]			
CARICOM			4.26 (10.48)[c]	3.93 (9.85)[c]	3.91 (6.47)[c]
CACM2					1.76 (5.38)[c]
MERCOSUR					1.36 (3.91)[c]
NAFTA					0.36 (0.88)
G3					0.41 (1.32)
ANDEAN_COM					1.51 (5.56)[c]
SAL_PAN			1.45 (2.87)[c]	1.74 (6.12)[c]	2.47 (4.13)[c]
GUAT_PAN			0.90 (2.23)[b]	1.54 (4.42)[c]	1.24 (2.97)[c]
USA_ISR				1.25 (6.05)[c]	1.11 (3.98)[c]
COS_COL				−0.32 (−0.42)	0.81 (1.12)
DOM_PAN				1.87 (1.30)	3.38[b] (2.40)
USA_CAN				−1.25 (−6.14)[c]	−1.45 (−3.38)[c]
BOL_MEX					0.92 (1.94)[a]
COS_MEX					0.77 (2.33)[b]
MER_CHIL					−0.31 (−0.74)
CAN_ISR					0.32 (1.77)[a]
MER_BOL					0.68 (1.54)
CHIL_CAN					0.81 (4.24)[c]
MEX_NIC					0.53 (0.96)
DOM_CARIC					3.05 (5.11)[c]
CHIL_MEX					0.78 (2.82)[c]
MEX_EU					−0.09 (−0.39)
MEX_ISR					0.00 (0.00)
OtherEIA	0.00 (0.01)	−0.17 (−1.06)	−1.54 (−11.99)[c]	−1.14 (−10.46)[c]	−0.55 (−5.49)[c]
Constant	−16.61 (−5.49)[c]	−14.52 (−16.35)[c]	−16.68 (−20.94)[c]	−17.21 (−24.31)[c]	−14.50 (−21.98)[c]
RMSE	1.183	1.4893	1.6613	1.7761	1.7738
R^2	0.5002	0.4403	0.3888	0.3687	0.3942
No. observations	2,789	4,030	5,494	6,474	7,302

Notes:
t-Statistics are in parentheses. a, b and c denote statistical significance at the 10 per cent, 5 per cent and 1 per cent levels, respectively, using two-tailed *t*-tests.

multilateral price terms does not improve the results for EIA effects relative to Table 3. For instance, the CARICOM coefficient estimate in 1980 is still implausibly large and *OtherEIAs*' coefficient estimate in 1980 is a large negative value.[6]

The reason why the EIA variables' coefficient estimates may be biased is perhaps due to the *endogenous* determination of EIAs in a competitive environment. For instance, in equations (1) or (2), the error term ε may be representing unobservable (to the empirical researcher) policy-related barriers tending to reduce trade between countries i and j that are not accounted for by standard gravity equation RHS variables, but may be correlated with the decision to form an EIA. Suppose two countries have extensive unmeasurable domestic regulations (say, internal shipping regulations) that inhibit trade (causing ε to be negative). The likelihood of the two countries' governments selecting into an EIA may be high if there is a large expected welfare gain from potential bilateral trade creation if the EIA deepens liberalisation beyond tariff barriers into domestic regulations (and other non-tariff barriers). Thus, $OtherEIA_{ijt}$ and the intensity of domestic regulations may be positively correlated in a cross-section of data, but the gravity equation error term ε_{ijt} and the intensity of domestic regulations may be negatively correlated. This suggests that $OtherEIA_{ijt}$ and ε_{ijt} are negatively correlated, and the *OtherEIA* coefficient estimate may be underestimated, as the evidence in Tables 3 and 4 indicates.

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 liberalising the domestic policies just named. As Lawrence states: 'Once tariffs are removed, complex problems remain because of differing regulatory policies among nations' (p. 7). He argues that in many cases, EIA 'agreements are also meant to achieve deeper integration of international competition and investment' (p. 7). Gilpin (2000) echoes this argument:

[6] It should be remembered throughout that the discussion of 'effects' of an EIA are limited only to the primary 'partial' effect associated with the dummy variable's coefficient estimates, and we are intentionally precluding from our discussion for now the full general-equilibrium comparative-static effects addressed in Anderson and van Wincoop (2003) and Baier and Bergstrand (2006). We will return to this topic later.

> 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).

However, as noted some of the coefficient estimates seem implausibly large, as is CARICOM's estimate. It is also possible for simultaneity to bias the coefficient estimates upward. Two countries' governments that trade more than their gravity-equation-suggested 'normal' level might be induced to form an EIA because there might potentially be less 'trade diversion' due to their extensive trading relationship, suggesting a positive simultaneity bias.

We believe the omitted variable (selection) bias is the major source of endogeneity facing estimation of EIA effects in gravity equations using cross-section data. Moreover, the arguments above suggest that policymakers' decisions to select into an EIA 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 *MERCOSUR, OtherEIA*, etc. are likely to be cross-sectional in nature.

4. ESTIMATING THE EFFECTS OF VARIOUS EIAs ON TRADE FLOWS USING PANEL DATA

With cross-section data, standard econometric techniques to address omitted variables (and selection) bias include estimation using instrumental variables and Heckman control functions. Only a small handful of studies in the past three years have attempted to do this; Baier and Bergstrand (2002) was the first. Of the few studies that have attempted to solve this dilemma using instrumental variables and other cross-section techniques, there has been little success; cf. Baier and Bergstrand (2007). The reason basically is that – in cross-section – it is virtually impossible in a convincing way to identify variables that are correlated with the EIA dummy variable and are uncorrelated with trade flows. That is, there are no observable variables to *identify* the respective equations.

However, some alternative techniques are available to address the problem. For example, if the decisions to form EIAs are 'slow-moving' – as they are likely to be – but trade flows are not slow moving (also likely), then panel data offers an opportunity to better identify unbiased effects of EIAs on trade flows. Bayoumi and Eichengreen (1997) pursued this using first-differences and Cheng and Wall (2005) using fixed effects, but both in the context of atheoretical gravity specifications with small samples.

Baier and Bergstrand (2007) used both approaches in the context of a theoretically-motivated gravity equation for a broad sample of countries and panel data. Starting from the conditional general equilibrium of Anderson and van

Wincoop (2003), Baier and Bergstrand (2007) motivated the panel version of the Anderson and van Wincoop gravity equation:

$$\ln[X_{ijt}/(\mathrm{RGDP}_{it}\mathrm{RGDP}_{jt})] = \beta_0 + \beta_3(\ln(\mathrm{DIST}_{ij}) + \beta_4(\mathrm{ADJ}_{ij}) + \beta_5(\mathrm{LANG}_{ij}) + \beta_6(\mathrm{EIA}_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \ln \varepsilon_{ijt}, \quad (5)$$

where X_{ijt} is the real (inflation-adjusted) trade flow from i to j in year t and RGDP_{it} is real GDP of country i in year t and EIA is used generically to represent the set of various EIA dummy variables.

Using fixed effects, Baier and Bergstrand (2007) find that the cumulative average treatment effect of an EIA on trade after 10–15 years is 0.76. Given that $e^{0.76}$ equals 2.14, this implies that an EIA on average increases two members' international trade by 114 per cent after 10–15 years, or roughly *5 per cent* annually over 15 years. This estimated (partial) effect is both considerably larger and more robust to sensitivity analyses than earlier estimates.

In this paper, we examine in particular the effects of membership in 23 explicit regional and bilateral EIAs in the Americas mentioned earlier, along with one overall dummy variable to capture membership in *all other EIAs*. Thus, in contrast to Baier and Bergstrand (2007) which treated the effects of all EIAs the same, this paper applies the *ex post* techniques of Baier and Bergstrand (2007) to examine several specific agreements, allowing here for changing membership over the 40-year period from 1960–2000. We have two goals in mind for the remainder of this analysis. First, we want to try to estimate *with precision* (and robustness) the *ex post* effects of various trade agreements in the Americas on members' international trade, accounting for the endogeneity of trade agreements' formation. Second, we want to establish that the economic effects of trade agreements in the Americas on members' trade were *much larger* than previous estimates have suggested, which will help to explain the proliferation of trade agreements in later years.

a. Alternative Panel Estimation Techniques: 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 3, 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 EIA and the volume of trade. Because these variables are likely correlated with $\mathrm{MERCOSUR}_{ij}$, NAFTA_{ij}, $\mathit{OtherEIA}_{ij}$, etc., they are best controlled for using bilateral 'fixed effects', as this approach allows for arbitrary correlations of w_{ij} with these variables. By contrast,

under 'random effects' one assumes zero correlation between unobservables w_{ij} with the various EIA dummy variables, which seems less plausible.

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

b. Fixed-effects Estimation of an Atheoretical Gravity Equation Ignoring Multilateral Price Terms

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

$$\ln X_{ijt} = \beta_0 + \beta_1(\ln \mathrm{RGDP}_{it}) + \beta_2(\ln \mathrm{RGDP}_{jt}) + \beta_3(\ln \mathrm{DIST}_{ij}) + \beta_4(\mathrm{ADJ}_{ij}) + \beta_5(\mathrm{LANG}_{ij}) + \beta_6(\mathrm{EIA}_{ijt}) + \varepsilon_{ijt}. \quad (6)$$

The essence of using a panel version of the gravity equation to extract the effect of an EIA on trade is seen by close scrutiny of equation (6). If the trade flow from i to j in year t is *actually* determined according to gravity equation (6), then the only variable with the same statistical properties as X_{ijt} is EIA_{ijt}. Holding constant RGDPs and the other three variables (or, as shown later, using ij fixed effects to account for these variables' variation) will allow unbiased estimation of β_6, which is the (partial) average treatment effect of an EIA.

Table 5 provides the empirical results of estimating gravity equation (6) using a panel of real trade flows (X_{ijt}), real GDPs (RGDP_{it}, RGDP_{jt}) and various EIA dummies, 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.

Similar to the Table 3 and 4 results, the coefficient estimates for the EIAs in column (1) of Table 5 have wide variation, and in many cases implausible values. CARICOM's effect on trade is over 7,300 per cent ($e^{4.31}$), which is again implausible. Numerous other EIAs have implausibly large effects, such as CACM and CARIFTA (the short-lived FTA precursor to CARICOM). On the other hand, the coefficient estimate for *OtherEIA* is negative and statistically significant, which seems implausible as well (since the EEC/EC/EU likely had strong effects).

Column (2) in Table 5 adds time dummies. For the most part the coefficient estimates change very little. In fact, some of the very large effects – such as for CARICOM and CACM – actually get larger. So several of the results remain implausible.

TABLE 5
Typical Panel (atheoretical) 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 $RGDP_i$	0.96 (218.42)[c]	0.98 (232.72)[c]	0.71 (34.53)[c]	1.27 (47.13)[c]
ln $RGDP_j$	0.95 (226.20)[c]	0.98 (237.09)[c]	0.58 (26.55)[c]	1.22 (41.57)[c]
ln $DIST_{ij}$	–1.03 (–78.16)[c]	–1.01 (–77.89)[c]		
ADJ_{ij}	0.41 (8.28)[c]	0.37 (7.20)[c]		
$LANG_{ij}$	0.52 (15.66)[c]	0.46 (13.87)[c]		
CAEIA	2.50 (20.90)[c]	1.83 (15.34)[c]	0.60 (1.94)[a]	0.30 (0.98)
CACM	2.85 (30.05)[c]	2.74 (29.81)[c]	1.16 (4.65)[c]	1.01 (4.09)[c]
CARIFTA	3.61 (46.16)[c]	3.39 (41.62)[c]	–0.14 (–0.15)[c]	0.19 (0.19)
CARICOM	4.31 (24.76)[c]	4.72 (28.57)[c]	0.27 (0.68)	0.95 (2.39)[b]
CACM2	1.90 (15.48)[c]	2.63 (21.00)[c]	0.61 (2.56)[b]	0.70 (2.98)[c]
MERCOSUR	0.79 (5.46)[c]	1.35 (9.03)[c]	0.73 (2.57)[b]	0.87 (3.11)[c]
NAFTA	0.61 (3.85)[c]	0.96 (5.54)[c]	1.21 (2.70)[c]	1.24 (2.81)[c]
G3	–0.23 (–1.56)	0.10 (0.56)	1.25 (3.04)[c]	1.28 (3.16)[c]
ANDEAN_COM	0.71 (4.92)[c]	1.34 (8.87)[c]	1.26 (4.65)[c]	1.43 (5.35)[c]
SAL_PAN	1.97 (8.17)[c]	2.33 (9.85)[c]	0.77 (1.26)	1.05 (1.74)[a]
GUAT_PAN	1.73 (11.29)[c]	2.17 (11.97)[c]	1.59 (2.50)[b]	1.68 (2.69)[c]
USA_ISR	1.42 (18.87)[c]	1.63 (19.68)[c]	0.54 (0.92)	0.54 (0.94)
COS_COL	–0.71 (–1.77)[a]	–0.24 (–0.55)	0.31 (0.53)	0.29 (0.50)
DOM_PAN	1.96 (2.60)[c]	2.57 (3.37)[c]	0.60 (0.90)	0.59 (0.90)
USA_CAN	–0.22 (–1.36)	–0.08 (–0.39)	–0.56 (–0.83)	–0.41 (–0.60)
BOL_MEX	–0.36 (–1.00)	0.21 (0.56)	1.12 (1.60)	1.27 (1.84)[a]
COS_MEX	0.11 (0.24)	0.68 (1.32)	0.98 (1.37)	0.86 (1.23)
MER_CHIL	0.27 (1.11)	0.94 (3.92)[c]	0.50 (1.09)	0.44 (0.97)
CAN_ISR	0.52 (4.71)[c]	1.04 (9.29)[c]	0.41 (0.45)	0.30 (0.33)
MER_BOL	–0.23 (–0.65)	0.52 (1.46)	0.45 (0.97)	0.67 (1.47)
CHIL_CAN	1.14 (12.24)[c]	1.67 (17.08)[c]	0.97 (1.05)	0.86 (0.95)
MEX_NIC	0.60 (1.03)	1.32 (2.25)[b]	1.63 (1.75)[a]	1.87 (2.04)[b]
DOM_CARIC	1.46 (3.11)[c]	2.17 (4.67)[c]	0.81 (1.34)	0.81 (1.36)
CHIL_MEX	1.11 (32.14)[c]	1.73 (42.74)[c]	1.39 (1.51)	1.21 (1.33)
MEX_EU	–0.47 (–2.85)[c]	0.03 (0.17)	0.47 (1.89)[a]	0.55 (2.22)[b]
MEX_ISR	–0.39 (–0.43)	0.13 (0.15)	1.12 (1.21)	0.94 (1.04)
OtherEIA	–0.09 (–2.41)[b]	0.01 (0.27)	0.42 (7.70)[c]	0.62 (11.43)[c]
RMSE	1.9201	1.8517		
Overall R^2	0.6602	0.6840		
Within R^2			0.2045	0.2277
No. observations	47,081	47,081	47,081	47,081

Notes:
t-Statistics are in parentheses. a, b and c denote statistical significance at the 10 per cent, 5 per cent and 1 per cent levels, respectively, using two-tailed *t*-tests.

Columns (3) and (4) in Table 5 include bilateral fixed effects. Column (3) has bilateral fixed effects, but no time dummies. Column (4) includes bilateral fixed effects and time dummies. Eliminating a major source of time-invariant unobserved heterogeneity yields some much more plausible findings. Focusing first on the

regional EIAs, the oldest regional agreement in the Americas, CAEIA, was a short-lived FTA among five of six Central American economies (excluding Panama) before CACM – the 'Common Market' (more accurately, a customs union) – went into full effect in 1966; El Salvador, Guatemala and Nicaragua joined the FTA in 1961, Honduras in 1962 and Costa Rica in 1963. While one would not expect a material effect from CAEIA, the coefficient estimate of 0.60 in column (3) is economically and statistically significant; however, in column (4) the coefficient estimate is notably smaller.

Much anecdotal evidence suggests that trade under CACM 'flourished'. The agreement was novel in that it provided *immediate* free trade on 95 per cent of all goods traded among the initial members. According to SICE (2006), CACM 'flourished as the most advanced and successful regional integration scheme of Latin America in the 1960s'. By contrast, the Latin American Free Trade Agreement (LAFTA) is not included in our sample because, by most accounts, it never reached effectively FTA status. Columns (3) and (4) report that CACM had an economically and statistically significant effect, supporting the anecdotal evidence that it flourished in the 1960s and, to a lesser extent, in the 1970s, becoming moribund by the late 1970s due to cross-border conflicts. The coefficient estimate of 1.01 in column (4) implies that CACM membership increased two members' average trade by 175 per cent over its likely effective period (mid-1960s to mid-1970s).

CARICOM is the second oldest EIA in the Americas, and likely the oldest *continually effective* agreement in the Americas, not suffering from the moribund status that CACM experienced in the 1970s and 1980s. While the coefficient estimate of 0.27 in column (3) is quite small, column (4) reports an economically and statistically significant coefficient estimate of 0.95, implying the CARICOM membership increased the average two members' trade by 159 per cent. We note that Table 5 also includes the short-lived CARIFTA – the initial FTA that began in 1966 before transforming into the 'common market' CARICOM in 1973. However, both columns (3) and (4) report little effect of this initial FTA. In sum, CACMFTA/CACM and CARIFTA/CARICOM were the only two effective regional EIAs in the Americas going back to the 1960s. Both apparently had economically and statistically significant (partial) effects on members' trade. While CACM became moribund during the late 1970s, its initial immediate liberalisation of trade on 95 per cent of products seemingly had a large effect. CARICOM seemingly benefited by being the oldest *continually effective* EIA in the Americas.

Outside of some bilateral EIAs during the 1980s, regional EIAs did not arise again until the 1990s with (in chronological order) the resuscitations of the Central American Common Market (CACM2, 1990) and the Andean Community (1990, with an FTA in 1994 and customs union in 1995) and formations of NAFTA (1994) and the Group of Three (G3, in 1995).

Table 5 reports that each of these agreements had economic success in terms of non-trivial impacts on trade. Focusing for brevity on column (4), common membership in MERCOSUR by 2000 had an impact on trade of 139 per cent over 10 years ($e^{0.87}$). The resuscitation of CACM in 1990 (CACM2) had a notable effect on trade among members. The coefficient estimate of 0.70 in column (4) implies that trade roughly doubled over the 10-year period. While the Andean Group was first formed in 1969, problems resulted in the customs union not being accepted until 1995. Over these years (1995–2000), the Andean Community still managed to improve trade non-trivially for its members, with the increase in trade estimated at approximately 300 per cent. This may well reflect earlier cooperation under the Andean 'Pact'. While this is the highest estimate of the EIAs' effects, one has to take into account that initial tariffs vary considerably among countries, with developing countries often having initial tariffs before liberalisation of three to four times those of developed economies. And the countries of the Andean Community were among the most protectionist of countries; in 2002, the members of the Andean Community had the highest average external tariff rates in all of the Americas, 11.65 per cent; cf. *Latin Business Chronicle* (2007).

NAFTA was signed in 1994, and so had six years phased in of a 10-year phase-in period. The coefficient estimate for NAFTA of 1.24 implies an increase of trade of 245 per cent over the six-year period, which seems somewhat high. We will find using the theoretically-motivated gravity equation that the likely estimate is much lower. Finally, the G3 went into effect in 1995. The G3 countries also maintained initially high tariffs; in 2002, Colombia, Mexico and Venezuela were also among the very high external tariff countries in the Americas. This may explain that the G3 FTA had a substantive effect on trade among members, even though the agreement was only five years old by 2000; Table 5 reports in column (4) that the G3 increased trade by 260 per cent. However, in the theoretically-motivated gravity equations below, we will find that effect diminished considerably.

Table 5 also reports a large number of average treatment effects from the numerous bilateral EIAs beginning with the El Salvador-Panama FTA of 1974 and ending with the Mexico-Israel and Mexico-EU FTAs of 2000. For brevity, we will not discuss all these treatment effects, but will note two interesting outcomes. Using Table 5's results in column (4), there were only four statistically significant treatment effects for bilateral agreements. Interestingly, two of the four were for the two oldest bilateral FTAs, the El Salvador-Panama FTA of 1974 and the Guatemala-Panama FTA of 1975; the next oldest bilateral FTA in the Americas was the Colombia-Costa Rica FTA of 1985. The estimated treatment effect for the El Salvador-Panama FTA was 1.05, implying it increased the two countries' trade by 186 per cent over the period. The estimated treatment effect for the Guatemala-Panama FTA of 1975 was 1.68, implying an impact on their trade of 436 per cent. Moreover, we will find shortly that these two results are robust to alternative specifications.

Finally, it is important to address the estimated treatment effect for *all other* EIAs over the period (*OtherEIA*). The coefficient estimate from column (4) of Table 5 is 0.62, implying that all other EIAs over the period improved members' trade by 86 per cent. This estimate is important because it is consistent with the average treatment effect for *all EIAs* estimated in Baier and Bergstrand (2007). This implies that the results here are relatively robust, for the specifications at hand.

However, we emphasise that all these estimates used an 'atheoretical' specification for the gravity equation. If we account for recent advances in theoretical foundations for the gravity equation, slightly different specifications from those above surface. The specifications above suffer *ex ante* from ignoring *time-varying* 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.

c. Fixed-effects Estimation of a Theoretically-motivated Gravity Equation with Phased-in Agreements

In this section, we consider two modifications to the previous specification. In Section 1, we include country-and-time effects to account for the theoretically-motivated multilateral price terms. In Section 2, we account for the fact that all EIAs 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 EIA may have a lagged impact on their bilateral trade.

(i) Accounting for multilateral price terms

While the results in the previous section are encouraging, the gravity equation suggested by recent formal theoretical developments – summarised in the system of equations (2)–(4) in Section 3 – suggests that one needs to account for the multilateral price variables. None of the four specifications in Table 5 accounts for these. First, accounting for the multilateral price variables in a panel context suggests estimating:

$$\ln X_{ijt} = \beta_0 + \beta_1(\ln \mathrm{RGDP}_{it}) + \beta_2(\ln \mathrm{RGDP}_{jt}) + \beta_3(\ln \mathrm{DIST}_{ij}) + \beta_4(\mathrm{ADJ}_{it}) + \beta_5(\mathrm{LANG}_{ij}) + \beta_6(\mathrm{EIA}_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \varepsilon_{ijt}. \quad (7)$$

Scaling the LHS variable by the product of real GDPs (as theory suggests) implies estimating:

$$\ln [X_{ijt}/(\mathrm{RGDP}_{it}\,\mathrm{RGDP}_{jt})] = \beta_0 + \beta_3(\ln \mathrm{DIST}_{ij}) + \beta_4(\mathrm{ADJ}_{it}) + \beta_5(\mathrm{LANG}_{ij}) + \beta_6(\mathrm{EIA}_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \varepsilon_{ijt}. \quad (8)$$

In a panel setting, the multilateral price variables would be *time varying*, and consequently the results in specifications (1)–(4) in Table 5 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 and time dummies using the panel data in its current form.[7] Moreover, the theoretical model in equation (2) suggests that the coefficient estimates for the real GDP variables should be unity.

We first estimate equation (7) using bilateral (*ij*) fixed effects to account for variation in DIST, ADJ and LANG along with 'country-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. These results are reported in column (1) of Table 6.[8]

Column (2) of Table 6 imposes explicitly unitary elasticities for real GDPs. However, in the presence of the *it* and *jt* dummies, this restriction is redundant, except for influencing the intercept estimate. Scaling or not scaling real trade flows by real GDPs will not matter for estimating the ATE in this specification. In log-linear form, the variation in the logs of real GDPs is captured by the country-time (*it*, *jt*) effects, and only the estimates of the intercept and the country-time effects' coefficients change; the EIA coefficient estimates are largely unaffected. In the remainder of the results, we use the real trade flow for the LHS variable; the EIA coefficient estimates are (virtually) identical using trade shares instead (and are available on request from the authors). Given the similarity of results, we discuss those for column (2).

The oldest agreement, CACM, maintains an economically and statistically significant (partial) effect on two members' average trade, increasing trade by 90 per cent ($e^{0.64}$). This is smaller than the corresponding result in Table 5, now accounting for the multilateral resistance terms. If we assume that CACM's trade effect took 15 years to fully take effect, this implies an average annual impact of 6.2 per cent.

The second oldest agreement, CARICOM, still had the largest trade impact of any regional agreement, 1.84. This implies an average treatment effect of 530 per cent, even larger than the effect in Table 5 of 159 per cent. However, if the increase in trade from the agreement took 15 years to complete, this estimate implies average annual growth of 13.1 per cent; if it took 20 years to complete, 9.6 per cent.

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

[8] We made two modifications in the data. Due to the inclusion of lagged terms in the next section and because both CAEIA (the predecessor of CACM) was short-lived (about 2–4 years) and CARIFTA (the predecessor of CARICOM) was also short-lived (7 years), we re-computed CACM to combine CAEIA and CACM and re-computed CARICOM to combine CARIFTA and CARICOM.

TABLE 6
Panel Gravity Equations with Theoretically-motivated Country-time Effects

Variable	*(1)*	*(2)*	*(3)*
$CACM_{ij,t}$	0.65 (2.86)[c]	0.64 (2.79)[c]	–0.34 (–1.32)
$CACM_{ij,t-5}$			1.18 (4.55)[c]
$CARICOM_{ij,t}$	1.86 (4.59)[c]	1.84 (4.56)[c]	1.09 (1.66)[a]
$CARICOM_{ij,t-5}$			0.73 (1.37)
$CACM2_{ij,t}$	0.61 (2.47)[b]	0.61 (2.45)[b]	0.48 (1.42)
$CACM2_{ij,t-5}$			–0.05 (–0.12)
$MERCOSUR_{ij,t}$	0.78 (2.76)[c]	0.77 (2.74)[c]	0.33 (0.88)
$MERCOSUR_{ij,t-5}$			0.48 (0.97)
$NAFTA_{ij,t}$	0.37 (0.85)	0.38 (0.88)	0.46 (0.78)
$NAFTA_{ij,t-5}$			0.07 (0.10)
$G3_{ij,t}$	0.67 (1.68)[a]	0.67 (1.68)[a]	0.63 (1.17)
$G3_{ij,t-5}$			–0.21 (–0.30)
$ANDEAN\_COM_{ij,t}$	1.34 (4.94)[c]	1.34[c] (4.97)	1.07 (3.59)[c]
$ANDEAN\_COM_{ij,t-5}$			0.31 (0.52)
$SAL\_PAN_{ij,t}$	1.04 (1.74)[a]	1.05 (1.76)[a]	0.56 (0.55)
$SAL\_PAN_{ij,t-5}$			0.32 (0.35)
$GUAT\_PAN_{ij,t}$	1.78 (2.91)[c]	1.76 (2.90)[c]	0.87 (0.88)
$GUAT\_PAN_{ij,t-5}$			0.78 (0.83)
$USA\_ISR_{ij,t}$	0.24 (0.43)	0.26 (0.47)	0.35 (0.38)
$USA\_ISR_{ij,t-5}$			–0.12 (–0.13)
$COS\_COL_{ij,t}$	–0.10 (–0.18)	–0.10 (–0.17)	–0.23 (–0.24)
$COS\_COL_{ij,t-5}$			0.57 (0.60)
$DOM\_PAN_{ij,t}$	1.02 (1.58)	1.01 (1.57)	0.31 (0.75)
$DOM\_PAN_{ij,t-5}$			0.77 (0.76)
$USA\_CAN_{ij,t}$	–0.17 (–0.26)	–0.14 (–0.22)	–0.01 (–0.02)
$USA\_CAN_{ij,t-5}$			–0.27 (–0.24)
$BOL\_MEX_{ij,t}$	0.44 (0.64)	0.43 (0.63)	0.33 (0.36)
$BOL\_MEX_{ij,t-5}$			0.09 (0.08)
$COS\_MEX_{ij,t}$	–0.11 (–0.16)	–0.11 (–0.15)	–0.52 (–0.57)
$COS\_MEX_{ij,t-5}$			0.72 (0.61)
$MER\_CHIL_{ij,t}$	–0.05 (–0.11)	–0.07 (–0.16)	–0.16 (–0.35)
$MER\_CHIL_{ij,t-5}$			
$CAN\_ISR_{ij,t}$	0.02 (0.02)	0.03 (0.03)	0.03 (0.03)
$CAN\_ISR_{ij,t-5}$			
$MER\_BOL_{ij,t}$	0.69 (1.50)	0.69 (1.50)	0.66 (1.43)
$MER\_BOL_{ij,t-5}$			
$CHIL\_CAN_{ij,t}$	0.25 (0.29)	0.25 (0.28)	0.15 (0.16)
$CHIL\_CAN_{ij,t-5}$			
$MEX\_NIC_{ij,t}$	1.10 (1.23)	1.10 (1.23)	1.05 (1.17)
$MEX\_NIC_{ij,t-5}$			
$DOM\_CARIC_{ij,t}$	0.67 (1.14)	0.67 (1.15)	0.61 (1.03)
$DOM\_CARIC_{ij,t-5}$			
$CHIL\_MEX_{ij,t}$	–0.20 (–0.22)	–0.22 (–0.25)	–0.29 (–0.33)
$CHIL\_MEX_{ij,t-5}$			
$MEX\_EU_{ij,t}$	–0.41 (–1.53)	–0.41 (–1.54)	–0.29 (–1.07)
$MEX\_EU_{ij,t-5}$			
$MEX\_ISR_{ij,t}$	–0.11 (–0.12)	–0.11 (–0.12)	0.03 (0.03)
$MEX\_ISR_{ij,t-5}$			
$OtherEIA_{ij,t}$	0.41 (7.14)[c]	0.41 (7.10)[c]	0.30 (3.97)[c]
$OtherEIA_{ij,t-5}$			0.27 (3.19)[c]
Constant	8.48 (252.97)[c]	–25.13 (–825.99)[c]	–25.49 (–994.17)[c]
Overall R^2	0.1020	0.0239	0.0228
No. observations	48,235	47,081	44,365

Notes:
t-Statistics are in parentheses. a, b and c denote statistical significance at the 10 per cent, 5 per cent and 1 per cent levels, respectively, using two-tailed *t*-tests.

As discussed earlier, CACM became effectively moribund in the late 1970s, perhaps even earlier. Nevertheless, the presidents of the five Central American economies of CACM agreed in 1990 to resuscitate CACM (CACM2). Column (2) of Table 6 indicates a coefficient estimate of 0.61 over the 10-year period of the 1990s, implying an average treatment effect of 6.3 per cent annually.

The impact of Mercosur on members' trade remained about the same in Table 6; a coefficient estimate of 0.77 is only slightly less than Table 5's estimate of 0.87 in column (4). The Mercosur coefficient estimate of 0.77 implies an average annual increase of eight per cent over the period 1991 to 2000 (formation in 1991).

Since NAFTA was phased in over a 10-year period starting in 1994, the coefficient estimate for NAFTA will only pick up six years of the agreement, and only partial liberalisation. Nevertheless, the coefficient estimate for NAFTA in column (2) of Table 6 of 0.38 implies an average annual impact on trade of 6.5 per cent, which is not that much smaller than the other estimates. However, the estimate is statistically insignificant.

The Andean Group, which actually formed an economic 'cooperation' agreement in 1969, did not implement an FTA until 1994. This was followed by the 'Andean Community' with a common external tariff in 1995. The coefficient estimate for the Andean Community dummy variable is 1.34, implying an average annual increase of 25 per cent. While this is considerably higher than the other agreements, part of the reason is that the resuscitation of the 'Andean Group' in 1989 preceded the actual FTA and customs union (the Andean Pact's effectiveness was stagnant in the previous decades). Moreover, *Latin Business Chronicle* (2007) documents that the Andean Group has had the highest tariffs in Latin America, so that an EIA between them would have created a very large decline in the members' intra-Andean-Community bilateral trade costs.

The final regional agreement is the Group of Three (G3) agreement among Colombia, Mexico and Venezuela, implemented in 1995. While only five years old by the end of the sample (2000), the coefficient estimate for this agreement was 0.67, implying an average annual trade impact of 14 per cent. As with the Andean Community, part of the large trade impact may have been associated with large initial tariffs. All three countries in 2002 were among the highest tariff countries in Latin America, ranging from 12 to 16 per cent, so that an EIA would have a substantive impact on members' trade.

Finally, it should be noted that – with the exception of the average treatment effect estimate for NAFTA – all six other regional agreements' treatment effects had both economic *and statistical* significance.

As before with Table 5, we will not explain in detail the estimated average treatment effects for all 17 bilateral agreements listed in Table 6. However, we

do draw attention to the two coefficient estimates for bilateral agreements that were both economically and statistically significant. The two include the 1974 El Salvador-Panama FTA and the 1975 Guatemala-Panama FTA, which as noted earlier are also the oldest bilateral FTAs in the Americas. Two other FTAs that had economically large effects, but their coefficient estimates lacked precision (statistically insignificant from zero), were the Dominican Republic-Panama FTA of 1987 and the Mexican-Nicaragua FTA of 1998. It should be noted that – with the exception of the politically-motivated Israel-USA FTA of 1985, the Colombia-Costa Rica FTA of 1985 and the Canada-USA FTA of 1989 – all the other bilateral FTAs (many of which had substantive, but not statistically significant, effects) were formed in the 1990s, many in the late 1990s.

(ii) Accounting for 'phased-in' agreements and lagged terms-of-trade effects

In this section, we introduce lagged effects of EIAs on trade, to provide some measure of robustness. The economic motivation for including lagged changes stems partly from the institutional nature of virtually all EIAs. The 0–1 variables were constructed using the 'Date of Entry into Force' of the agreement, as best surmised by scrutinising multiple data sources provided earlier. However, virtually every EIA is 'phased-in', often over 5–10 years; CACM was an exception noted earlier. 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 EIA 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 EIA dummy. Because most of our agreements have been in place for a short period – and CACM faced a long period of dormancy – we include only one lagged level; in most cases, the second lagged level had to be 'dropped' in the regression because no data were available.[9]

Moreover, economic effects of an EIA 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 EIA 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 column (3) in Table 6 provide some evidence that the results discussed in the previous section are largely robust to having a contemporaneous impact as well as a lagged impact. Consider first the seven regional agreements (counting CACM and CACM2 as separate agreements due to the 15–20-year gap

[9] For regional EIAs, only one EIA could be estimated with two lags, and for bilateral EIAs, only six of 17 EIAs could include two lags.

in its enforcement). Summing the average treatment effects for the current and lagged values in column (3) and comparing the sum to the current effect only in column (2) provides some economically plausible findings, with the exception of CACM. For instance, for CARICOM the separation of current and lagged effects shows that not all of the trade impact of the agreement occurs in the concurrent period, with the lagged term allowing both for 'phasing-in' of the agreement as well as lagged terms-of-trade effects. In many cases, by separating the effects into current and lagged effects, economically plausible coefficient estimates become statistically significant. However, in no case did any of the coefficients that were unlikely (negative coefficients) have statistical significance. In general, the results in column (3) of Table 6 are largely consistent with those in columns (1) and (2) of Table 6.

This same pattern of corroboration holds when the bilateral EIA coefficient estimates in column (3) are compared with the corresponding estimates in columns (1) and (2) of Table 6. However, because of the short durations of many of the recent bilateral FTAs, coefficient estimates for these agreements in columns (1), (2) or (3) – while economically plausible – are not statistically significant. Our conclusion is that an *extension of the sample* to 2005 is necessary to uncover more economically and statistically significant trade effects of bilateral EIAs in the Americas.

5. PARTIAL VS. FULL (GENERAL-EQUILIBRIUM) EFFECTS

As noted earlier, Anderson and van Wincoop (2003) remind researchers that the coefficient estimate for an EIA dummy variable represents only the 'partial' (equilibrium) effect of the formation of that EIA – *not* the (full) 'general-equilibrium' (comparative static) effect. Since trade flows and prices are likely determined simultaneously in the context of a (full) general-equilibrium system, such as equations (2)–(4) earlier, an estimate of β_6 only captures the partial effect, as discussed earlier. In reality, the formation of an EIA between countries i and j alters the relative price an importer pays for the trade flow from i to j, but it also influences the prices of *all* trade flows in the world economy. In equations (2)–(4), this implies that an EIA_{ijt} will alter P_{it} and P_{jt}. However, the consequent (likely) reductions in P_{it} and P_{jt} will cause *less* trade from i to j (for instance, the fall in P_{jt} causes consumers in j to increase demand for all *non-ij* trade flows to j). This 'general-equilibrium' (feedback) effect tends to offset the partial effect on X_{ijt} of a change in EIA_{ijt}. Thus, the full effect on X_{ijt} of EIA_{ijt} needs to account for the (likely dampening) general-equilibrium effects. Hence, Anderson and van Wincoop emphasise estimation of the full system of equations to generate the multilateral price terms to conduct appropriate comparative statics.

That said, how important *quantitatively* are the general-equilibrium effects in the context of this study? There is reason to believe that the general-equilibrium effects in the context here are quantitatively quite small. In related research, we have employed the Anderson-van Wincoop technique in the context of estimating the effects of EIAs using *simulated* trade flows from the system of equations (2)–(4), based upon actual real GDPs, distance, adjacency, language and 'NoFTA' dummy variables among 96 countries (including the Americas). By using simulated trade flows generated by the (true) model, we avoid the endogeneity of FTAs. In that research, we define a dummy variable 'NoFTA', which assumes a value of 1 where no FTA exists and 0 where an FTA exists (the *opposite* of 'EIA' defined earlier), with no loss of generality. We assign (exogenously) a coefficient for the NoFTA dummy of –0.7, which implies a 'partial' effect (from NoFTA) of reducing trade by 50 per cent, which is consistent with the empirical estimate of an average treatment effect for FTA (or EIA) of 0.7; cf. Baier and Bergstrand (2007).

In this context, general-equilibrium effects should dampen the direct effect, which is what we find in our world (based upon actual GDPs, distances, etc., but *simulated* trade flows). However, in the case of EIAs for the Americas, we find that the dampening effects are less than 10 per cent in more than half of the Americas' cases (precisely, 15 of 26), are between 10 and 20 per cent in 11 more cases, and exceed 20 per cent *in only two of 26 country pairs*. Thus, while important, we find that the dampening effect is less than 20 per cent in the vast bulk of country pairs.

The rationale for the limited dampening effects on the partial effects is that – in a world with nearly 100 countries, vast bilateral distances, and distance, adjacency, language and FTA elasticities mirroring typical estimates – the multilateral price terms do not change very much quantitatively, limiting the role of general-equilibrium (feedback) effects for offsetting the 'partial' effects. Thus, while the partial effects discussed in preceding sections provide only limited information on the full comparative static effects of EIAs on trade, the limitations are likely not quantitatively large in many cases. Estimation of the full *ex post* general-equilibrium empirical effects is the subject of future research.

6. IMPLICATIONS AND CONCLUSIONS

What do these empirical results mean for better understanding the 'growth' of regional economic integration agreements? National policy makers around the world, operating in an increasingly competitive global environment, face strong pressure from their national constituents (firms, households) to maximise their economic status (profits and consumer welfare, respectively). Such policy makers

must make decisions about trade policies in such an environment. The proliferation of bilateral and regional EIAs in the world economy likely mirrors the proliferation of bilateral and regional trade in the world economy. There is a world market for goods and services that is met efficiently by bilateral trade flows. Correspondingly, there has likely emerged a world 'market' for bilateral and regional trade policies/institutions to *facilitate* the bilateral exchange of products, owing largely to the gains from specialisation and the welfare benefits of product diversity for final goods producers (i.e. product differentiation in intermediates) and consumers (i.e. product differentiation in final goods). The vast bulk of EIAs are among countries: (1) that are close in distance and consequently share low bilateral transaction costs; (2) that are large in economic size and consequently benefit from greater specialisation in production and variety in terms of consumption; (3) that differ in relative factor endowments, benefiting from the exchange of traditional comparative advantages. That is, countries that have selected into EIAs have 'chosen well'; cf. Baier and Bergstrand (2004b).

Traditional *ex ante* estimates of the trade and economic welfare gains from EIAs have often suggested relatively modest economic benefits. Much anecdotal evidence from policy makers suggests that the anticipated economic gains are much larger than traditional CGE models have implied. However, sufficient time has now passed – and econometric and theoretical developments advanced – such that policy makers can now examine with more precision the *ex post* effects of EIAs on trade patterns. The evidence in this paper suggests that the trade effects of membership in the EIAs in the Americas have been *much larger* than those suggested by *ex ante* considerations *and* much larger than even earlier empirical estimates using cross-sectional gravity equations suggested; cf. Frankel (1997). The results here suggest that EIA memberships over the past 40 years (1960–2000) in the Americas are of an economically-significant magnitude, considerably larger than that postulated a decade ago in Bayoumi and Eichengreen's excellent analysis of early EEC6 effects between 1957 and 1972.

Policy makers around the world have likely drawn lessons from the apparent success of the major economic integration agreement experiment of 1957, the Treaty of Rome. They have likely pursued the seeming trade enhancements for bilateral and regional EIAs. And the evidence in this paper suggests that their 'economic expectations' have largely been correct. After accounting for the pitfalls associated with the 'endogeneity of EIAs' determination', the vast bulk of EIAs have tended to augment members' trade by about *100 per cent* over a 15-year period. This is consistent with anecdotal evidence from policy makers that the economic benefits from EIAs are much larger than conventional *ex ante* economic analyses have previously suggested.

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

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