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Do Economic Integration Agreements Actually Work? Issues in Understanding the Causes and Consequences of the Growth of Regionalism

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1. INTRODUCTION

ONE of the most notable phenomena in the world economy over the past 20 years has been the enormous growth in the number of international economic integration agreements (EIAs). EIAs are treaties between economic units – in the case of international EIAs, between nations – to reduce policy-controlled barriers to the flow of goods, services, capital, labour, etc. Most – though not all – EIAs tend to be 'regional' (or continental) in scope and most tend to be free (or preferential) *trade* agreements (henceforth, FTAs). According to the World Trade Organization (WTO) website, in 2006 there were approximately 300 regional trade agreements that were either planned, had concluded negotiations, or were in force. Interestingly, of the 250 agreements notified to the General Agreement on Tariffs and Trade (GATT) and WTO between 1947 and 2002, about half have been notified since 1995. Thus, there has been a virtual explosion

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in the number of EIAs in the past decade. This is the 'latest wave' of regional trade and cooperation agreements that comes on the heels of the 50th anniversary of the most noted economic integration agreement of modern times, the 1957 Treaty of Rome.

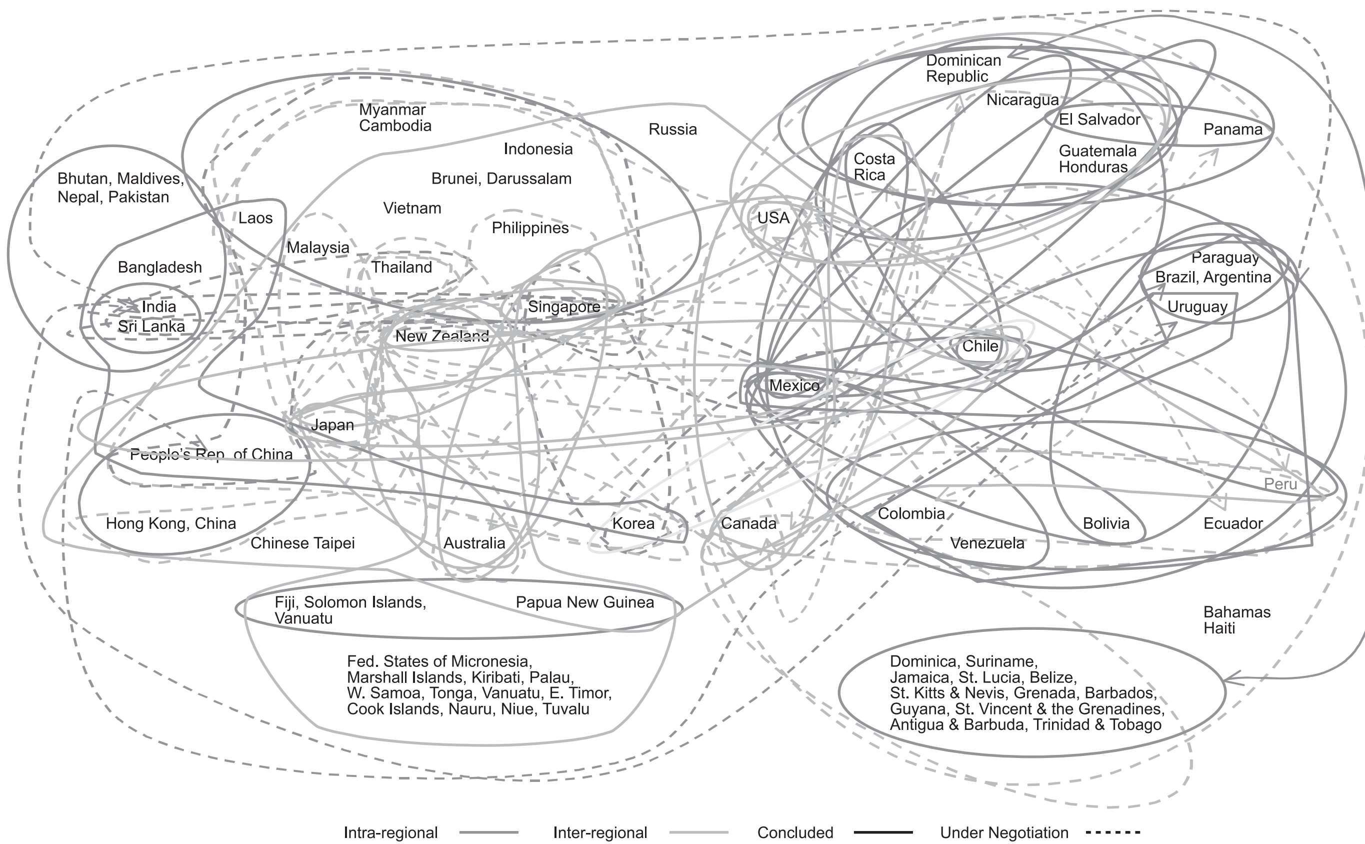
This wave has culminated in – what Jagdish Bhagwati and Arvind Panagariya (1999) have famously termed – a seeming 'spaghetti bowl' of EIAs. Figure 1, from Estevadeordal (2006), illustrates vividly this 'spaghetti bowl', with each line representing an EIA between one country and another (or with a group of countries). However, one aim of this paper is to convince the reader that, instead of looking at this web of agreements as a spaghetti bowl, economists and policy makers should see this as a 'market for regionalism'.

This paper synthesises and develops further a line of research pursued by the authors on the causes and consequences of the growth of regionalism. In this paper, we hope to accomplish four goals. First, we address conceptually why it is useful to consider this web of agreements as a type of 'market'. In a world with approximately 200 countries and national governments, there exist approximately 20,000 potential bilateral EIAs ($200 \times 199/2 = 19{,}900$). To the extent that national governments promote the welfare of their nations' firms and consumers, the rules of engagement in bilateral trade are likely determined in a highly competitive political environment. We discuss the notion of 'competitive liberalisation', coined by Fred Bergsten more than a decade ago, and suggest a systematic conceptual framework for analysing determinants of EIAs, initially in a static context. While bilateral trade agreements are ultimately negotiated by national governments, the rules are negotiated in the context of a type of 'market' of 20,000 potential bilateral agreements, which can provide potentially for the beneficiaries of such agreements – various nations' firms and consumers – to influence their national policy makers to negotiate in a competitive manner. To a large extent, one might interpret our approach in the context of the 'new institutionalism'. We discuss empirical evidence consistent with the notion that EIAs are determined in a competitive economic environment.

Our second goal is to argue that the market for (bilateral) EIAs exists contemporaneously with the market for (bilateral) trade flows, obscuring *ex post* evaluation of the effects of EIAs on trade. For instance, country pairs that are physically close and are large economically tend to have very large trade flows, e.g. US–Canada and France–Germany. Moreover, countries that choose to form EIAs are physically close and are large economically, e.g. US–Canada and France–Germany. However, if trade flows and EIAs are determined simultaneously, this raises problems for evaluating *ex post* the effects of EIAs on trade.

Our third goal is then to address issues concerned with providing better estimates of the *ex post* effects of EIAs on trade in the context of this world. While computable general equilibrium (CGE) models have long dominated policy makers' analyses of the potential economic benefits from changing trade policies

FIGURE 1
The 'Spaghetti Bowl' of FTAs in the Americas and Asia-Pacific (2005)



Source: Integration and Regional Programs Department, IDB.

(including formation of EIAs), such models can only provide *ex ante* forecasts of the effects of eliminating (or reducing) measurable government-imposed trade barriers on trade, production, consumption and welfare of a nation. These models cannot address what actually *did* happen as a result of forming a specific EIA. Moreover, many have argued that CGE models have tended to underestimate considerably the effects of EIAs on trade (cf. DeRosa and Gilbert, 2005). Policy makers should be interested – and, we conjecture, *are* interested – in *ex post* quantitative estimates of the effects of an EIA on trade flows (and, subsequently, on production, incomes, etc.). As John Whalley puts it in his article in this same collection of symposium papers: 'A recent World Bank (*Global Economic Prospects*, 2005) estimate is that perhaps around 43 per cent of world trade was covered by agreements in force in 2003 and was projected to increase to 55 per cent by 2005 (OECD, 2003). But such calculations only raise more questions: *How large are the impacts of these agreements on covered trade*?' (Whalley, 2008, this issue; emphasis added).

Surprisingly, estimates to date using the workhorse for *ex post* empirical analysis of the effect of EIAs on trade flows (the 'gravity equation') often find economically and statistically *insignificant* effects of EIAs on trade (cf. Frankel, 1997). Moreover, recent empirical evidence shows that such estimates are quite *fragile* (cf. Ghosh and Yamarik, 2004). We address estimation techniques that suggest that previous estimates are likely biased downward. Moreover, we provide empirical evidence of much more 'sturdy' (*ex post*) estimates of the trade effects of EIAs. In fact, one of the advantages of using the gravity equation approach is that, when properly specified, it may actually be able to capture *ex post* the trade effects for EIA members of liberalisation of the 'complex and elaborate' barriers (beyond simple tariff cuts) that previous approaches (such as CGE models) cannot offer.

Our fourth goal is, then, to address how the previous three issues help us to better understand the 'latest wave' of regional trade agreements. We argue that policy makers have tended to expect larger trade effects from EIAs than *ex ante* CGE models have suggested. Because policy makers have self-selected into EIAs due to larger expected effects, previous *ex post* estimates of the trade effects of EIAs (ignoring self-selection) have been biased downward. Using 'sturdier' estimates, we then confirm this conjecture for Europe, demonstrating much stronger EIA effects on trade than witnessed previously.

2. DETERMINANTS OF BILATERAL TRADE FLOWS AND BILATERAL ECONOMIC INTEGRATION AGREEMENTS

International economists such as Richard Baldwin (1995) and C. Fred Bergsten (1996) noted more than a decade ago that there were seemingly strong competitive pressures in the world economy – sensed by nations' governments – that induced such governments to liberalise trade both bilaterally and regionally. The large numbers of nations party to the GATT/WTO has grown over the past 50 years to approximately 150 countries. This large number of parties has likely made the ability of negotiators to liberalise trade in agriculture, goods, services, capital and labour under one agreement much more difficult.[1] Nevertheless, governments are pressured by individual voters and firms' lobbies to provide a framework of policies (or 'institutions') well-suited to both constituencies' interests (maximising economic welfare and economic profits, respectively). In the face of these pressures and an impasse in multilateral trade and investment liberalisation at the WTO level, governments have sought alternative policy changes to improve economic welfare and firms' profits. One alternative – potentially a 'building block' for further multilateral liberalisation – is economic integration agreements (which include bilateral agreements). As shown in Figure 1, the proliferation of EIAs over the past 50 years has created the so-called 'spaghetti bowl' of EIAs.

However, Baldwin's 'domino theory' of regionalism and Bergsten's 'competitive liberalisation hypothesis' are implicitly dynamic stories. In our view, before one can conceptualise about the 'latest wave' of regionalism (which is also implicitly dynamic), we consider it imperative to address first 'Regionalism'. That is, we start with a *static* long-run view of the determinants of regionalism (and bilateralism). The notion of 'competitive liberalisation' can be consistent with a static concept of regionalism as well as a dynamic one. As is traditional in economics, one should probably examine the *long-run* economic factors influencing the equilibrium outcome *before* modelling explicitly the short- and medium-run factors influencing EIA formation, where the latter are often more easily observed and often discussed less technically.

We have intentionally used the term 'economic integration agreements' initially to be inclusive. The term 'economic integration' spans integration of goods, services, capital and labour markets; in even broader views, it encompasses integration in economic activity that goes beyond economists' traditional categorisations of 'goods' and 'factors'. We also used 'economic integration' – not 'regional economic integration' – to be inclusive in geographic scope of coverage. Many recent economic integration agreements – the recently-signed Australian–US FTA, for example – involve countries on different continents; economists have occasionally referred to these as 'unnatural' EIAs, in the sense that they are not in the same geographic region or on the same continent.[2]

[1] See Mansfield and Reinhardt (2003) and Moravcsik (2008). Also, Moravcsik (2008) argues that competitive liberalisation pressures have been the dominant force behind much of European economic integration, with the likely exception of Germany's motivation in the 1950s.

[2] See, for example, Krugman (1991a,b), Frankel et al. (1995, 1996), and Frankel (1997).

However, the vast bulk of EIAs are regional free trade agreements, limited in scope to countries sharing common continents and to goods (and, in many cases, services) sectors. In the remainder of this paper, we will continue to use the acronym EIA to be inclusive of FTAs, customs unions, common markets and economic unions, although most of the focus will be on the trade implications of EIAs. One reason for this is that, in the empirical analysis later, our EIAs will include some deeper integration agreements, such as the European Union.

a. Determinants of Bilateral Trade

Before addressing directly static determinants of EIAs, 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 and economic welfare 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.[3]

Because regionalism typically entails bilateralism,[4] 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 by '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 multi-country (or N country) theoretical foundations for a 'gravity equation' of bilateral international trade over the past 25 years, and in a manner consistent with established theories of intra- and inter-industry international trade. For instance, the first formal economic theoretical foundation for the gravity equation with a one-sector endowment economy, but many countries, was

[3] Our analysis initially will take as given exogenously the prevailing level of policy-oriented trade barriers, such as tariff rates. In reality, the ideal approach would be to consider the endogenously-determined Nash equilibrium tariff rates pre- and post-integration, as the pre-integration Nash equilibrium tariffs are likely to differ from the post-integration ones. Addressing this limitation, however, is beyond the scope of this paper.

[4] In the remainder of the paper, we often use the terms 'bilateralism' and 'regionalism' interchangeably.

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:[5]

$$PX_{ij} = \beta_0 (GDP_i)^{\beta_1} (GDP_j)^{\beta_2} (DIST_{ij})^{\beta_3} \varepsilon_{ij}, \tag{1}$$

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 with trade flows 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, and showed empirically their importance for explaining 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, 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- and 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.

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 a system of nonlinear equations (2)–(4) below using custom nonlinear least squares (NLLS) programming could account properly for endogeneity of prices:

$$PX_{ij} = \beta_0 (GDP_i)^1 (GDP_j)^1 (t_{ij})^{1-\sigma} P_i^{\sigma-1} P_j^{\sigma-1} \varepsilon_{ij}, \tag{2}$$

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

[5] 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.

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 GDP_i/GDP^{T} (GDP_j/GDP^{T}). Details of estimating (2) for aggregate trade flows using either nonlinear least squares or fixed effects for P_i and P_j are addressed in Anderson and van Wincoop (2003), Feenstra (2004), and Baier and Bergstrand (2002, 2006, 2007).[6] 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–Samuelson inter-industry trade and Chamberlin–Helpman–Krugman intra-industry trade (cf. Carrere, 2006).

While acknowledging the endogeneity of prices and efficiency of estimating equations (2)–(4) using NLLS, Baier and Bergstrand (2006) suggest a method for estimating coefficient estimates in equations (2)–(4) using ordinary least squares (OLS) that are virtually identical to those estimated using Anderson and van Wincoop's NLLS program or fixed effects. Using a first-order log-linear Taylor series approximation of the theory in Anderson and van Wincoop (2003), Baier and Bergstrand show that OLS estimation of:

$$\frac{PX_{ij}}{GDP_i GDP_j / GDP^{\mathrm{T}}} = \left(\frac{t_{ij}}{t_i(\theta)t_j(\theta)/t^{\mathrm{T}}(\theta)}\right)^{-(\sigma-1)}, \tag{5}$$

where $t_i(\theta) = \Pi_{j=1}^{N} t_{ij}^{\theta_j}$, $t_j(\theta) = \Pi_{i=1}^{N} t_{ji}^{\theta_i}$, $t^{\mathrm{T}}(\theta) = \Pi_{i=1}^{N}\Pi_{j=1}^{N} t_{ij}^{\theta_i\theta_j}$, and recall $\theta_i = GDP_i/GDP^{\mathrm{T}}$ and $t_{ij} = t_{ji}$, provides identical coefficient estimates to the other two methods

[6] 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, 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.

and can generate approximations of the multilateral price terms that are highly correlated with those generated by OLS.

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, 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. As mentioned earlier, 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. Estimates of gravity equation (2) for EIAs are scarce, since equation (2) surfaced in the past five years. Baier and Bergstrand (2002, 2007) and Carrere (2006) provide some early estimates.

b. Determinants of EIAs

A key notion in this paper is that bilateral EIAs are – like bilateral trade flows – endogenous and under certain assumptions may be considered to be determined in a competitive setting as well. In considering what factors might explain whether or not certain country pairs are likely or unlikely to have an EIA, one needs to distinguish along two dimensions. First, we address static versus dynamic determinants of EIAs. In the static view taken in this section, we consider a world in 'long-run equilibrium'. We ask the question: what are some economic factors that explain theoretically whether or not a pair of countries is likely to have an EIA (in equilibrium)? We then examine empirically using a cross-section qualitative choice econometric model whether or not the pairs of countries that have EIAs are the most likely ones to have such agreements, conditioned upon a set of economic determinants suggested by theory (relative economic sizes, relative factor endowments, trade costs, etc.) and that full *multilateral* free trade liberalisation under the WTO is prohibitively expensive.[7]

[7] In our theory, we assume that the decision to have or not to have an FTA takes as exogenous the current WTO structure that impedes achieving 'free' trade. We assume, as Bergsten (1996) states, 'It simply turns out to be less time-consuming and less complicated to work out mutually agreeable arrangements with a few neighbors than with the full membership of well over 100 countries in the WTO' (p. 4). This is also consistent with the approach taken in Grossman and Helpman (1995b) that, 'As in Grossman and Helpman (1994, 1995a), we suppose the incumbent government is in a position to set trade policy, which means here that it can either work toward a free-trade agreement or terminate the discussions' (p. 670). A multilateral trade policy alternative is ruled out by assumption. Also, since Bergsten wrote, there are now 150 parties to the WTO. Zissimos (2007) demonstrates in a game-theoretic setting the relevance of geography (i.e. trade costs) for the formation of FTAs.

Second, we must distinguish between the 'economics' of EIAs versus the 'political economy' of EIAs.[8] In reality, of course, national governments are empowered to sign treaties regarding international commerce and factor mobility. In the international trade literature, it is common to assume that a representative (national) government's objective is to maximise a weighted average of the welfare of individuals (in economic terms, voters' utilities) and the influence of firms (in economic terms, firms' economic 'rents' or profits), which likely operate through lobbies.[9] While both factors play a role in reality, we follow the intuitive suggestion by Bergsten (1996) that – in a long-run view – economic welfare is likely to be the dominant force, and that political factors (lobbies, special interest groups, etc.) are likely to be relatively more important in the short to medium run. Bergsten (1996, p. 2; emphasis added) states:

> There are of course different national circumstances which explain the detailed strategies and timing of the individual initiatives. The overarching force, however, has been the process of competitive liberalization. The rapid increase of global interdependence has forced all countries, whatever their prior policies or philosophies, to liberalize their trade (and usually investment) regimes. *Economic success in today's world requires countries to compete aggressively* for the footloose international investment that goes far to determine the distribution of global production and thus jobs, profits and technology.

In our initial static analysis of selection into EIAs, we assume that the economic welfare of two nations' representative consumers determines whether or not the governments of that pair choose to have an EIA or not. To avoid the role of economic rents (or excessive profits), we assume monopolistically competitive markets for the production of goods, with large numbers of profit-maximising firms that find political coordination prohibitively costly; this simplifies the model.[10] In a dynamic analysis that addresses more the 'timing' of formations of EIAs, political-economy considerations and economic rents could surface.[11]

Following in the spirit of Krugman (1991a,b), Frankel et al. (1996), and Frankel (1997), Baier and Bergstrand (2004) created a model of a world economy with asymmetric countries recognising explicitly inter- and intra-continental trade

[8] We borrow this useful distinction from Krugman (1991a).

[9] See, for example, Grossman and Helpman (1995b) or Gawande et al. (2005).

[10] Even in a monopolistically competitive framework, countries might optimally choose higher tariffs in equilibrium. We assume they do not for three reasons: (1) the spirit of the GATT/WTO, where EIA members are precluded from raising their average external tariffs; (2) the Nash equilibrium may even yield a lowering of external tariffs (see work by Yi, 2000, and Ornelas, 2005); and (3) we have not observed increases in external tariffs (see empirical work by Estevadeordal et al., 2005).

[11] One may further distinguish the economic and political-economy determinants of EIAs from the 'politics' (or political science) of determination of EIAs, where the latter literature deals theoretically and empirically with the role of democratic institutions, etc.; see, for instance, Mansfield and Reinhardt (2003) on such issues.

costs. Krugman (1991a) used a simple model of three symmetric (or identical) economies where firms produced slightly differentiated goods under increasing returns to scale in production to show that – in a world with *no* trade costs – regional EIAs decreased economic welfare of households unambiguously. However, Krugman (1991b) showed that in the same model – but with prohibitive inter-continental trade costs – regional EIAs increased economic welfare unambiguously. Frankel et al. (1996) cleverly labelled this the 'Krugman vs. Krugman' debate. Frankel et al.'s extension of Krugman's model usefully allowed for a continuum of inter-continental trade costs, distinguishing 'natural' EIAs (within continents) from 'unnatural' EIAs (across continents). Frankel et al. could then show the cross-over point – in terms of inter-continental trade costs – at which net welfare changed from positive to negative. Using some empirical estimates of the costs of inter-continental trade based upon a gravity model of trade, one conclusion from Frankel's (1997) book was that, if all continents followed the European example, the regionalisation of the world economy would be '*excessive*'.

In order to establish a quantitative model to predict which pairs of countries should or should not have an EIA, Baier and Bergstrand (2004) extended the Frankel–Stein–Wei model to allow for asymmetric economies – both in terms of economic size and in relative factor endowments – and for asymmetric inter- and intra-continental transport costs. The model has six countries on three continents with countries on the same continent facing (Samuelson) iceberg-type intra-continental trade costs and countries on different continents facing additional iceberg-type inter-continental trade costs. Each country is endowed with two factors of production, capital (K) and labour (L). There are two industries, goods and services, with preferences for the two sectors' outputs of the Cobb–Douglas type. Preferences for each sector's output are of the constant elasticity-of-substitution (CES) type, common to the trade literature. Each sector's products are slightly differentiated, with each product produced under increasing returns to scale; consumers value variety. The production of goods and of services uses capital and labour in different relative factor intensities. Standard demand functions are generated, the details of which are discussed in Baier and Bergstrand (2004).

If governments are welfare maximisers, then – in the context of this theoretical model – certain *economic characteristics* are likely to favour EIA formation in some pairs of countries relative to others.[12] For example, two important economic factors influencing trade and utility are intra-continental and inter-continental trade costs. First, countries that are closer together (on the same continent)

[12] Moreover, in the context of 20,000 potential bilateral interactions, each government is assumed to operate competitively, taking as given the behaviour of other governments (and welfare of their consumers).

benefit more from an EIA because, with lower intra-continental trade costs, they are already large traders. Second, the net benefits of a natural EIA increase (and the net costs of an unnatural EIA decrease) as inter-continental trade costs rise, because more remote countries trade little with distant countries.

Baier and Bergstrand (2004) demonstrate also that pairs of larger GDP economies tend to benefit more from EIAs than pairs of smaller countries, due to economies of scale in production and increased varieties of products available. As two countries' GDPs become more different, the likelihood of an EIA decreases. A larger economy's benefit from an EIA diminishes as the two countries become more dissimilar in size (for a given total economic size) because the breadth of variety in imports from a small EIA partner contracts for the larger economy.

Due to the presence of two industries and two factors, the wider the relative factor endowments of a country pair, the more likely an EIA (if inter-continental transports are sufficiently high) due to the gains of exchange relative comparative advantages, i.e. inter-industry trade. However, the wider the difference in two partners' relative factor endowments relative to the rest of the world, the less likely an EIA. It is important to note – as perhaps surmised already – that most (if not all) of these economic factors are also well established as economic determinants of bilateral trade flows.

Based upon the qualitative choice econometric model of McFadden, Baier and Bergstrand (2004) used a probit model to try to establish empirically the relative importance of these factors for explaining – and potentially predicting – the likelihood of an EIA between country pairs. We employed a sample of bilateral pairings among 54 countries, or 1,431 observations for EIAs observed in 1996 [$(54 \times 53)/2 = 1{,}431$]. These probabilities are predicted using bilateral distances, GDP sizes, GDP similarities, relative K/L ratios, and indices of remoteness (or multilateral resistance) as explanatory variables (Baier and Bergstrand, 2004).

We draw attention to three empirical outcomes. First, the empirical probit model actually works quite well. Every economic relationship described above between GDPs, relative factor endowments and distance is found empirically in the probit results.[13] As a measure of overall fit, the pseudo-R^2 value of the full specification is 73 per cent for 1,431 country pairs. We note that for a (more recently constructed) wider sample of 96 countries in 1995, the pseudo-R^2 remains high at 67 per cent. Of the 286 EIAs in 1996 in our original sample, the model predicted 85 per cent (or 243) correctly. Of the remaining 1,145 pairs with no EIAs, the model predicted correctly 97 per cent ($1{,}114 = 1{,}145 - 31$). Details

[13] Egger and Larch (2006) find the same qualitative relationships between these economic variables and the probability of an FTA using a much wider sample of 178 economies and 15,753 country pairs for the year 2005.

are available in Baier and Bergstrand (2004). We note that the most likely EIAs in 1996 (using exogenous geographic variables and GDPs and *K*/*L* ratios from 1960) were the earliest EIAs.

Second, of the top 200 pairs (of 1,431) that were the most likely to have an EIA in 1996, only six pairs did not have one: Iran–Iraq, Iran–Turkey, Chile–Peru (EIA being negotiated), Japan–South Korea (EIA being negotiated), Hong Kong–South Korea and Panama–Venezuela.

Third, of the 1,000 pairs (of 1,431) that were the least likely to have an EIA in 1996, only four pairs actually had an EIA: Portugal–Turkey, Egypt–Iraq, Mexico–Chile and Mexico–Bolivia.

3. SIMULTANEOUS MARKETS FOR TRADE FLOWS AND EIAs

Why does the model work so well? We believe the model is consistent with the notion of 'competitive liberalisation'. National governments realise countries are unique in economic characteristics. In the interest of liberalising markets to improve productivity and standard of living levels, national governments select into arrangements with other countries for which they share certain economic characteristics, such as similar economic size or low trade costs (close in distance). Empirically, most pairs of countries with EIAs tend to have the key economic characteristics that the theoretical model suggests should be present for an EIA to enhance (on net) the welfare of pairs' representative consumers. In many (if not most) cases, these are pairings where countries already trade extensively with one another. This is consistent with Bergsten's 'competitive liberalisation' notion that economic welfare may be the dominant long-run 'overarching' force driving regionalism, despite political factors influencing timing, etc. Hence, the same observable variables that explain trade patterns – gravity-equation variables – also explain the likelihood of an EIA because of likely net benefits for producers and consumers from creating such an EIA. Hence, one can argue that *ex post* country pairs that have chosen to have EIAs have 'chosen well'.

The reader might ask a seemingly obvious question: if national governments are simply maximising consumers' welfare, why not simply predict bilateral EIAs with bilateral trade flows? First, there is an 'endogeneity' issue. Predicting the likelihood of an EIA based upon a probit regression using trade flows on the right-hand side (RHS) will likely yield biased coefficient estimates. The reason is that 'unobservable' variables – such as institutional and political factors – that likely influence the decision by governments to form EIAs also tend to influence trade flows. In cross-sectional data, these unobservable – to the econometrician – variables likely influence both EIA and trade variables. The coefficient estimates in the probit regression would be biased. Second, the probit specification we use helps identify the (exogenous) 'economic characteristics' that influence the decision to form an EIA: economic geography variables, factors influencing intra-industry trade, and factors influencing inter-industry trade.

The approach and results just discussed have some potentially important implications for the 45 years of empirical research using the gravity equation with cross-sectional data discussed in Section 2*a*. Since Nobel Laureate Jan Tinbergen (1962) first employed the gravity equation, the equation has been used increasingly to estimate the impact of EIAs on members' trade flows. Tinbergen (1962) studied bilateral international trade flows among several countries in a cross-section from the 1950s including dummy variables for the Benelux FTA and the British Commonwealth members; he found that membership in either of these agreements increased trade by only 5 per cent. However, the previous discussion suggests that cross-section estimates of EIAs' effects on trade over these 45 years suffer from potential selection bias. If country pairs select into EIAs for unobservable reasons correlated with potential trade flows, OLS estimates will likely be biased.[14]

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 Appendix. These estimates are derived including separate EIA dummy variables for the European Union (EU), the European Free Trade Association (EFTA), the European Economic Area (EEA), and all 'other' EIAs (OEIAs). EU_{ijt} is defined to equal 1 if a country pair ij in year t were members of the European Economic Community (1960 to 1970), the European Community (1975 to 1990), or the European Union (1995 and 2000), and 0 otherwise. $EFTA_{ijt}$ is defined to equal 1 if a country pair ij in year t were members of EFTA, and 0 otherwise. EEA_{ijt} is defined to equal 1 if one country was in *EU* and the other was in *EFTA* in year t; members of the EC (EU) formed (maintained) FTAs with remaining EFTA members in 1973 (1994). $OEIA_{ijt}$ 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

[14] A case where this is least likely to occur is the original EEC6 countries, formed based upon strong political and national security considerations. Consequently, plausible estimates of the trade effects of the EEC6 in Aitken (1973), as well as estimates of FTAs prior to the early 1970s, may well be less biased.

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 compiled also 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; hence, (one- or two-way) preferential trade agreements that were not intended to liberalise the bulk (typically, 80 per cent or more) of their trade were excluded from the sample. Table 1 lists the trade agreements used and sources.[15]

As Table 2 shows, common membership in *EU* had an economically significant effect in 1960 and 1970 only, with the sole statistically significant positive effect in 1960 – only three years into the original EEC agreement. These results are surprising. Second, common membership in *EFTA* had an economically and statistically significant effect on trade in 1960 (the year the agreement came into effect!) and in 1970 only. In fact, common membership in EFTA had more than twice the effect on members' trade than common membership in *EU*. These results are surprising. Third, common membership in any other EIA (*OEIA*) had a positive and economically significant effect in all five years examined, although the coefficient estimate is statistically different from zero in only three of the sample years (1960, 1980, 2000). Moreover, in 1970 the effect of other FTAs was to increase trade by 1,900 per cent. Consequently, the results for *OEIA* are quite fragile. All in all, the empirical results using a typical gravity equation specification – assuming the EIA variables are exogenous – are not very supportive that EIAs actually work.

As discussed earlier, the typical gravity equation (1) is likely misspecified owing to ignoring theoretical foundations that have developed over the past several decades. Table 3 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) and restricting the coefficient estimates for GDPs to be unity (as suggested by theory). As Table 3 reports, accounting for the theoretically-motivated multilateral price terms does not improve the results for EIA effects relative to Table 2. If anything, estimates from the theoretically-motivated gravity equation (2) using country fixed effects lend even less support to the notion that *ex post* EIAs actually work.[16]

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

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

TABLE 1
Economic Integration 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 (1986–95), Norway, Portugal (until 1986), Sweden (until 1945), Switzerland, United Kingdom (until 1973)
Latin American Free Trade Agreement/Latin American Integration Agreement, or LAFTA/LAIA (1961–79, 1993–): Argentina, Bolivia, Brazil, Chile, Ecuador, Mexico, Paraguay, Peru, Uruguay, Venezuela (became inoperative during 1980–90, but reinitiated in 1993)
African Common Market (1963): Algeria, Egypt, Ghana, Morocco
Central American Common Market (1961–75, 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 (1973/94)
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 Comun 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, Peru, Venezuela (1997)

TABLE 1 *Continued*

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
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 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, Lawrence (1996), Frankel (1997).

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

TABLE 2
Typical Cross-Section Gravity Equation Coefficient Estimates

Variable	*(1) 1960*	*(2) 1970*	*(3) 1980*	*(4) 1990*	*(5) 2000*
ln GDP_i	0.76 (46.57)	0.89 (57.77)	1.01 (69.37)	1.09 (85.00)	1.19 (103.97)
ln GDP_j	0.76 (49.65)	0.92 (64.17)	1.01 (73.56)	0.97 (77.96)	0.98 (87.36)
ln $DIST_{ij}$	−0.65 (−16.81)	−0.84 (−20.95)	−1.06 (−27.65)	−1.07 (−28.68)	−1.20 (−33.00)
ADJ_{ij}	0.14 (0.93)	0.13 (0.78)	0.35 (2.24)	0.58 (3.65)	0.67 (6.90)
$LANG_{ij}$	0.05 (0.54)	0.27 (2.75)	0.55 (5.83)	0.79 (8.07)	0.65 (6.90)
EU_{ij}	0.67 (2.00)	0.48 (1.16)	−0.36 (−1.32)	−0.25 (−1.15)	−0.29 (−1.76)
$EFTA_{ij}$	0.56 (2.41)	1.04 (4.25)	0.32 (0.91)	−0.19 (0.41)	−0.98 (−0.71)
EEA_{ij}			−0.07 (−0.31)	−0.15 (−0.71)	−0.11 (−0.29)
$OEIA_{ij}$	0.72 (1.77)	3.01 (0.38)	0.86 (1.81)	0.61 (1.42)	0.61 (5.05)
Constant	−10.17 (−21.63)	−14.36 (−30.74)	−17.16 (−37.62)	−18.34 (−43.34)	−19.72 (−51.56)
RMSE	1.4144	1.7548	1.8935	1.9919	1.9616
R^2	0.6035	0.6364	0.6453	0.6651	0.7147
No. observations	2,789	4,030	5,494	6,474	7,302

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

TABLE 3
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 (−17.43)	−0.87 (−21.27)	−1.31 (−31.83)	−1.31 (−31.82)	−1.49 (−36.52)
ADJ_{ij}	0.36 (2.64)	0.39 (2.64)	0.43 (2.93)	0.56 (3.76)	0.52 (3.60)
$LANG_{ij}$	0.36 (3.84)	0.77 (7.70)	0.78 (7.82)	0.95 (9.19)	0.89 (8.91)
EU_{ij}	−0.84 (−2.87)	−1.23 (−3.41)	−2.26 (−8.65)	−1.54 (−7.41)	−1.26 (−7.68)
$EFTA_{ij}$	0.21 (1.01)	0.30 (1.32)	−0.62 (−1.89)	−0.74 (−1.73)	−0.72 (−0.56)
EEA_{ij}			−1.45 (−7.17)	−1.01 (−5.08)	−0.23 (−0.60)
$OEIA_{ij}$	0.67 (1.79)	3.27 (9.44)	0.05 (0.13)	−0.09 (−0.22)	0.40 (3.29)
Constant	−16.73 (−9.79)	−14.54 (−21.38)	−16.42 (−21.11)	−17.04 (−30.71)	−14.10 (−26.92)
RMSE	1.1806	1.4853	1.6638	1.7786	1.7757
Within R^2	0.5026	0.4433	0.3870	0.3665	0.3912
No. observations	2,789	4,030	5,494	6,474	7,302

Notes:
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.

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, EU_{ijt} 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 suggests that EU_{ij} and ε_{ij} are negatively correlated, and the *EU* coefficient estimate may be underestimated.

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-Second World War 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: '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; emphasis added).

We believe this 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 policy makers' 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 *EU*, *EFTA*, *EEA* and *OEIA* 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 (see Baier and Bergstrand, 2007). The reason basically is that – in cross-section – it is very difficult 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 appropriately *identify* the respective equations.[17]

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 (2002) used 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:

$$\ln[X_{ijt}/(RGDP_{it}RGDP_{jt})] = \beta_0 + \beta_3(\ln DIST_{ij}) + \beta_4(ADJ_{ij}) + \beta_5(LANG_{ij}) + \beta_6(EIA_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \ln \varepsilon_{ijt}, \qquad (6)$$

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 EU, $EFTA$, EEA and $OEIA$.

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. This estimated effect is both considerably larger and more robust to sensitivity analyses than earlier estimates.

In this paper, we examine in particular the effects of EU membership, EFTA membership, EEA membership, and membership in all other EIAs using these techniques. 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

[17] By this, we mean identification in the usual econometric sense needed for applying instrumental variables techniques appropriately. One may be able to provide identification using the 'functional forms' of the relationships, but some econometricians have reservations about this technique.

and Bergstrand (2007) to examine some specific agreements, allowing here for changing membership over the 40-year period from 1960 to 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 Western European trade agreements on members' international trade, accounting for the endogeneity of trade agreements' formation. Second, we want to establish that the economic effects of trade agreements 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 EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$, 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 EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$, 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 versus First Differencing

A standard discussion on the treatment of endogeneity bias using panel data focuses on a choice between estimation using fixed effects versus using first-differenced data (see Wooldridge, 2002, Ch. 10). Wooldridge notes that when the number of time periods (T) exceeds 2, a fixed-effects estimator is more efficient under the assumption of serially uncorrelated error terms. When $T > 2$ and the error term ε_{ijt} follows a random walk (i.e. that the difference in the error terms, $\varepsilon_{ijt} - \varepsilon_{ij,t-1}$, is white noise), the first-differencing estimator is more efficient.[18]

[18] 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}.

It is possible that 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. That is, unobservable (to the econometrician) variables such as domestic shipping regulations, which cause trade to be 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 becomes large. This suggests that differencing the data will increase estimation efficiency for our large T panel. Second, 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. Third, 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. In the following, we use fixed effects in Sections *c* and *d*, and for robustness we use differenced data in Section *e*.

c. Fixed-Effects Estimation of an Atheoretical Gravity Equation Ignoring Multilateral Price Terms

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

$$\ln X_{ijt} = \beta_0 + \beta_1(\ln RGDP_{it}) + \beta_2(\ln RGDP_{jt}) + \beta_3(\ln DIST_{ij}) \\ + \beta_4(ADJ_{ij}) + \beta_5(LANG_{ij}) + \beta_6(EIA_{ijt}) + \varepsilon_{ijt}. \tag{7}$$

Table 4 provides the empirical results of estimating gravity equation (7) using a panel of real trade flows (X_{ijt}), real GDPs ($RGDP_{it}$, $RGDP_{jt}$) and EIA dummies (EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$), 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.

However, other than $OEIA_{ij}$ the coefficient estimates for the Western European EIAs are quite unstable across agreements, suggesting fragile estimates. Although $EFTA_{ij}$ has an economically and statistically significant value of 0.33 (suggesting that EFTA increased trade by $e^{0.33}$ = 39 per cent), membership in various stages of the EEC/EC/EU had a statistically significant *negative* effect on members' trade, as did the EEA's EU–EFTA free trade agreements. Such results seem implausible.

Column (2) provides the empirical results including a time dummy, where (for brevity) we omit reporting the (statistically significant) coefficient estimates for these time dummies. Although the inclusion of the time dummies causes the

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*	*(5) With Time and Bilateral Fixed Effects, GDP Elasticities Restricted to Unity*
ln $RGDP_{it}$	0.95 (217.57)	0.98 (231.55)	0.71 (34.52)	1.27 (47.29)	
ln $RGDP_{jt}$	0.95 (225.07)	0.97 (236.17)	0.58 (26.53)	1.23 (41.72)	
ln $DIST_{ij}$	−1.04 (−78.42)	−1.02 (−78.34)			
ADJ_{ij}	0.38 (7.66)	0.34 (6.56)			
$LANG_{ij}$	0.60 (18.06)	0.53 (16.25)			
EU_{ijt}	−0.25 (−7.16)	−0.11 (−2.77)	0.58 (7.57)	0.89 (11.58)	0.82 (10.65)
$EFTA_{ijt}$	0.33 (7.36)	−0.17 (−3.49)	0.55 (4.23)	0.45 (3.48)	0.50 (3.88)
EEA_{ijt}	−0.12 (−2.83)	−0.11 (−2.53)	0.34 (3.92)	0.57 (6.64)	0.53 (6.24)
$OEIA_{ijt}$	0.72 (10.24)	1.12 (16.07)	0.57 (8.86)	0.65 (10.25)	0.63 (9.92)
RMSE	1.9252	1.8567			
Overall R^2	0.6582	0.6821			
Within R^2			0.2038	0.2273	0.0880
No. observations	47,081	47,081	47,081	47,081	47,081

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

RGDP elasticities to move closer to unity, the coefficient estimates for the time-invariant variables (distance, adjacency and language) are unaffected. However, coefficient estimates for EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ are all affected. Now, even the coefficient estimate for $EFTA_{ij}$ is surprisingly negative and statistically insignificant. Moreover, the $OEIA_{ij}$ coefficient estimate becomes very large, 1.12, implying that non-Western European EIAs on average increase trade by 200 per cent. This result also seems implausible. However, time dummies do not adjust for the endogeneity of EIAs.

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 coefficient estimates for EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ are now *all* plausible and are statistically significant. It is worth noting now that the coefficient estimates for EU_{ij}, $EFTA_{ij}$ and $OEIA_{ij}$ are also *all virtually identical* quantitatively (0.58, 0.55 and 0.57, respectively), each implying that the particular agreement increases trade by about 75 per cent. Membership in EEA_{ij} increases bilateral trade by about 40 per cent.[19]

Column (4) in Table 4 combines the inclusion of bilateral fixed effects and time dummies. One notable change occurred in the coefficient estimates for EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ for this specification relative to the previous one. First, the coefficient estimate for EU_{ij} increases substantively, suggesting that membership in the EU increased trade of the typical country pair during the period by 144 per cent. A second more minor difference is that the coefficient estimate for EEA_{ij} increased while those for $EFTA_{ij}$ and $OEIA_{ij}$ stayed approximately the same.

Column (5)'s specification differs from column (4)'s only by restricting the coefficient estimates for the (time-varying) real GDP variables to be unity. This reduces the overall explanatory power (within R^2), but has only minor implications for the EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ coefficient estimates.

Overall, the inclusion of bilateral fixed effects and time-varying dummies has made the coefficient estimates for EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ more economically plausible and statistically significant. If (as we will argue shortly) the effects of an EIA on trade took 15 years to play themselves out, the coefficient estimates from column (5) imply that common membership in the EU (beginning with the original six EEC countries) increased trade (in real terms)

[19] The only other published studies that have estimated the ATE of an EIA using a panel of data spanning as many years and countries are Rose (2004) and Tomz et al. (2004). Using fixed effects, Rose found an ATE of $e^{0.94}$ or 156 per cent. However, using a classification of formal and informal GATT members, Tomz et al. (2007) estimate an ATE for EIAs (with fixed effects) of only $e^{0.76}$ or 114 per cent. Cheng and Wall (2005) 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.

about 5.6 per cent annually over 15 years. Common membership in EFTA (or the EC–EFTA trade pacts) increased trade by about 3.5 per cent annually and membership in any other EIA increased trade by about 4.3 per cent annually.

How do these results compare to previous ones? Bayoumi and Eichengreen (1997) examined the impacts of common membership in the original EEC6 and in the original EFTA7, but only over a much shorter period, 1956–73. They found implied annualised impacts of only 3.2 and 2.3 per cent, respectively, over the period. These are significantly lower than our estimates of 5.6 and 3.5 per cent annually, respectively, over 1960–2000.[20] By contrast, our estimate for $OEIA_{ij}$ membership was 0.63, which is considerably lower than comparable estimates using similar specifications in Rose (2004) and Tomz et al. (2007) of 0.94 and 0.76, respectively.

However, we emphasise that all these estimates used an 'atheoretical' specification for the gravity equation. If we account for recent theoretical advances in 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.

d. Fixed-Effects Estimation of a Theoretically-Motivated Gravity Equation with Phased-In Agreements

In this section, we consider three modifications to the previous specification. Initially, we include country-and-time effects to account for the theoretically-motivated multilateral price terms. Then we move on to 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. Finally, we address 'strict exogeneity' issues; we test for the possibility of reverse causality by addressing the effect of future EIA dummies on current trade flows.

(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 2 – suggests that one needs to account for the multilateral price variables. None of the four specifications in Table 4 accounts for these. First, accounting for the multilateral price variables in a panel context suggests estimating:

[20] Bayoumi and Eichengreen (1997) only estimated models using first-differenced data, which may not be appropriate given the error structures discussed earlier.

$$\ln X_{ijt} = \beta_0 + \beta_1(\ln RGDP_{it}) + \beta_2(\ln RGDP_{jt}) + \beta_3(\ln DIST_{ij}) + \beta_4(ADJ_{ij}) + \beta_5(LANG_{ij}) + \beta_6(EIA_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \varepsilon_{ijt}. \quad (8)$$

As before, scaling the left-hand side (LHS) variable by the product of real GDPs suggests estimating:

$$\ln[X_{ijt}/(RGDP_{it}RGDP_{jt})] = \beta_0 + \beta_3(\ln DIST_{ij}) + \beta_4(ADJ_{ij}) + \beta_5(LANG_{ij}) + \beta_6(EIA_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \varepsilon_{ijt}. \quad (9)$$

In a panel setting, the multilateral price variables would be time varying, and consequently the results in specifications (1)–(5) 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 and time dummies using the panel data in its current form.[21] Moreover, the theoretical model in equation (2) suggests that the coefficient estimates for the real GDP variables should be unity, as reported in specification (5) in Table 4.

We first estimate equation (8) 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.[22]

Column (1) in Table 5 provides the results of estimating this equation using bilateral fixed effects and the country-and-time effects. We note two observations. First, all the coefficient estimates for the effects of EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ on trade are diminished (relative to those in Table 4) by accounting for the theoretically-motivated multilateral price terms. Second, there is a notable change in the relative effects of the agreements. Common membership in the European Union (or, as appropriate in early years, EEC or EC) declines only slightly. Membership in the EU still increased trade by almost 100 per cent. Membership in any other EIA increased trade by almost 60 per cent. However, membership in EFTA had no effect. The EC–EFTA free trade agreements that began in 1973, and continued in the 1994 EEA agreement, boosted trade by about 20 per cent, considerably less than in the results in Table 4.

Column (2) of Table 5 imposes explicitly unitary elasticities for real GDPs. However, in the presence of the *it* and *jt* dummies, this restriction is redundant,

[21] 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.

[22] As noted in footnote 14, the estimate of β_6 reflects the primary 'direct' (or partial) effect associated with EIA membership, and not the full general equilibrium comparative static effect addressed in Anderson and van Wincoop (2003).

TABLE 5
Panel Gravity Equations with Bilateral Fixed and Country-and-Time Effects

Variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*
EU_{ijt}	0.65** (7.86)	0.65** (7.85)	0.45** (4.01)	0.47** (3.90)	0.50** (3.74)
EU_{ijt-1}			0.37** (3.13)	0.19 (1.36)	0.04 (0.29)
EU_{ijt-2}				0.24* (1.78)	0.26 (1.57)
EU_{ijt+1}					−0.08 (−0.63)
$EFTA_{ijt}$	−0.01 (−0.09)	−0.01 (−0.11)	−0.18 (−1.10)	−0.12 (−0.61)	0.04 (0.16)
$EFTA_{ijt-1}$			0.29* (1.83)	0.13 (0.60)	0.17 (0.74)
$EFTA_{ijt-2}$				0.07 (0.41)	−0.05 (−0.28)
$EFTA_{ijt+1}$					−0.22 (−1.02)
EEA_{ijt}	0.19* (2.11)	0.19* (2.10)	0.05 (0.48)	0.10 (0.85)	0.19 (1.61)
EEA_{ijt-1}			0.29** (2.85)	0.09 (0.76)	0.06 (0.47)
EEA_{ijt-2}				0.27** (2.51)	0.13 (1.00)
EEA_{ijt+1}					−0.24* (−1.66)
$OEIA_{ijt}$	0.46** (7.02)	0.46** (7.01)	0.31** (4.55)	0.29** (4.10)	0.39** (3.64)
$OEIA_{ijt-1}$			0.46** (4.77)	0.37** (3.52)	0.29* (1.79)
$OEIA_{ijt-2}$				0.17 (1.26)	0.11 (0.67)
$OEIA_{ijt+1}$					−0.04 (−0.58)
Constant	8.43 (279.58)	−25.05 (−870.87)	8.92 (346.63)	9.00 (263.34)	9.16 (282.92)
Within R^2	0.3106	0.1896	0.3050	0.2759	0.2523
No. observations	47,081	47,081	36,563	34,105	27,575

Notes:
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. * (**) denote statistical significance at 5 (1) per cent level in one-tailed *t*-test. Coefficient estimates for bilateral fixed and country-and-time effects are not reported for brevity.

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-and-time (*it*, *jt*) effects, and only the estimates of the intercept and the country-and-time effects' coefficients change; the *EIA* coefficient estimate is unaffected. In the remainder of the results, we use the real trade flow for the LHS variable; the EIA coefficient estimates are identical using trade shares instead (and are available on request).

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

In this section, we introduce lagged effects of EIAs on trade. The economic motivation for including lagged changes stems partly from the institutional nature of virtually all EIAs. The 0–1 EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ 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', typically over ten years. For instance, the original EEC agreement of 1958 had a ten-year phase-in period; NAFTA had a similar ten-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 (e.g. $EU_{ij,t-1}$ and/or $EU_{ij,t-2}$). Since our data is a panel with five-year intervals, $t-1$ ($t-2$) denotes a variable lagged five (ten) years.

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 columns (3) and (4) in Table 5 reveal that EU_{ij} has a statistically significant lagged effect on trade flows. Moreover, the coefficient estimates have economically plausible values, balanced across periods. In column (3), the sum of the two ATEs for EU_{ij} is 0.82 – identical in magnitude to the EU_{ij} coefficient estimate in column (5) of Table 4. With two lags, the coefficient estimate for one of the two lagged terms is statistically insignificant; however, summing the coefficient estimates yields a total ATE of 0.90. Since this ATE reflects the effect of EU membership over approximately 15 years, the implied average annual effect on members' trade across the 15-year transition period is 6.2 per cent. This is only slightly larger than our earlier estimate (using the atheoretical gravity equation), and is roughly twice the average annual ATE found in Bayoumi and Eichengreen for the original EEC6 countries.

We will discuss the results and implications for all other EIAs ($OEIA_{ijt}$) in Section 5.

(iii) Strict exogeneity

The results of the previous sections suggest that – after accounting for endogeneity using panel data – one can find economically significant ATEs for *EIA*. However, to confirm that there are no 'feedback effects' from trade changes to *EIA* changes, we run one more specification using the fixed-effects approach.[23] Wooldridge (2002, p. 285) suggests that it is easy to test for the 'strict exogeneity' of EIAs in our context. To do this, we add future levels of EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ to the regression model. In the panel context here, if EU_{ij}, $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$ changes are strictly exogenous to trade flow changes, $EU_{ij,t+1}$, $EFTA_{ij,t+1}$, $EEA_{ij,t+1}$ and $OEIA_{ij,t+1}$ should be uncorrelated with the concurrent trade flow. The results in column (5) of Table 5 confirm this. In only one case did $EU_{ij,t+1}$, $EFTA_{ij,t+1}$, $EEA_{ij,t+1}$ and $OEIA_{ij,t+1}$ affect the trade flow X_{ijt} materially; except for $EEA_{ij,t+1}$, in all cases the coefficient estimate is not significantly different from zero. Moreover, the consistently negative coefficient estimates suggest, if anything, that firms delay trade temporarily in anticipation of an impending agreement.

e. First-Differenced Panel Gravity Equation Estimates

As discussed in Section *b*, for econometric reasons one might expect first-differenced data to provide better estimates of the average treatment effect than using 'fixed effects'. 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 the EIAs' ATEs. As before, with country-and-time effects the coefficient estimates of the EIA treatment effects are insensitive to the real bilateral trade flow being scaled or not scaled by real GDPs; for consistency with earlier results, we present those for the flows (the virtually identical results are available on request using trade flows scaled by the product of real GDPs). We start by first-differencing the natural logarithm of X_{ijt}, creating $d\ln X_{ij,t-(t-1)}$. As before, since our data set is a panel with five-year intervals, $t-(t-1)$ represents a five-year difference. Second, we regress $d\ln X_{ij,t-(t-1)}$ on 768 country-and-time effects ($Dum_{i,t-(t-1)}$, where i denotes a country and $t-(t-1)$ a five-year period, e.g. 1995–2000) and retain the residuals. Third, we difference EU_{ijt}, creating $dEU_{ij,t-(t-1)}$, and regress $dEU_{ij,t-(t-1)}$ on the same 768 country-and-time fixed effects and retain these residuals (and do the same for $EFTA_{ij}$, EEA_{ij} and $OEIA_{ij}$). Fourth, a regression of the residuals from the first ($d\ln X$) regression on the residuals from the other

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

regressions will yield unbiased estimates of the ATE effect of an EIA holding constant time-varying multilateral price terms.

The procedure described above is equivalent to estimating:

$$d\ln X_{ij,t-(t-1)} = \beta_6 dEIA_{ij,t-(t-1)} + \beta_{i,t-(t-1)} Dum_{i,t-(t-1)} + \beta_{j,t-(t-1)} Dum_{j,t-(t-1)} + \nu_{ij,t-(t-1)}, \tag{10}$$

where $dEIA$ represents any of the four trade agreements we have been investigating and $\nu_{ij,t-(t-1)} = \varepsilon_{ijt} - \varepsilon_{ij,t-1}$ is white noise. With nine years in the panel, we have eight time periods $t-(t-1)$. Since there are 96 countries that can potentially trade, our procedure above effectively introduces 768 $(= 8 \times 96)$ country-and-time fixed effects ($Dum_{i,t-(t-1)}$ and $Dum_{j,t-(t-1)}$) to account for the changes in the unobservable theoretical multilateral resistance terms, $d\ln P^{1-\sigma}_{i,t-(t-1)}$ and $d\ln P^{1-\sigma}_{j,t-(t-1)}$, 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 four agreements on trade flow changes. For the European

TABLE 6
First-Differenced Panel Gravity Equations with Country-and-Time Effects

Variable	*(1)*	*(2)*	*(3)*	*(4)*
$EU_{ij,t-(t-1)}$	0.48** (8.91)	0.47** (8.63)	0.46** (8.54)	0.46** (8.16)
$EU_{ij(t-1)-(t-2)}$		0.23** (4.41)	0.19** (3.70)	0.04 (0.72)
$EU_{ij(t-2)-(t-3)}$			−0.11** (−2.82)	−0.07 (−1.17)
$EU_{ij(t+1)-t}$				0.06 (0.82)
$EFTA_{ij,t-(t-1)}$	0.08 (1.28)	0.02 (0.27)	0.01 (0.85)	0.03 (0.40)
$EFTA_{ij(t-1)-(t-2)}$		0.20** (3.09)	0.14* (2.06)	0.23** (2.74)
$EFTA_{ij(t-2)-(t-3)}$			0.02 (0.23)	−0.01 (−0.13)
$EFTA_{ij(t+1)-t}$				−0.25* (−2.25)
$EEA_{ij,t-(t-1)}$	0.19** (4.02)	0.17** (3.49)	0.16** (3.43)	0.15** (2.92)
$EEA_{ij,(t-1)-(t-2)}$		0.06 (1.40)	0.05 (1.08)	0.05 (1.00)
$EEA_{ij(t-2)-(t-3)}$			−0.02 (−0.40)	−0.01 (0.09)
$EEA_{ij(t+1)-t}$				−0.20** (−2.59)
$OEIA_{ij,t-(t-1)}$	0.31** (6.66)	0.30** (6.30)	0.28** (6.04)	0.27** (4.55)
$OEIA_{ij(t-1)-(t-2)}$		0.29** (4.57)	0.25** (3.79)	0.30 (1.72)
$OEIA_{ij(t-2)-(t-3)}$			0.05 (0.29)	0.04 (0.21)
$OEIA_{ij(t+1)-t}$				−0.06 (0.91)
Constant	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)
R^2	0.0009	0.0011	0.0011	0.0010
No. observations	36,563	34,105	31,172	24,642

Notes:
t-Statistics are in parentheses. The dependent variable is the (natural log of the) real bilateral trade flow from country *i* to country *j*. * (**) denote statistical significance at 5 (1) per cent level in one-tailed *t*-test. Coefficient estimates for bilateral fixed and country-and-time effects are not reported for brevity.

Union, columns (1)–(4) all report slightly smaller coefficient estimates for the *EU* effect than the respective estimates in Table 5 using fixed effects. For EFTA, the results are more plausible. However, as in Table 5, the effects of EFTA are quite small. Using first differences, the effects of EC–EFTA free trade agreements are small as well, but largely similar to those in Table 5. As with the EU, the effects for all other EIAs are diminished using first differences relative to fixed effects.

The major point worth noting from an empirical standpoint is that the results using first differencing provide strong support for the robustness of the previous estimates in this section using fixed effects for the theoretically-motivated gravity equation. Membership in the EEC/EC/EU had an economically and statistically significant effect on trade among members between 1960 and 2000. This result is robust across many specifications. The small variation in results, say, between column (2) in Table 5 and Table 6 – total ATEs of 0.82 and 0.70, respectively (depending upon one's preferences over underlying assumptions about the error structure) – suggest that these results are fairly precise and robust. In average annual percent changes, the two effects are 5.6 and 4.8 per cent, respectively, over a 15-year period. For all other EIAs, the results for the two approaches (using column (2) results again) are 0.77 and 0.59.

5. IMPLICATIONS FOR UNDERSTANDING THE 'LATEST WAVE' OF REGIONALISM

What do these empirical results mean for better understanding the 'latest wave' of regional trade and cooperation 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 these constituents' economic status (profits and consumer welfare, respectively). Such policy makers are likely making decisions about trade policies in a competitive 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. The world market for goods and services 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, but are also remote from the rest of the world; (2) that are large and similar in economic size and consequently benefit from greater specialisation in production and variety in terms of consumption; and (3) that differ in relative factor endowments, benefiting from the exchange of traditional comparative advantages. Our probit estimates of the determinants of EIAs confirmed this. Hence, the vast bulk of EIAs are among countries that trade extensively; that is, countries that have formed EIAs have chosen well.

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 EEC/EC/EU 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 (Frankel, 1997). One reason is that the approach taken here does not require measurement of the 'complex and elaborate' barriers (beyond tariff cuts) that EIA agreements often liberalise. The results here suggest that EEC/EC/EU membership over the past 40 years (1960–2000) is of an economically significant magnitude and even larger than that postulated a decade ago in Bayoumi and Eichengreen's excellent analysis of EEC6 effects between 1957 and 1972.

Policy makers beyond Europe 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 similar expected trade enhancements from bilateral and regional EIAs. And the evidence in this paper suggests that their 'economic expectations' have largely been correct. Our results suggest that other EIAs that have formed over the 1960–2000 period have also yielded 'average treatment effects' of nearly the same magnitudes as the trade effects of EEC/EC/EU membership (notwithstanding the issues raised in Anderson and van Wincoop, 2003, discussed earlier). Naturally, the deeper integration of the EU has likely boosted the trade effects of that particular agreement relative to most other agreements, which have been FTAs.

Our overall message is twofold. First, *ex post* empirical evidence is consistent with the notion that policy makers are operating in a competitive environment, pursuing economic integration agreements in 'natural cases' where the members already trade extensively (based upon bilateral, multilateral and world levels of GDP and trade costs). Second, after accounting for the pitfalls associated with the 'endogeneity of country pairs that select into EIAs', 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.

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.	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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