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TRADE CREATION AND DIVERSION REVISITED: ACCOUNTING FOR MODEL UNCERTAINTY AND NATURAL TRADING PARTNER EFFECTS

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SUMMARY

The effect of preferential trade agreements (PTAs) on trade flows is subject to model uncertainty stemming from the diverse and even contradictory effects suggested by the theoretical PTA literature. The existing empirical literature has produced remarkably disparate results and the wide variety of empirical approaches reflects the uncertainty about the 'correct' set of explanatory variables that ought to be included in the analysis. To account for the model uncertainty that surrounds the validity of the competing PTA theories, we introduce Bayesian model averaging (BMA) to the PTA literature. Statistical theory shows that BMA successfully incorporates model uncertainty in linear regression analysis by minimizing the mean squared error, and by generating predictive distributions with optimal predictive performance. Once model uncertainty is addressed as part of the empirical strategy, we find strong evidence of trade creation, trade diversion, and open bloc effects. Our results are robust to a range of alternative empirical specifications proposed by the recent PTA literature.

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1. INTRODUCTION

Bhagwati and Panagariya (1996) call preferential trading arrangements (PTAs) 'two faced' because PTAs introduce trade liberalization at the cost of discrimination. The controversy regarding the costs and benefits of PTAs has raged since the 1950s, due to the potential for trade creation and trade diversion (Viner, 1950). Time has not provided a consensus; to the contrary, with the proliferation of PTAs in the 1990s, the number of PTA theories that predict either increasing or decreasing trade flows among (non-)members increased in tandem. And as the number of theories expanded, the set of candidate regressors suggested by empirical PTA research approached the point where comprehensive robustness has become virtually unfeasible. Consequently, it has become common practice in this literature to juxtapose results that represent alternative PTA theories. It is therefore not surprising that PTA coefficient estimates have been found to be highly sensitive to the specific set of regressors used in any given study (see Baxter and Kouparitsas, 2006).

Ghosh and Yamarik (2004) provide the most extensive PTA robustness analysis to date. Not only do they include a large set of PTAs, but they also employ extreme bound analysis (Leamer, 1983) to examine a diverse set of PTA theories. Ghosh and Yamarik (2004) find little evidence for either trade-creating or trade-diverting PTAs. They conclude that 'the pervasive trade creation effect found in the literature reflects not the information content of the data but rather the unacknowledged beliefs of the researchers'.

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In this paper we apply Bayesian model averaging (BMA) to the PTA literature to re-examine model uncertainty. BMA is specifically designed to incorporate model uncertainty into the estimation process and is firmly rooted in statistical theory. It is a methodology that explores the model space without restrictions, weighs each model according to quality, and provides a probability distribution for each coefficient estimate. Raftery and Zheng (2003) show that BMA maximizes predictive performance while minimizing the total error rate when compared to any individual model. The rapidly growing list of economics applications using BMA include policy evaluations (e.g. Brock *et al*., 2003), monetary policy (e.g. Levin and Williams, 2003), macroeconomic forecasting (e.g. Garratt *et al*., 2003), economic growth (e.g. Fernandez *et al*., 2001), and international economics (e.g. Chen and Rogoff, 2006).

The issue of model uncertainty surrounding PTA effects is well known in the PTA literature. Seldom do papers present less than a dozen different PTA regression specifications. We show that BMA overturns the fundamental Ghosh and Yamarik result by identifying a number of PTAs that exert decisive effects on trade flows. Since Ghosh and Yamarik, the PTA literature has evolved to introduce a number of innovations that address omitted variable bias. We show that our main finding of measurable PTA effects on trade flows is robust, even when the Ghosh and Yamarik (2004) dataset is updated to include additional years, additional PTAs, and alternative fixed-effect specifications.[1] Our methodological extensions include a full account of multilateral resistance (see, for example, Anderson and van Wincoop, 2003; Subramanian and Wei, 2007), bilateral unobserved heterogeneity (see, for example, Glick and Rose, 2002; Egger and Pfaffenmayr, 2003), and an approach to control for both multilateral resistance and heterogeneity simultaneously (Baier and Bergstrand, 2007; Baldwin and Taglioni, 2006). We also consider accession dynamics (Freund and McLaren, 1999). Our analysis follows a voluminous literature spanned by Frankel *et al*. (1995, 1997), Rose and van Wincoop (2001), Frankel and Rose (2002), and Rose (2004).[2]

Our BMA benchmark specification, using Ghosh and Yamarik's (2004) original dataset, shows strong trade creation, trade diversion, and open bloc effects for 12 PTAs.[3] Our results are at odds with Ghosh and Yamarik (2004), even if we use their identical dataset. The differences arise for the following two reasons. First, BMA inference is based on an unrestricted search of the model space spanned by all candidate regressors, while extreme bound analysis covers only a fraction of the model space due to the researcher's categorization of variables into 'free' (variables that should always be included in the regression specification) and 'doubtful' (variables that *may* be effective in the regression specification). Second, BMA theory requires that each model is weighed according to its posterior model probability (which is associated with the model's quality or performance), while extreme bound analysis weighs all models equally and thus attributes the same power of inference to both strong and exceptionally weak models.[4]

Even after we extend the Ghosh and Yamarik data from 1970–1995 to 1960–2000 and include more recent bilateral trade agreements, our results remain robust. In fact, a number of PTAs are estimated with increased precision, which allows us to identify additional trade-creating PTAs. The updated dataset also modifies the counterintuitive trade diversion effects (for NAFTA) and

[1] It is important to note that most of the literature has ignored general equilibrium effects and estimates. The primary goal of this paper is to flag more robust estimates of the 'partial' or direct effects of PTAs and other controls, in order to provide potentially better inputs for general equilibrium comparative statics.

[2] An appealing alternative is to examine the intensive and extensive margins of trade as proposed by Helpman *et al*. (2008) and Felbermayr and Kohler (2006). We leave this to future research.

[3] It is common in extreme bound analysis to attach all the weight of the posterior to the prior distribution. While extreme bound analysis provides no guidelines, Ghosh and Yamarik (2004) also examine the case where 95% of the weight of the posterior distribution is on the prior and 5% on the sampling distribution—in this case they find trade creation in four PTAs (CACM, CARICOM, MERCOSUR and APEC).

[4] Previous comparisons between extreme bound analysis and BMA results have also found extreme bound analysis to be excessively stringent (see Sala-i-Martin, 1997; Fernandez *et al*., 2001).

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the unexpectedly large open bloc effects (for MERCOSUR) that were implied by the Ghosh and Yamarik data. Controlling for multilateral resistance does not affect our result qualitatively, and the vast majority of PTAs are shown to exert influence on trade flows, mostly through trade creation among member countries.

Our approach to addressing multilateral resistance follows directly from Anderson and van Wincoop (2003) and Novy (2006, 2007), as implemented by Subramanian and Wei (2007), whose context was different and it did not address individual PTA effects. We also show that estimates based on multilateral resistance are generally larger than estimates that account for unobserved country-pair heterogeneity (an approach advocated by Glick and Rose, 2002; Rose, 2004, 2005). This may be due to the methodological difference, whereby country-pair fixed effects render estimates that measure only those PTA effects that are directly related to accession. This raises the question of accession dynamics; we show that PTA trade effects generally appear around accession or thereafter.[5]

Our most comprehensive specification controls simultaneously for multilateral resistance and unobserved heterogeneity among countries. This specification is inspired by Baier and Bergstrand (2007), who produce a similar specification, but without emphasis on the heterogeneous effects that individual PTAs exert on trade flows. Even in this most comprehensive specification, we find strong effects of PTAs on trade flows, for the Andean Pact, Central American Common Market, European Economic Area, Latin American Integration Association, and for bilateral trade agreements.

The remainder of the paper is organized as follows. Section 2 discusses the basic framework of the BMA methodology used in our estimation. In Section 3 we take a look at the datasets employed, and in Section 4 we report and discuss our results. Section 5 concludes.

2. THE EMPIRICAL FRAMEWORK

2.1. Baseline Specification

Econometric studies that seek to identify the impact of PTAs on trade flows are generally based on the gravity model.[6] The approach fits the application particularly well, due to the gravity model's proven efficiency in predicting trade flows (see Frankel and Romer, 1999). This allows PTA coefficients to pick up on deviations between predicted and actual trade.

Ghosh and Yamarik (2004) include dummies that capture PTA effects on bilateral trade alongside a matrix of other covariates, Z_{ijt},[7] obtaining

$$\log T_{ijt} = \alpha_t + \beta_1 \log Y_{it} Y_{jt} + \beta_2 \log D_{ij} + \beta_3 Z_{ijt} + \beta_4 \mathrm{PTA}_{ijt} + \beta_5 \mathrm{PTA}_{it} + \varepsilon_{ijt} \tag{1}$$

where average bilateral trade, T_{ijt}, between countries i and j at time t depends positively on national incomes, Y_{it} and Y_{jt}, and negatively on bilateral distance, D_{ij}. The matrix of other covariates, Z_{ijt}, is included to represent alternative trade theories and to proxy for unobservable trade costs. The inclusion of time fixed effects, α_t, is standard in the literature to eliminate bias resulting from aggregate shocks to world trade, such as global income shocks. Time fixed effects

[5] The accession dynamics results are interesting in light of the emerging 'endogenous PTA' strand of literature (Baier and Bergstrand, 2007). This paper does not address endogeneity explicitly, although regressors used to control for endogeneity by Baier and Bergstrand (2007) are included here. We discuss endogeneity bias later on. Similarly, while we are unable to prove that our specifications are free of omitted variable bias, our expanded dataset is one of the most comprehensive to date.

[6] The theoretical foundations of the gravity model are presented in Frankel (1997) and Deardorff (1998).

[7] The set of specific correlates used is discussed in Section 2.3.

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also mitigate any spurious correlation introduced, for example, by the use of a US price index to deflate all trade flows. To capture PTA effects, two sets of zero–one dummy variables are included for each time interval, t. PTA_{ijt} indicates that both trading partners are members of the same PTA in a given year, and PTA_{it} indicates that only one member has joined. These dummies enable us to isolate the three distinct effects that PTAs may exert on trade flows. A positive coefficient on PTA_{ijt} captures trade creation among PTA members, while trade diversion registers a negative PTA_{it} coefficient. Finally, open bloc trade creation is simply the opposite of trade diversion, characterized by a positive PTA_{it} coefficient.

2.2. Multilateral Resistance and Unobserved Heterogeneity

Equation (1) can be extended to control for multilateral resistance and unobserved country-pair heterogeneity. In place of average trade, multilateral resistance requires the use of either bilateral imports (Subramanian and Wei, 2007) or bilateral exports (Novy 2006, 2007) as the dependent variable.[8] Here we largely follow Subramanian and Wei (2007) to generate results that are comparable to their benchmark:

$$\log(\text{Imports}_{ijt}) = \alpha_t + \alpha_{it} + \alpha_{jt} + \beta_2 \log D_{ij} + \beta_3 \tilde{Z}_{ijt} + \beta_4 PTA_{ijt} + \varepsilon_{ijt} \tag{2}$$

The added advantage of using bilateral imports, Imports_{ijt}, as the dependent variable is that it avoids bias induced from averaging trade flows (see Baldwin and Taglioni, 2006).[9] Since any nation faces only one import/export price index at any point in time, multilateral resistance can be accounted for with time-varying importer/exporter fixed effects (represented by α_{it} and α_{jt}).[10] The inclusion of time-varying importer/exporter effects does not allow for average trade flows as the dependent variable and we follow Subramanian and Wei (2007) and choose bilateral imports instead. Multilateral resistance controls in (2) absorb some of the covariates, which reduces Z_{ijt} to $\tilde{Z}_{ijt}$. Most notably the remoteness measure is now absorbed. Remoteness speaks only to GDP-weighted geographic distance, which changes only slightly over time because the GDP weights are time-varying (see Section 2.3). Multilateral resistance, instead, also accounts for variations in prices of all trading partners over time, which can imply considerable fluctuations.

In addition, multicollinearity no longer allows for the identification of separate trade creation and diversion effects. In the presence of time-varying importer effects, the PTA_{ijt} dummy partitions an importer's observations in any given year into (a) imports originating from fellow PTA members and (b) imports from non-members. As a consequence, the PTA_{ijt} dummies now express *net trade creation*, or how much greater intra-PTA trade is compared to trade between PTA members and non-members. This implies that when trade between members and non-members decreases because of trade diversion, the PTA_{ijt} coefficient increases in this specification.

Unobserved country-pair heterogeneity can be addressed by controlling for all time-invariant bilateral heterogeneity with country-pair fixed effects, α_{ij}, as follows:

$$\log(\text{Imports}_{ijt}) = \alpha_t + \alpha_{ij} + \beta_1 \log Y_{it} Y_{jt} + \beta_3 \overline{Z}_{ijt} + \beta_4 \text{PTA}_{ijt} + \varepsilon_{ijt} \tag{3}$$

[8] Some argue that this is advantageous, since trade theories yield predictions on unidirectional trade (see Freund, 2000; Anderson and van Wincoop, 2003; Baldwin and Taglioni, 2006).

[9] Alternative estimation approaches can also address measurement error bias (see Felbermayr and Kohler, 2006; Santos Silva and Tenreyro, 2006).

[10] Time-varying importer/exporter fixed effects are lucidly motivated by Baldwin and Taglioni (2006).

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Note that now all time-invariant regressors are absorbed into the pair-specific fixed effects.[11] Pair fixed effects capture similarities of trading partners that are constant over time. With these pair-specific constants, our regression only relies on time series variation, comparing each country pair's observations before and after PTA accession to determine the PTA_{ijt} coefficient. Therefore here, as in equation (1), PTA_{ijt} expresses only intra-PTA trade creation. The country-pair fixed-effect specification, together with Rose's remoteness variable to (imperfectly) capture multilateral resistance, represents a general formulation of the gravity equation to address unobserved heterogeneity (e.g. Egger, 2000; Baldwin, 2005). If country-pair fixed effects are omitted, the PTA coefficients tend to be biased upward because they pick up trade creation that is not specifically PTA related, but simply due to unobservables. The introduction of country-pair fixed effects absorbs non-time-varying control variables, which reduces the original matrix of other covariates to $\overline{Z}_{ijt}$ (in equation (3)) to $\tilde{Z}_{ijt}$ (in equation (2)).

The most comprehensive approach to controlling for unobserved heterogeneity, multilateral resistance, and all other unobserved time-varying importer and exporter specific effects, is to combine (2) and (3). This yields a specification that is most likely to generate unbiased coefficient estimates, while adhering to theoretical foundations. This specification was suggested by Baier and Bergstrand (2007) in the context of estimating average trade effects across all PTA member countries. It can be obtained by adding country-pair (*ij*) fixed effects along with the time-varying importer/exporter (*it*, *jt*) fixed effects to equation (1):[12]

$$\log(\mathrm{Imports}_{ijt}) = \alpha_{ij} + \alpha_{it} + \alpha_{jt} + \beta_2 \log D_{ij} + \beta_3 \tilde{Z}_{ijt} + \beta_4 PTA_{ijt} + \varepsilon_{ijt} \tag{4}$$

Baier and Bergstrand (2007) also point out that in panel data fixed effects or first differencing can be employed to address some of the potential endogeneity in the PTA regressions.[13] The fixed effects in equations (2)–(4) can partially address two out of three sources of endogeneity bias. The first type of endogeneity bias may arise in equation (1) between GDP and trade flows (see Frankel and Romer, 1999). The inclusion of time-varying importer/exporter fixed effects will contain this source of bias. The second type of bias arises due to the endogeneity between trade flows and trade policies. Trefler (1993) first used instruments to addess the endogeneity of trade policies and found the effect of such policies to increase tenfold. Lee and Swagel (1997) also document that the effect of trade liberalization on imports is biased downward in the absence of instrumenting for endogeneity.

The third source of possible endogeneity is that countries might endogenously select into (specific) PTAs. This bias is less likely addressed by fixed effects. Baier and Bergstrand (2004a) find cross-section evidence that country pairs with common economic characteristics also tend to share PTA memberships. Baier and Bergstrand (2007) suggest that the endogeneity of PTA membership likely renders the PTA coefficient biased downward in cross-sections.[14]

[11] We estimate equations (1) using the Andrews *et al*. (2006) 'FEiLSDVj' estimator, which relies on partitioned regression techniques to reduce computational burden; it delivers identical results to LSDV regressions.

[12] The alternative would be to first-difference. Wooldridge (2002, Ch. 10) shows that when the number of time periods exceeds two, the fixed-effects estimator is more efficient under the assumption of serially uncorrelated error terms. Baier and Bergstrand (2007) provide a comprehensive discussion of the two approaches whose results might differ slightly depending on the length of the panel and the structure of the error terms. Although both approaches have advantages and disadvantages, Baier and Bergstrand show in a panel that is basically identical in error structure to ours that the results are very similar. Hence we present the fixed-effects results below.

[13] Aside from Baier and Bergstrand (2007), the potential PTA endogeneity bias in cross-section gravity models is also addressed by Baier and Bergstrand (2002, 2004b) and Magee (2003), but with mixed success. Baier and Bergstrand (2007) also lament that 'other methods to identify the impact, such as instrumental variables using cross-section data, are compromised by a lack of suitable instruments'.

[14] To paraphrase, their reasoning is that PTA membership of a trading pair and the intensity of their domestic regulations may be positively correlated in a cross-section of data, but the gravity equation's error term and the intensity of domestic

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They argue that a key source of this endogeneity may be bilateral unobserved characteristics, for example, common institutions or regulations. Such unobserved bilateral characteristics may be the determinants of countries' trade and of their PTA membership decisions. In this case the endogeneity bias would be largely cross-sectional in nature, and it can be controlled by the country-pair fixed effects that we include in our panel regressions.

Egger (2004) argues that the estimates obtained in the regressions above may be biased downward if there exist cross-section dependencies that result in correlations between explanatory variables and unobserved bilateral effects exists. Thus he proposes the Hausman and Taylor (1981) two-stage least squares error components model. Serlenga and Shin (2007) incorporate the Egger methodology into the correlated common effect pooled (CCEP) estimation approach, which was advanced by Pesaran (2006). Serlenga and Shin (2007) highlight that the bias can go either way, depending on the specific coefficient estimate and time period examined. This is not surprising since the exact bias depends on the specific correlation structure. Unfortunately, the methods developed by Hausman and Taylor (1981), Pesaran (2006), and Serlenga and Shin (2007) are not available for BMA. Thankfully, however, this type of bias is not of crucial importance in our application since we focus primarily on the coefficients of the time-varying PTA dummies (which are still estimated consistently) and not on the magnitude of time-invariant regressors (such as distance).

Serlenga and Shin (2007) further show that the gravity equation may be biased due to possible cross-section dependence arising from unobserved (heterogeneous) time-specific factors. The authors thus adopt an alternative estimator, originally proposed by Pesaran (2006), to explicitly address such dependencies. The Pesaran estimator has not been implemented in a BMA context; hence we limit ourselves to the approaches that we introduced above. While we are mindful of this bias, we nevertheless regard it as important to consider results that have been obtained via a principled approach to model uncertainty. If anything, the previous literature seems to suggest that the derived estimates are generally too low.

Another bias may arise due to spatial heterogeneity (structural instability or heteroskedasticity; see Anselin and Griffith, 1988). This may lead to biased parameter estimates or misleading significance levels. Bougheas *et al*. (2003) explore the spatially autocorrelated error terms and use instrumental variables in an attempt to address the issue. Their results show the bias can go either way, depending on the application. Baltagi *et al*. (2007) also highlight the importance of spatial autoregressive error processes that apply to both the individual and remainder error components. They suggest a maximum likelihood estimator for a general spatial panel with random effects. Here we presume, consistent with previous literature, that the fixed-effects model is predominant for our trade application.

2.3. Model Uncertainty in PTA Theory

A voluminous theoretical literature discusses appropriate controls in gravity models, which include proxies for geography, history, economic policy, and development and factor endowments. Each control is motivated by a particular theory. At times the same control is claimed for different theories (with the opposite sign), underlining the rampant model uncertainty. Below we provide a brief description of the theoretical underpinnings of the various controls suggested by the previous literature. It is crucial to outline this diversity of approaches to justify the use of the model averaging methodology.

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regulations may be negatively correlated. Hence the PTA_{ijt} dummies and the error term are negatively correlated, and the PTA coefficient will tend to be underestimated.

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Table I summarizes the extent of the model uncertainty by tabulating the covariates suggested by earlier studies. It highlights the numerous attempts to identify determinants of trade flows and the associated diversity of results. The table shows how important it is to incorportate the model uncertainty that is inherent in gravity/PTA regressions as part of the empirical strategy. When the uncertainty about the true specification is not accounted for in the econometric method, the precision of estimates is inflated, since they neglect the uncertainty surrounding the true theory.

It is important to outline the theoretical backbone for each covariate included in the analysis. Without theoretical support, the results are difficult to interpret. The first set of control variables captures historical ties, such as *Common Language*, *Common Colonizer*, or *Colony*. These covariates are commonly included to capture transaction costs due to communication and/or cultural differences.[15] Common historical ties lead to similar institutions and similar levels of development, implying reliable contractual and legal standards, as well as trust in shared values. Controlling for model uncertainty addresses not only which one of these regressors (or regressor combinations) is appropriate, but also whether their inclusion is indeed approximating the true model.

Geographic factors have been introduced as proxies for transport costs (e.g. Aitken, 1973), trade-and-geography theories (e.g. Helpman and Krugman, 1985), or New Trade Theories (e.g. Rivera-Batiz and Romer, 1991). *Remoteness* (developed by Rose, 2000) is the GDP-weighted negative of distance that is often included to capture the notion that relatively remote country pairs are expected to trade more, because they have fewer options in choosing trade partners.[16] It has also been motivated as a proxy for multilateral resistance, or the average trade costs facing a country (Carrere, 2006). *Land Area* is intended to capture self-sufficiency and scale effects that are prominent in both the new trade and growth theories (e.g. Rose, 2000; Rose and van Wincoop, 2001). Scale effects are also proxies for technology or knowledge spillovers (e.g. Grossman and Helpman, 1991).

Alternative proxies in the geography category, such as *Border*, *Landlocked*, and *Island*, have previously been utilized by a variety of authors, although it is not immediately clear why adjacency should matter after having controlled for distance.[17] Perhaps variables that measure distance center-to-center introduce errors that are mitigated by the additional controls, because neighboring countries often engage in large volumes of trade. BMA addresses the uncertainty surrounding the inclusion of geography variables by indicating which covariates are relevant to explaining how PTAs influence trade patterns.

Covariates for development and factor endowments juxtapose the Heckscher–Ohlin factor endowments trade theory with Linder's (1961) hypothesis, which holds that similar countries should trade more because of their similar tastes. Davis (1995) presents an augmented Heckscher–Ohlin–Ricardo model that provides support for either theory, depending on the technological distance between the countries, and Spilimbergo and Stein (1998) examine the issue empirically. Common proxies for factor endowment differences are based on *Per Capita GDP*, *Schooling*, and *Population Density*.[18] The theoretical rationale for *Per Capita GDP* is based on the strategic trade literature (e.g. Helpman and Krugman, 1985), which predicts intra-industry trade to increase as countries become more similar in their levels of development. Furthermore, countries with higher per capita GDP are likely to have better access to less distortionary revenue sources. Hence they may experience more bilateral trade since they can afford lower tariffs.

[15] See Wei (1996); Frankel (1997); Rose (2000); Soloaga and Winters (2001); Rose and van Wincoop (2001); Frankel and Rose (2002).

[16] See Wei (1996); Rose (2000); Soloaga and Winters (2001); Baier and Bergstrand (2007).

[17] See Frankel and Romer (1999); Rose (2000); Feenstra *et al*. (2001); Rose and van Wincoop (2001); Soloaga and Winters (2001); Frankel and Rose (2002).

[18] They have been introduced by Frankel (1992, 1997), Frankel and Wei (1993), Frankel *et al*. (1995), Freund (2000), Rose and van Wincoop (2001), and Frankel and Rose (2002).

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Table I. Relationship between gravity model controls and bilateral trade in past studies

		Relationship in past studies		
		Positive	None	Negative
Trade creation 0–1 dummies	$AFTA_{ij}$	3	2	
	$ANZCERTA_{ij}$	1		
	$APEC_{ij}$	3		
	AP_{ij}	3	2	
	$CACM_{ij}$	4	2	
	$CARICOM_{ij}$			
	EEA_{ij}			
	$EFTA_{ij}$	3	5	
	EU_{ij}	9	9	
	$LAIA_{ij}$	4	2	
	$MERCOSUR_{ij}$	2	3	
	$NAFTA_{ij}$	1	3	
Trade diversion/open bloc 0–1 dummies	$AFTA_i$	2	1	1
	$ANZCERTA_i$			
	$APEC_i$			
	AP_i		1	2
	$CACM_i$		2	2
	$CARICOM_i$			
	EEA_i			
	$EFTA_i$	1	1	
	EU_i	2	1	
	$LAIA_i$		2	2
	$MERCOSUR_i$		2	2
	$NAFTA_i$	1	2	1
Core gravity	$\log(DISTANCE_{ij})$		1	23
	$\log(GDP_i\ GDP_j)$	23	2	1
	$\log(gdp_i\ gdp_j)$	9	1	2
Economic policy variables	$SACHS_i+SACHS_j$	1		
	CU_{ij}	3	1	
	$FLOAT_{ij}$	1		1
	$VOLATILITY_{ij}$	1	1	4
Dev't/factor endowment	$\text{abs}(gdp\_DIFF)$	3	1	1
	$\text{abs}(DENS\_DIFF)$	1	1	
	$\text{abs}(SCHOOL\_DIFF)$		1	
Geography	$BORDER_{ij}$	19	5	
	$REMOTE_{ij}$	4	3	
	$LANDLOCK_{ij}$	3	2	2
	$\log(AREA_i\ AREA_j)$	4		
	$ISLAND_{ij}$	3	1	1
Historical ties	$COMLANG_{ij}$	12	1	1
	$COMCOL_{ij}$	3		
	$COLONY_{ij}$	5		2

Note: Following Ghosh and Yamarik (2004), from whom parts of this table are adapted, an estimated relationship is reported positive or negative when a paper reports the coefficient significant at the 1% level. One paper may have multiple entries for the same regressor, if different regressions in the paper yield different relationships. See Tables II(a)–(c) for additional variable description.

Sources: Aitken (1973); Aitken and Lowry (1973); Baier and Bergstrand (2007); Baldwin and Taglioni (2006); Baxter and Kouparitsas (2006); Bergstrand (1985); Brada and Mendez (1988); Carrere (2006); Cheng (2005); Coe and Hoffmaister (1999); Eichengreen and Irwin (1996); Egger (2000); Egger and Pfaffermayr (2003); Feenstra *et al*. (2001); Frankel (1992); Frankel and Rose (1998); Frankel *et al*. (1995); Frankel and Wei (1993, 1996); Freund (2000); Montenegro and Soto (1996); Rose (2000); Soloaga and Winters (2001); Thursby and Thursby (1987); Wei (1996); Wei and Frankel (1998); Wei and Zhang (2006).

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Economic policy variables that are commonly included relate to trade/financial openness and exchange rate management. These are important controls as trade restrictions can explain deviations from trade patterns implied by the pure gravity equation. The Sachs and Warner (1995) *Trade Openness* variable is inserted into the gravity equation to account for trade policy effects. In addition, proxies that measure capital account openness, and financial transaction costs such as *Currency Union*, *Floating FX Rate*, and *FX Volatility* are usually included, although it is not clear what coefficient estimates are to be expected. Clark *et al*. (2004) survey the literature and highlight that just this subset of regressors alone is so deeply affected by model uncertainty that the impact of exchange rate fluctuations depends on the specific assumptions of each model.[19]

Finally we address model uncertainty in the PTA theory itself.[20] Not only do we have opposing implications suggested by different theories, but at times opposing theories have been suggested by the same author (see, for example, Krugman, 1991a,b). The theory of PTAs is based on Viner's (1950) theory of trade creation and diversion. By the 1990s, a full-scale discussion erupted regarding the drivers of trade creation and diversion. Krugman (1991a,b) examined the relative merits of PTAs in a static, monopolistically competitive framework that emphasized economic geography. His first model implied PTAs should not be welfare creating in the absence of intercontinental transport costs. At the other extreme, Krugman's second model suggested regional PTAs increase trade flows and subsequently welfare in the presence of prohibitive intercontinental transport costs.

Krugman's theories led Frankel *et al*. (1995), Frankel (1997), and Wei and Frankel (1998) to develop theories based on a continuum of transport costs. Their work characterizes trade partners as 'natural' on the basis of relatively low intercontinental transport costs and their approach implies that trade creation among 'natural' trading partners should dominate small trade diversion among remote country pairs from a welfare perspective. As trade costs fall, however, trade diversion may become larger since 'natural' trading partners overly skew their trade toward PTA partners. Frankel *et al*. (1995) suggest two hypotheses. First, the more remote trading partners are from the rest of the world, the more likely they are to form PTAs due to less potential trade diversion. This effect could be picked up by the *Remoteness* proxy. Second, the more 'natural' trading partners are, the more likely PTAs are to lead to trade creation.

Krugman's and Frankel *et al*.'s theories are based on one factor/one industry models. Deardorff and Stern (1994) note that these models preclude trade due to comparative advantage. Deardorff and Stern point out that this 'stacks the deck' against bilateralism and argue that, given differences in factor endowments, trade with a few countries suffices in order to maximize gains from trade. Thus trade diversion would be minimal. In response, Baier and Bergstrand (2004a) construct a model that builds upon Frankel *et al*. (1995) to allow for comparative advantage and scale effects. Freund (2000) argues strongly for PTA open bloc trade creation effects (even if trade creation among members is absent) since PTAs help outside exporters overcome fixed trade costs. Trade-diverting effects, instead, are highlighted by Bond and Syropoulos (1996), who indicate that the increased market power of PTAs, relative to the market power of each member taken individually, may lead to higher external tariffs.

2.4. Bayesian Model Averaging

This section briefly outlines the BMA methodology used in the estimation. We limit ourselves to discussing the properties relevant to our application. The interested reader is referred to the

[19] Authors who introduced such regressors into the gravity equation include Rose (2000), Frankel and Rose (2002), Rose and van Wincoop (2001), Glick and Rose (2002), and Tenreyro and Barro (2007).
[20] For a more detailed literature review, see Panagariya (1999, 2000).

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comprehensive tutorial by Raftery *et al*. (1997) for further discussion.[21] BMA is a natural candidate to address model uncertainty surrounding the correct controls in equations (1)–(4), since it provides probability distributions over both the model space and the parameter space. In our PTA estimation, the model space consists of all the possible subsets of candidate regressors that have been suggested by the distinct theories summarized above.

For linear regression models, the basic BMA setup can be concisely summarized as follows. Given a dependent variable, Y, a number of observations, n, and a set of candidate regressors, $X_1, X_2, \ldots, X_k$, the variable selection problem is to assess the quality of model:

$$Y = \alpha + \sum_{j=1}^{p} \beta_j X_j + \varepsilon \tag{5}$$

where $X_1, X_2, \ldots, X_p$ is a subset of $X_1, X_2, \ldots, X_k$, and β is a vector of regression coefficients to be estimated. Note that (5) is specified for linear models. Given the data, d, BMA first estimates a posterior distribution $P(\beta_r|d, M_k)$, for every candidate regressor, r, in every model M_k that includes β_r. It then combines all posterior distributions into a weighted averaged posterior distribution, $P(\beta_r|d)$, using each model's posterior probability, $P(M_k|d)$, as model weight:

$$P(\beta_r|d) = \sum_{r \,\in\, M_k} P(\beta_r|d, M_k)\ P(M_k|d) \tag{6}$$

The posterior model probability of M_k is simply the ratio of its marginal likelihood to the sum of the marginal likelihoods over all other models:

$$P(M_k|d) = \frac{l(d|M_k)}{\sum_{h=1}^{2^k} l(d|M_h)} \tag{7}$$

where posterior model probabilities are also the weights used to establish the posterior means and variances:

$$\mu \equiv E[\beta_k|d] = \sum_{k \in M}^{K} \hat{\beta}_k P(M_k|d) \tag{8}$$

$$\sigma \equiv Var[\beta_k|d] = \sum_{k \in M}^{K} (Var[\beta_k|d, M_k] + \hat{\beta}_k^2)\ P(M_k|d) - E[\beta_k|d]^2 \tag{9}$$

Summing the posterior model probabilities over all models that include a candidate regressor, we obtain the posterior inclusion probability:

$$P(\beta_k \neq 0|d) = \sum_{r \in M} P(M_k|d) \tag{10}$$

The posterior inclusion probability provides a probability statement regarding the importance of a regressor that directly addresses the researchers' prime concern: what is the probability that

[21] For recent methodological contributions to BMA see, for example, Doppelhofer and Weeks (2009), Ley and Steel (2009), and Eicher *et al*. (2010).

DOI: 10.1002/jae

the regressor has a non-zero relationship with the dependent variable? The general rule developed by Jeffries (1961) and refined by Kass and Raftery (1995) stipulates effect thresholds for posterior probability. Posterior probabilities <50% are seen as *evidence against* an effect, while the evidence for an effect is either *weak*, *positive*, *strong*, or *decisive* for posterior probabilities of 50–75%, 75–95%, 95–99%, and >99%, respectively. In our analysis, we refer to a regressor as 'effective' if its posterior inclusion probability exceeds 50%.

BMA has a number of key advantages over estimating a single model, and over extreme bound analysis. Raftery and Zheng (2003) show that BMA (a) minimizes the total error rate (sum of Type I and Type II error probabilities), (b) its point estimates and predictions minimize mean squared error (MSE), and (c) its predictive distributions have optimal predictive performance relative to other approaches. Contrary to extreme bound analysis, BMA examines the entire model space and imposes no restrictions on the model size. Ghosh and Yamarik (2004) only consider models that contain a specific number of fixed variables. In addition to these fixed regressors, a fixed number of regressors is rotated in and out of ech regression. This approach limits the model search to a fraction of the model space that is spanned by all candidate regressors. This has been shown to render extreme bound analysis excessively stringent (see Sala-i-Martin, 1997).

3. DATA

Our dataset is based on the Ghosh and Yamarik (2004) dataset to allow for a direct re-examination of their evidence using BMA as our alternative statistical methodology. The Ghosh and Yamarik dataset is based on Frankel and Rose (2002) and it includes 12 PTAs,[22] 3420 bilateral trade pairs at 5-year intervals from 1970 to 1995, and a total of 14,522 observations.[23] This dataset features average bilateral trade as the dependent variable, recorded in US dollars and deflated by the US GDP chained price index. In addition to the basic gravity and trade agreement variables, 16 control variables have been suggested by various gravity approaches discussed above.

To address refinements in the theoretical and empirical trade flow specifications suggested by the recent literature, we expand the baseline dataset in several dimensions. We extend the time horizon from 1960 to 2000 and allow for 60 additional (bilateral) trade agreements that are included in the Subramanian and Wei (2007) dataset, which features 164 importers and 177 exporters. This increases the total number of observations to 37,983.[24] We follow Subramanian and Wei (2007) and choose bilateral imports as the dependent variable; nominal imports are obtained from the IMF's *Direction of Trade Statistics*.[25] Overall our updated dataset extends the unbalanced panel of Subramanian and Wei (2007) in the following three dimensions: (a) it disaggregates the Subramanian–Wei catch-all PTA variable; (b) it allows for additional PTAs not considered in

[22] The PTAs are the European Union (EU), European Free Trade Arrangement (EFTA), European Economic Area (EEA), Central American Common Market (CACM), Caribbean Community (CARICOM), North American Free Trade Agreement (NAFTA), Latin American Integration Association (LAIA), Andean Pact (AP), Southern Cone Common Market (MERCOSUR), Association of South-East Asian Nations Free Trade Area (AFTA), Australia–New Zealand Trade Agreement (ANZCERTA), and Asian Pacific Economic Cooperation (APEC).
[23] See Ghosh and Yamarik (2004, Appendix C) for further details.
[24] With 177 countries in the IMF's *Direction of Trade Statistics*, as obtained by Subramanian and Wei (2007), potentially trading in the nine time periods from 1960 to 2000, we have $177 \times 176 \times 9 = 280,368$ potential observations. Of these, 72,211 report non-zero values. Dropping observations with import values of less than \$500,000 reduces the dataset to 52,340. Missing values for key covariates reduce the dataset by another 14,357 observations to yield our final dataset of 37,983 observations.
[25] Note that Subramanian and Wei (2007) deflate bilateral imports by the US CPI. Here we use nominal import values as they yield the same results once time fixed effects are included (see Baldwin and Taglioni, 2006).

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Table II(A). Preferential trading arrangements

Abbreviation	Name of PTA	Start	Member countries
ANZCERTA	Australia–New Zealand Closer Economic Relations Trade Agreement	1983	Australia, New Zealand
APEC	Asia Pacific Economic Community	1989	Australia, Brunei, Canada, China (1991), Chile (1994), Taiwan (1991), Hong Kong (1991), Indonesia, Japan, South Korea, Malaysia, Mexico (1993), New Zealand, Papua New Guinea (1993), Peru (1998), Philippines, Singapore, Thailand, United States, Vietnam (1998)
AP	Andean Community/Andean Pact	1969	Bolivia, Colombia, Ecuador, Peru, Venezuela (1973). Former: Chile (1969–76)
AFTA	Association of South East Asian Nations (ASEAN) Free Trade Area	1967	Brunei (1984), Cambodia (1998), Indonesia, Laos (1997), Malaysia, Myanmar (1997), Philippines, Singapore, Thailand, Vietnam (1995)
CACM	Central American Common Market	1960	Costa Rica (1963), El Salvador, Guatemala, Honduras, Nicaragua
CARICOM	Caribbean Community/Carifta	1968	Antigua and Barbuda, Bahamas (1983), Barbados, Belize (1995), Dominica (1974), Guyana (1995), Grenada (1974), Jamaica, Montserrat (1974), St Kitts and Nevis, St Lucia (1974), St Vincent and the Grenadines, Suriname (1995), Trinidad and Tobago
EEA	European Economic Area	1994	Austria, Belgium, Denmark, Finland, France, Germany, Greece, Luxembourg, Iceland, Italy, Ireland, Liechtenstein, Netherlands, Norway, Portugal, Spain, Sweden, UK
EFTA	European Free Trade Association	1960	Iceland, Liechtenstein (1991), Norway (1986), Switzerland. Former: Denmark (1960–72), UK (1960–72), Portugal (1960–85), Austria (1960–94), Sweden (1960–94), Finland (1986–94)
EU	European Union	1958	Austria (1995), Belgium, Denmark (1973), Finland (1995), France, Germany, Greece (1981), Luxembourg, Ireland (1973), Italy, Netherlands, Portugal (1986), Spain (1986), Sweden (1995), UK (1973)
LAIA/LAFTA	Latin America Integration Agreement	1960	Argentina, Bolivia (1967), Brazil, Chile, Colombia (1961) Ecuador (1961), Mexico, Paraguay, Peru, Uruguay, Venezuela (1966)
MERCOSUR	Southern Cone Common Market	1991	Argentina, Brazil, Paraguay, Uruguay
NAFTA	Canada–US Free Trade Arrangement/North America Free Trade Agreement	1988	Canada, USA, Mexico (1994)
BilateralPTA	Bilateral Preferential Trade Agreements		All bilateral agreements considered are listed in Table II(b)

Note: This table is based on Ghosh and Yamarik (2004) and includes corrections to some of the original PTA coding as follows. ASEAN, which is no free trade area, was changed to AFTA, with AFTA membership starting in 1992 instead of 1980. For the Andean Pact, Chile had to be excluded post 1976, when it left the AP. Finally, CARICOM membership for Guyana is corrected to start in 1973 (instead of 1995). The corrections do not alter the qualitative results.

Subramanian and Wei (2007);[26] and (c) it incorporates a comprehensive list of additional controls suggested by the previous literature. Detailed descriptions of the PTAs and the other control variables included in the extended dataset can be found in Tables II(a)–(c).

[26] This extension adds the European Free Trade Agreement (EFTA), the European Economic Area (EEA), the Andean Pact (AP), the Latin American Integration Association and the Asia Pacific Economic Community (APEC) to the analysis.

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Table II(B). Bilateral Preferential Trade Agreements considered in BilateralPTA

USA–Israel	Slovak Republic–Turkey
Turkey–Slovenia	Papua New Guinea–Australia Trade & Commercial Relations Agreement (PATCRA)
EC–Slovenia	EC–Tunisia
EC–Lithuania	Estonia–Turkey
EC–Estonia	Slovenia–Israel
EC–Latvia	Poland–Israel
Chile–Mexico	Estonia–Faroe Islands
Mexico–Israel	Czech Republic–Estonia
Georgia–Armenia	Slovak Republic–Estonia
Georgia–Azerbaijan	Lithuania–Turkey
Georgia–Kazakhstan	Israel–Turkey
Georgia–Turkmenistan	Romania–Turkey
Georgia–Ukraine	Hungary–Turkey
Latvia–Turkey	Czech Republic–Israel
Turkey–former Yugoslav Rep. of Macedonia	Slovak Republic–Israel
EC–South Africa	Slovenia–Croatia
EC–Morocco	Hungary–Israel
EC–Israel	CEFTA accession of Romania
EC–Mexico	CEFTA accession of Slovenia
Estonia–Ukraine	Poland–Lithuania
Poland–Turkey	Slovak Republic–Latvia
EFTA–Morocco	Slovak Republic–Lithuania
Bulgaria–former Yugoslav Rep. of Macedonia	Canada–Chile
Hungary–Latvia	Czech Republic–Latvia
Hungary–Lithuania	Czech Republic–Lithuania
Poland–Latvia	Slovenia–Estonia
Poland–Faeroe Islands	Slovenia–Lithuania
Kyrgyz Republic–Moldova	EC–Faeroe Islands
Kyrgyz Republic–Ukraine	Canada–Israel
Kyrgyz Republic–Uzbekistan	EFTA–Estonia
Bulgaria–Turkey	EFTA–Latvia
Czech Republic–Turkey	EFTA–Lithuania
EAEC	EC–Turkey
CEFTA accession of Bulgaria	

Source: Subramanian and Wei (2007).

The Subramanian and Wei (2007) data are in turn based on Rose (2004) and Glick and Rose (2002) work on the determinants of trade flows; we maintain their convention of including only those of the roughly 280,000 observations whose trade values exceed \$500,000. There exists, however, an important literature that seeks to understand the true nature of the data when zero trade flows are observed. Zero trade values may also be due to a rounding error or missing observations, and in a log-linear gravity equation zeros are automatically excluded. If a zero trade value were to be an accurate representation of two countries' goods trade, the observation should not be excluded, since it holds information and its absence may induce selection bias.

Santos Silva and Tenreyro (2006) suggest the Poisson pseudo-maximum likelihood (PPML) estimator to appropriately address the issue of zero trade values. This method has been shown to reduce estimates by as much as 40%. Martin and Pham (2008) suggest that the PPML estimator is efficient in addressing heterogeneity, but still biased in the presence of zero trade values. Based on the results of their simulations, they instead recommend a Heckman maximum likelihood approach to control for selection bias. Below we follow Rose's and Subramanian and Wei's OLS approach not only to maintain comparability with their results, but also because neither a PPML nor a Heckman estimator has been developed to date for BMA application.

DOI: 10.1002/jae

Table II(C). Description of non-PTA related variables

Variable	Description	Source
$\log(IMPORTS_{ijt})$	Natural log of bilateral imports (current US dollars)	IMF Direction of Trade Statistics
$\log(DISTANCE_{ij})$	Natural log of the bilateral distance	Subramanian and Wei (2007)
$\log(GDP_{it}\ GDP_{jt})$	Natural log of the product of nominal GDP	Penn World Tables
$\log(gdp_{it}\ gdp_{jt})$	Natural log of the product of real GDP per capita	Penn World Tables
$SACHS_{it}+SACHS_{jt}$	The sum of the Sachs–Warner index of an open trade policy	Sachs and Warner (1995); Wacziarg and Welch (2003)
CU_{ijt}	Dummy (1 if the two share a common currency)	Subramanian and Wei (2007)
$FLOAT_{ijt}$	Number of countries with a floating exchange rate (0, 1, 2)	IMF Annual Report on Exchange Rate Arrangements and Restrictions
$VOLATILITY_{ijt}$	Standard deviation of the first difference in the bilateral exchange rate during the previous 3 years	IMF International Financial Statistics
abs(*gdp_DIFF*)	Absolute log difference of real GDP per capita	Penn World Tables
abs(*DENS_DIFF*)	Absolute log difference in population density	CIA World Fact Book
abs(*SCHOOL_DIFF*)	Absolute log difference in the average years of secondary schooling in the 25+ population	Barro and Lee (2001)
$BORDER_{ij}$	Dummy (1 if the two share a common land border and 0 otherwise)	Subramanian and Wei (2007)
$REMOTE_{ijt}$	Natural log of the product of the average distance (weighted by relative GDP) of each country from all trading partners[a]	CIA World Fact Book and Penn World Tables
$LANDLOCK_{ij}$	Number of landlocked countries (0, 1, 2)	Subramanian and Wei (2007)
$\log(AREA_i\ AREA_j)$	Natural log of the product of the surface area of the two countries	CIA World Fact Book
$ISLAND_{ij}$	Number of island countries (0, 1, 2)	Subramanian and Wei (2007)
$COMLANG_{ij}$	Dummy (1 if the two share a common language and 0 otherwise)	Subramanian and Wei (2007)
$COMCOL_{ij}$	Dummy (1 if the two share a common colonizer and 0 otherwise)	Subramanian and Wei (2007)
$COLONY_{ij}$	Dummy (1 if one was a former colony of the other and 0 otherwise)	Subramanian and Wei (2007)

[a] The remoteness variable is constructed as in Rose (2000). The use of the Ghosh and Yamarik (2004) remoteness variable does not alter the qualitative results.

4. RESULTS

4.1. PTA Trade Creation: Differences Due to Methodologies

Ghosh and Yamarik (2004) embarked on the most comprehensive robustness test of PTAs to date. They considered not just a subset but all major PTAs and employed extreme bound analysis to explore the model space far beyond what ordinary robustness exercises can hope to represent. Our first objective is to replicate Ghosh and Yamarik's (2004) results using BMA methodology. Table III reports results for two specifications. Specification 1 employs BMA on the exact same data and regression equation in Ghosh and Yamarik (2004, equation 1). Specification 2 differs from Specification 1 only in that it uses our new updated dataset based on Subramanian and Wei (2007).

Table III highlights that our key result is independent of the choice of PTA dataset that is used. Once model uncertainty is addressed in a principled fashion using BMA, Ghosh and Yamarik's

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Table III. PTA trade creation and trade diversion

		Specification 1 Time fixed effects Original Ghosh and Yamarik (2004) specification and data			Specification 2 Time fixed effects Ghosh and Yamarik (2004) specification, updated Subramanian/Wei data		
		$p \neq 0$	μ	σ	$p \neq 0$	μ	σ
Trade creation	$AFTA_{ijt}$	0	−0.22	0.54	1	0.36	0.35
	$ANZCERTA_{ijt}$	1	0.89	0.96	1	0.88	0.62
	$APEC_{ijt}$	100	1.48***	0.15	100	1.71***	0.09
	AP_{ijt}	1	−0.05	0.27	99	0.67***	0.15
	$CACM_{ijt}$	100	2.25***	0.23	100	2.30***	0.15
	$CARICOM_{ijt}$	100	2.08***	0.41	100	2.83***	0.30
	EEA_{ijt}	1	0.26	0.19	2	0.22	0.15
	$EFTA_{ijt}$	0	0.02	0.26	100	0.67***	0.13
	EU_{ijt}	0	0.03	0.14	100	0.51***	0.09
	$LAIA_{ijt}$	91	0.46***	0.13	1	−0.05	0.08
	$MERCOSUR_{ijt}$	12	1.66	0.7	14	0.96	0.36
	$NAFTA_{ijt}$	1	−0.89	0.84	0	0.20	0.47
	$BILATERAL_{ijt}$	n.a	n.a	n.a	1	0.13	0.13
Trade diversion, open bloc	$AFTA_{it}$	3	0.17	0.11	100	0.41***	0.06
	$ANZCERTA_{it}$	100	−0.47***	0.1	100	−0.81***	0.06
	$APEC_{it}$	100	0.55***	0.06	100	0.48***	0.04
	AP_{it}	52	−0.19*	0.06	2	0.07	0.04
	$CACM_{it}$	85	−0.18**	0.05	100	−0.17***	0.03
	$CARICOM_{it}$	100	−0.74***	0.07	100	−0.58***	0.05
	EEA_{it}	0	0.01	0.08	92	−0.17**	0.04
	$EFTA_{it}$	100	0.35***	0.05	100	0.37***	0.03
	EU_{it}	100	0.56***	0.04	100	0.65***	0.03
	$LAIA_{it}$	100	−0.40***	0.07	100	−0.52***	0.03
	$MERCOSUR_{it}$	79	0.42**	0.12	0	−0.04	0.06
	$NAFTA_{it}$	100	−0.63***	0.1	4	0.13	0.06
	$BILATERAL_{it}$	n.a	n.a	n.a	100	−0.27***	0.04
Core gravity	$\log(GDP_{it}\ GDP_{jt})$	100	0.88***	0.01	100	0.94***	0.01
	$\log(DISTANCE_{ij})$	100	−1.19***	0.02	100	−1.08***	0.02
	$\log(gdp_{it}\ gdp_{jt})$	100	0.55***	0.02	100	0.28***	0.01
Economic policy	$SACHS_{it}+SACHS_{jt}$	100	0.35***	0.03	100	0.22***	0.02
	$VOLATILITY_{ijt}$	25	0.006	0.002	0	−0.0003	0.00
	$FLOAT_{ijt}$	0	−0.01	0.02	100	0.09***	0.02
	CU_{ijt}	100	1.40***	0.29	100	1.22***	0.10
Development, factor endowments	abs(*SCHOOL_DIFF*)	1	0.02	0.02	14	0.04	0.02
	abs(*DENS_DIFF*)	100	0.23***	0.01	100	0.13***	0.01
	abs(*gdp_DIFF*)	100	0.18***	0.02	100	0.08***	0.01
Geography	$BORDER_{ij}$	100	0.53***	0.1	100	0.40***	0.06
	$ISLAND_{ij}$	2	−0.05	0.03	100	−0.22***	0.03
	$LANDLOCK_{ij}$	100	−0.42***	0.04	100	−0.26***	0.02
	$\log(AREA_i\ AREA_j)$	92	−0.03**	0.01	100	−0.08***	0.01
	$REMOTE_{ijt}$	100	342***	39.79	100	1.31***	0.04
History	$COLONY_{ij}$	100	1.44***	0.12	100	1.12***	0.06
	$COMCOL_{ij}$	100	0.77***	0.07	100	0.55***	0.04
	$COMLANG_{ij}$	100	0.47***	0.05	100	0.28***	0.02

Note: Fixed effect coefficients are omitted. Asterisks represent *weak, positive*, and *decisive* evidence for an effect of the regressor, corresponding to posterior inclusion probabilities of * 50–75%, ** 75–99%, and *** >99%, respectively (see Jefferies, 1961; Kass and Raftery, 1995). $p \neq 0$ is the inclusion probability, μ is the posterior mean, and σ is the posterior standard deviation.

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(2004) own econometric specification produces a host of PTA effects that range from trade creating to open bloc and even trade diverting. We obtain effective coefficients (indicated with asterisks) whose signs and magnitudes are similar to those commonly reported in the previous literature. BMA thus provides evidence that the model space spanned by 'free and doubtful variables' through extreme bound analysis was too restrictive. The models flagged out by extreme bound analysis did not contain those that feature the highest posterior probabilities, and the heuristic model weighting assigned by extreme bound analysis generated excessively conservative results that indicated no PTA effects. The expanded model space, combined with the principled weighting of effective models, generate BMA's superior predictive performance.

Of the 13 major trade agreements, eight are found to be either trade creating and/or to exhibit open bloc effects in Specification 1. All Western Hemisphere PTAs are identified as trade diverting in the original Ghosh and Yamarik dataset (Specification 1). The additional years and controls for bilateral agreements in our updated dataset (Specification 2) increase the precision of the estimates, but our key insights remain the same. Specification 2 produces four additional trade creation effects (for key PTAs such as the EFTA, AFTA and the EU), and erases the odd implication of NAFTA trade diversion that was reported by Specification 1. These changes are most certainly due to the extension of the time horizon from the mid 1990s to 2000. In summary, the BMA results robustly link PTAs to changes in trade flows, although the effects vary across PTAs.

A substantial literature addresses the possibility of PTA coefficient bias due to omitted variables or inaccurate model specification. We extend our analysis to incorporate the insights of this recent literature to examine the robustness of our results. The scale of some PTA coefficients in Table III is certainly suspicious if not implausible. Coefficients that exceed unity imply that a PTA increased trade more than twofold (since the regression is in logs). Such aberrant magnitudes have previously been noted and questioned in the literature (e.g. Frankel, 1992, 1997; Frankel and Wei, 1993; Frankel *et al*., 1995). We take up the issue of omitted variable bias in the following section.

4.2. Multilateral Resistance

Ghosh and Yamarik (2004) and our Specifications 1 and 2 (Table III) include time fixed effects, but the recent PTA literature suggests the inclusion of additional fixed effects to account for multilateral trade costs. Wei (1996), Deardorff (1998), Anderson and van Wincoop (2003), and Subramanian and Wei (2007) emphasized that the standard gravity model is subject to misspecification bias if multilateral trade costs are ignored. The crucial insight is that bilateral trade is influenced by the average multilateral trade cost faced by a country in any given period. Anderson and van Wincoop (2003) suggest that, empirically, the inclusion of country fixed effects captures multilateral resistance. Since bilateral trade between any two countries depends on the multilateral resistance of *both* importers and exporters, the Anderson and van Wincoop (2003) model requires fixed effects for both countries involved in any bilateral trading relationship.[27] In a panel, these importer and exporter fixed effects must be time varying, which allows the PTA dummies in equation (2) to identify net trade creation. This fixed-effect approach has been popularized by Subramanian and Wei (2007) in their analysis of WTO trade effects (although these authors do not break out the effects of individual PTAs).

Specification 4 reports results that control for multilateral resistance. Specification 3 replicates the results in Specification 2, without separate trade diversion/open bloc effects. As expected, the results for most trade agreements are very similar to the sum of trade creation and diversion in Specification 2. For example, the Central American Common Market (CACM) featured a

[27] Helpman *et al*. (2008) suggest an alternative rationale for importer and exporter fixed effects based on firm heterogeneity.

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coefficient for trade creation of 2.3 in Specification 2, and a trade diversion effect with the rest of the world of 0.17. The combined net trade creation for PTA members is then an implied 2.47, which is closely matched by the estimate of 2.45 in Specification 3.

More importantly, however, Table IV shows that even after controlling for multilateral resistance the fundamental result of our analysis remains unchanged: PTAs have a strong impact on bilateral trade. Of the 13 major PTAs covered, 10 PTAs exhibit an effect on bilateral trade, only one of which is negative. This implies that controlling for multilateral resistance identifies four additional PTAs with significant positive impacts on bilateral trade flows.

The one surprise in Specification 4 is the implied negative net trade creation for the EU. The attractiveness of the EU market with its large size and strong harmonization likely exerts a significant pull on non-EU exporters, resulting in the large open bloc effects estimated at about 0.6 in Specifications 1 and 2. The drag of open bloc effects on net trade creation by itself thus explains roughly half of the negative coefficient estimate. In addition, it is well known that the gravity equation overpredicts EU trade when estimated on a global sample. Given their close proximity and other bilateral characteristics, EU countries undertrade relative to the globally based prediction, resulting in a negative EU coefficient. This may be related to the gravity equation's inability to proxy firms' fixed costs in establishing trade relations (e.g. Freund, 2000). Empirically, Aitken (1973) and Rose (2004) find similarly negative EU results. Inclusion of country-pair fixed effects is commonly suggested to control for such time-invariant bilateral heterogeneity. It represents the main alternative to time-varying importer/exporter fixed effects for our robustness analysis. By examining EEA, EFTA and EU effects across alternative fixed-effects specifications, Baier *et al.* (2008) also find a similar instability and turn to country-pair fixed effects to obtain robust effects.

4.3. Unobserved Heterogeneity

To capture unobserved time-invariant heterogeneity among trade partners, we re-estimate Specification 4, accounting for country-pair fixed effects. This specification does not address multilateral trade costs as comprehensively as suggested by Anderson and van Wincoop (2003), especially if they exhibit large fluctuations over time. However, Rose (2004) makes the point that country-pair fixed effects constitute a valid proxy for average multilateral resistance exhibited in country pairs. Hummels and Levinsohn (1995) first introduced country-pair fixed effects to better distinguish between factor endowments and market structure as trade flow drivers. Egger and Pfaffermayr (2003) advocate country-pair fixed effects to account for heterogeneity induced by time-invariant factors (e.g., geography, history, policy, and culture) that are only partially accounted for by the explanatory variables or completely unobserved. Glick and Rose (2002) use the same specification as Egger and Pfaffermayr (2003), but motivate country-pair fixed effects as proxies for trade resistance. Here we employ it as a robustness test of the estimated parameter magnitudes for specific PTAs, such as the EU.

Note that the introduction of country-pair fixed effects removes the cross-sectional information so that Specification 5 relies *only* on the time series information contained in the data. Specification 5, therefore, expresses only PTA effects directly caused by PTA accession or exit. Nevertheless, our central result remains robust: PTAs exert a significant effect on trade flows. The rewarding aspect of the country-pair analysis is that BMA confirms the hypothesis that the gravity model overpredicts intra-European trade flows only when pair-specific heterogeneity is ignored. Once these effects are accounted for, EU trade creation is indeed positive. On the other hand, some effects that seemed unreasonably large before are now significantly reduced. ANZCERTA, AP, EEA and MERCOSUR lose their influence on net trade flows, which indicates considerable unobserved bilateral heterogeneity members of these PTAs. With the exception of the Latin

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Table IV. PTA net trade creation controlling for multilateral resistance and bilateral heterogeneity

	Specification 3			Specification 4			Specification 5			Specification 6		
Subramanian/ Wei data		Yes			Yes			Yes			Yes	
Time fixed effects		Yes			Yes			Yes			Yes	
Imp. exp. fixed effects		No			Yes			No			Yes	
Country-pair fixed effects		No			No			Yes			Yes	
Accession dynamics		No			No			No			No	
	$p \neq 0$	μ	σ	$p \neq 0$	μ	σ	$p \neq 0$	μ	σ	$p \neq 0$	μ	σ
$AFTA_{ijt}$	1	0.46	0.35	1	−0.36	0.32	1	0.27	0.24	2	−0.38	0.29
$ANZCERTA_{ijt}$	19	1.73	0.63	94	2.15**	0.56	0	−0.02	0.51	3	0.89	0.60
$APEC_{ijt}$	100	1.39***	0.08	100	0.62***	0.09	100	0.52***	0.06	1	0.09	0.08
AP_{ijt}	88	0.63**	0.16	93	0.59**	0.15	6	0.44	0.19	99	0.80**	0.24
$CACM_{ijt}$	100	2.45***	0.15	100	2.34***	0.14	100	2.19***	0.26	100	1.59***	0.37
$CARICOM_{ijt}$	100	2.89***	0.31	100	4.27***	0.28	63	1.45*	0.43	92	1.59**	0.53
EEA_{ijt}	10	0.35	0.13	99	0.49**	0.12	25	−0.24	0.08	100	0.49***	0.11
$EFTA_{ijt}$	1	0.15	0.12	2	−0.17	0.11	5	0.26	0.12	1	−0.08	0.15
EU_{ijt}	19	−0.29	0.13	100	−1.18***	0.10	100	0.41***	0.10	21	0.25	0.12
$LAIA_{ijt}$	100	0.40***	0.09	100	0.93***	0.08	100	1.68***	0.19	100	1.18***	0.27
$MERCOSUR_{ijt}$	6	0.79	0.37	80	1.19**	0.34	1	0.38	0.26	1	0.35	0.31
$NAFTA_{ijt}$	1	−0.25	0.48	0	0.09	0.43	1	0.48	0.34	2	0.53	0.40
$BILATERAL_{ijt}$	12	0.34	0.13	53	0.41*	0.13	14	0.24	0.09	0	0.01	0.12
$\log(GDP_{it}\ GDP_{jt})$	100	1.02***	0.01				15	0.01	0.00			
$\log(DISTANCE_{ij})$	100	−1.09***	0.02	100	−1.18***	0.02						
$\log(gdp_{it}\ gdp_{jt})$	100	0.18***	0.01				100	1.06***	0.02			
$SACHS_{ijt}$	100	0.34***	0.02				100	0.20***	0.02			
$VOLATILITY_{ijt}$	1	0	0.00	1	0.00	0.01	100	−0.01***	0.00	1	0.00	0.01
$FLOAT_{ijt}$	3	0.03	0.02				100	−0.06***	0.01			
CU_{ijt}	100	1.26***	0.10	100	1.15***	0.10	100	0.66***	0.13	6	0.28	0.16
$\lvert SCHOOL\_DIFF \rvert$	100	0.08***	0.01	23	0.05**	0.02	100	0.23***	0.02	20	0.07	0.03
$\lvert DENS\_DIFF \rvert$	100	0.12***	0.01	100	0.12***	0.01	28	0.11	0.04	3	0.08	0.05
$\lvert gdp\_DIFF \rvert$	100	0.09***	0.01	6	−0.04	0.02	100	−0.19***	0.03	100	−0.31***	0.04
$BORDER_{ij}$	100	0.42***	0.06	98	0.23**	0.05						
$ISLAND_{ij}$	100	−0.26***	0.02									
$LANDLOCK_{ij}$	100	−0.20***	0.02									
$\log(AREA_i AREA_j)$	100	−0.13***	0.00									
$REMOTE_{ijt}$	100	0.81***	0.04				0	−0.01	0.02			
$COLONY_{ij}$	100	1.19***	0.06	100	1.13***	0.06						
$COMCOL_{ij}$	100	0.27***	0.02	100	0.35***	0.04						
$COMLANG_{ij}$	100	0.59***	0.04	100	0.30***	0.03						

Note: Fixed-effect coefficients are omitted. Asterisks represent *weak, positive*, and *decisive* evidence for an effect of the regressor, corresponding to posterior inclusion probabilities of * 50–75%, ** 75–99, and *** >99%, respectively (see Jefferies, 1961; Kass and Raftery, 1995). $p \neq 0$ is the inclusion probability, μ is the posterior mean, and σ is the posterior standard deviation.

American Integration Association (LAIA), magnitudes of significant PTA impacts are uniformly smaller when we explore only the specific effect of entering and exiting a trade agreement.

4.4. A Comprehensive Approach

The previous sections illustrated how each individual fixed-effect approach influences PTA estimates. In this section we present results from our most comprehensive approach, which controls for both unobserved heterogeneity and multilateral resistance simultaneously. The comprehensive

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approach adds a large number of fixed-effects controls to the regression, and is identical to the Baier and Bergstrand (2007) methodology. However, their focus was on the *average* PTA effect, while the motivation for this paper was to show the heterogeneity of trade effects across individual PTAs and to resolve model uncertainty. As outlined above, the comprehensive approach is also best suited to control for the various biases may be contained in a gravity equation, especially endogeneity bias.

Specification 6 in Table IV presents new results and presents a number of additional insights. Even after accounting for the large number of fixed effects, and after accounting for model uncertainty, a series of PTAs show strong effects on trade flows. CACM, CARICOM, EEA and LAIA all exhibit high inclusion probabilities and positive trade effects. The EU which oscillated from negative to positive coefficients is now economically significant but only marginally statistically significant. Note, however, that the EEA picks up important recent trade effects among a large number of EU members. We also find a dramatic reduction in predicted trade flows due to a PTA among the APEC countries. This is comforting since APEC did not institute actual tariff reductions, and it has been well known that the gravity models must have attributed some of the bilateral or individual country effects to the creation of APEC. Once we control for these effects, and for the potential endogenous selection of fast trade-growing countries into AFTA, we find that the actual affect of APEC is nil.

Among the non-PTA variables, only the difference in GDP remains significant. The coefficient indicates that countries with similar GDP generate larger trade volumes, which supports intra-industry trade theories rather than Heckscher–Ohlin. The variation of the results across different fixed effects raises the general question of integration dynamics. Are average estimates over the life of PTA membership appropriate, or can we observe accession dynamics where static effects (before or at accession) differ fundamentally from subsequent dynamic changes in trade flows? If we are interested in the specific effects of individual trade flows, there may well be accession dynamics in play that suggest that the simple averaging of effects over time may be misleading. We examine this hypothesis in the following section.

4.5. Accession Dynamics

Further investigation of accession dynamics may also yield benefits beyond the reconciliation of remaining differences between Specifications 4 and 5. Namely, accession dynamics provide insights whether gains from trade tend to be static, as advocated by neoclassical trade theory, or dynamic (e.g. Young 1991). Indeed the gain might even commence *before* the PTA accession. Hence we recode the PTA dummy into three separate effects. If accession occurs at time t, an *accession* dummy captures trade creation when the country joined a PTA, a *pre-accession* dummy captures the 5 years prior to joining a PTA ($t - 1$ in our notation), and a *post-accession* dummy captures the 5 years following accession to the end of the sample, $(t + 1, n)$, where n indicates either the year 2000 or the year a country exited the PTA.

Results that include accession dynamics are presented in Table V, where we present specifications that control for multilateral resistance (Specification 4a), unobserved bilateral heterogeneity (Specification 5a), and for both of the former (Specification 6a). For expositional purposes, Table V does not report non-PTA regressors that were included in the analysis to save space. The posterior estimates and inclusion probabilities are very similar to the corresponding specifications without accession dynamics. Table V also includes the average PTA effects (t, n) established in Specifications 4–6 to allow for quick comparisons between average effects and accession dynamics for each PTA.

The accession dynamics highlight the timing of the trade gains for each PTA. In general, the PTAs' effects on trade materialize in the accession and post-accession phases with the appropriate

Table V. PTA accession dynamics

		Specification 4a			Specification 5a			Specification 6a		
Subramanian /Wei data		Yes			Yes			Yes		
Time fixed effects		Yes			Yes			Yes		
Imp. exp. fixed effects		Yes			No			Yes		
Country-pair fixed effects		No			Yes			Yes		
Accession dynamics		Yes			Yes			Yes		
		$p \neq 0$	μ	σ	$p \neq 0$	μ	σ	$p \neq 0$	μ	σ
$AFTA_{ijt}$	Average (t, n)	1	−0.36	0.32	1	0.27	0.24	2	−0.38	0.29
	Pre-accession $(t-1)$	0	−0.26	0.34	1	0.00	0.07	0	−0.36	0.32
	Accession (t)	0	−0.43	0.41	0	0.00	0.02	0	−0.54	0.36
	Post-accession $(t+1, n)$	0	0.00	0.00	1	0.01	0.07	0	0.00	0.00
$ANZCERTA_{ijt}$	Average (t, n)	94	2.15**	0.56	0	−0.02	0.51	3	0.89	0.60
	Pre-accession $(t-1)$	2	2.17	1.07	1	0.00	0.03	0	0.00	0.00
	Accession (t)	1	1.74	1.07	0	0.00	0.05	0	0.00	0.00
	Post-accession $(t+1, n)$	66	2.12*	0.62	0	0.00	0.04	0	0.92	0.63
$APEC_{ijt}$	Average (t, n)	100	0.62***	0.09	100	0.52***	0.06	1	0.09	0.08
	Pre-accession $(t-1)$	100	0.74***	0.12	100	0.54***	0.08	10	0.23	0.11
	Accession (t)	100	0.75***	0.12	100	0.64***	0.08	21	0.24	0.11
	Post-accession $(t+1, n)$	100	0.57***	0.11	100	0.66***	0.08	0	−0.08	0.10
AP_{ijt}	Average (t, n)	93	0.59**	0.15	6	0.44	0.19	99	0.80**	0.24
	Pre-accession $(t-1)$	2	−0.65	0.32	0	0.00	0.02	0	0.00	0.00
	Accession (t)	0	0.00	0.00	2	−0.01	0.07	0	0.00	0.00
	Post-accession $(t+1, n)$	98	0.67**	0.16	4	0.01	0.08	87	0.68**	0.24
$CACM_{ijt}$	Average (t, n)	100	2.34***	0.14	100	2.19***	0.26	100	1.59***	0.37
	Pre-accession $(t-1)$	2	−0.79	0.42	2	−0.02	0.21	88	−1.6**	0.38
	Accession (t)	2	0.79	0.39	100	1.85***	0.35	5	0.22	0.95
	Post-accession $(t+1, n)$	100	2.52***	0.15	100	2.24***	0.27	15	1.14	0.43
$CARICOM_{ijt}$	Average (t, n)	100	4.27***	0.28	63	1.45*	0.43	92	1.59**	0.53
	Pre-accession $(t-1)$	0	0.00	0.00	2	−0.02	0.21	0	−0.86	0.99
	Accession (t)	100	4.13***	0.42	2	0.02	0.17	4	1.43	0.59
	Post-accession $(t+1, n)$	100	4.34***	0.35	2	0.02	0.14	7	1.32	0.69
EEA_{ijt}	average (t, n)	99	0.49**	0.12	25	−0.24	0.08	100	0.49***	0.11
	Pre-accession $(t-1)$	1	0.20	0.14	92	0.34**	0.13	5	0.24	0.12
	Accession (t)	1	0.26	0.14	11	0.03	0.08	42	0.29	0.12
	Post-accession $(t+1, n)$	100	0.74***	0.14	30	−0.10	0.16	100	0.69***	0.14
$EFTA_{ijt}$	Average (t, n)	2	−0.17	0.11	5	0.26	0.12	1	−0.08	0.15
	Pre-accession $(t-1)$	3	−0.64	0.30	1	0.00	0.03	0	0.00	0.00
	Accession (t)	0	0.00	0.00	1	0.00	0.01	0	−0.28	0.21
	Post-accession $(t+1, n)$	1	−0.17	0.12	30	−0.10	0.16	0	0.00	0.00

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Table V. (*Continued*)

		Specification 4a			Specification 5a			Specification 6a		
Subramanian /Wei data			Yes			Yes			Yes	
Time fixed effects			Yes			Yes			Yes	
Imp. exp. fixed effects			Yes			No			Yes	
Country-pair fixed effects			No			Yes			Yes	
Accession dynamics			Yes			Yes			Yes	
		$p \neq 0$	μ	σ	$p \neq 0$	μ	σ	$p \neq 0$	μ	σ
EU_{ijt}	average (t, n)	100	−1.18***	0.10	100	0.41***	0.10	21	0.25	0.12
	Pre-accession $(t-1)$	100	−0.97***	0.14	3	0.01	0.05	0	−0.17	0.12
	Accession (t)	100	−0.81***	0.13	38	0.12	0.17	2	0.26	0.12
	Post-accession $(t+1, n)$	100	−1.32***	0.10	95	0.37**	0.15	15	0.27	0.12
$LAIA_{ijt}$	Average (t, n)	100	0.93***	0.08	100	1.68***	0.19	100	1.18***	0.27
	Pre-accession $(t-1)$	46	−0.97	0.30	2	0.00	0.09	0	0.21	0.73
	Accession (t)	0	−0.20	0.24	42	0.31	0.40	29	0.70	0.30
	Post-accession $(t+1, n)$	100	1.03***	0.08	100	1.62***	0.28	100	1.20***	0.31
$MERCOSUR_{ijt}$	Average (t, n)	80	1.19**	0.34	1	0.38	0.26	1	0.35	0.31
	Pre-accession $(t-1)$	2	0.88	0.46	0	0.00	0.02	0	0.00	0.00
	Accession (t)	2	0.93	0.46	1	0.00	0.03	0	0.00	0.00
	Post-accession $(t+1, n)$	16	1.26	0.46	1	0.00	0.06	0	0.47	0.40
$NAFTA_{ijt}$	Average (t, n)	0	0.09	0.43	1	0.48	0.34	2	0.53	0.40
	Pre-accession $(t-1)$	0	0.00	0.00	0	0.00	0.03	0	0.00	0.00
	Accession (t)	0	0.00	0.00	1	0.00	0.04	0	0.00	0.00
	Post-accession $(t+1, n)$	0	0.00	0.00	1	0.00	0.05	0	0.60	0.49
$BILATERAL_{ijt}$	Average (t, n)	53	0.41*	0.13	14	0.24	0.09	0	0.01	0.12
	Pre-accession $(t-1)$	0	0.00	0.00	3	0.01	0.04	97	−0.4**	0.13
	Accession (t)	3	0.28	0.13	8	0.02	0.07	0	0.00	0.00
	Post-accession $(t+1, n)$	99	1.94***	0.44	0	0.00	0.03	0	0.42	0.50

Note: Fixed-effect coefficients are omitted. Asterisks represent *weak, positive*, and *decisive* evidence for an effect of the regressor, corresponding to posterior inclusion probabilities of * 50–75%, ** 75–99, and *** >99%, respectively (see Jefferies, 1961; Kass and Raftery, 1995). $p \neq 0$ is the inclusion probability, μ is the posterior mean, and σ is the posterior standard deviation.

magnitudes. The accession results also show a high coincidence between average effects and dynamic effects; namely only PTAs that produce average effects in Table IV also produce dynamic PTA effects in Table V. There are two interesting exceptions to this rule. In specification 6, where we control for both country-pair and time-varying importer/exporter time fixed effects, we find that there is a gradual onset of trade creation for those PTAs with overall net trade creation. The Latin America Integration Association (LAIA) and the European Economic Area (EEA) illustrate that net trade creation first becomes notable in the accession period and fortifies thereafter. The patterns in the Central American Common Market (CACM) and bilateral trade agreements differ slightly.

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These PTAs show negative net trade creation pre-accession, which is remedied by accession to the PTA. PTA accession thus created trade, correcting for members' previously observed undertrade relative to the gravity prediction.

4.6. Other Determinants of Trade Flows

So far we have discussed only the impact of PTAs on trade flows. However, the BMA exercise holds important additional information regarding other determinants of trade flows. The geography and history controls are highly significant in Specifications 1 and 2 (in agreement with the previous literature). Althought the magnitues of their effects are reduced by the fixed effects, they generally remain significant.

BMA identifies trade openness as a key variable in all specifications, which is not surprising since we are attempting to explain trade flows. More interesting is that a number of variables related to exchange rate policy are not significant unless we control for bilateral unobservables. The currency union variable, on the other hand, shows a strong effect independent of dataset or empirical specification. Additional variables that might influence trade flows are factor endowments. Here BMA allows us to examine the competing hypotheses that trade flows are either driven by differences in endowments (Heckscher–Ohlin) or by similarities (Lindner). In Specifications 1 and 2, the Heckscher–Ohlin factor endowment theory finds strong support, as differences in per capita GDPs and population densities are strongly associated with greater trade flows. The endowment effect vanishes, however, when we consider multilateral resistance. Effects of population density disappear once we account for bilateral heterogeneity. Finally, the BMA methodology shows that differences in schooling increase bilateral trade flows when we control for either unobserved heterogeneity or multilateral resistance.

5. CONCLUSION

The literature on preferential trade agreements (PTAs) features an unusual diversity of theoretical and empirical approaches. In this paper we incorporate model uncertainty into the empirical strategy by applying Bayesian model averaging (BMA). To date, the most extensive robustness analysis by Ghosh and Yamarik (2004) used extreme bound analysis and found evidence *against any* effects of PTAs at the extreme bounds. In contrast, applying BMA to Ghosh and Yamark's original dataset we find that PTA trade creation is strong. In addition, the BMA approach produces coefficient estimates that resolve a number of empirical puzzles.

We confirm strong PTA effects not only with Ghosh and Yamarik's original dataset, but also with an updated dataset that includes additional years and PTAs. Our results are robust to the inclusion of multilateral resistance, accession dynamics, and unobserved bilateral heterogeneity. Overall, the observed PTA effects reflect the diversity of PTAs and the degree of tariff reductions they encompass. BMA allows us to also account for model uncertainty in the set of additional control variables usually featured in PTA regressions. Our approach highlights the importance of including all controls for policy, development, factor endowments, geography, and history that have been suggested by the previous literature. Among these regressors, the only ones that receive mixed evidence are those related to exchange rate fluctuations.

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