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In search of WTO trade effects: Preferential trade agreements promote trade strongly, but unevenly[☆]

Theo S. Eicher [a,*], Christian Henn [b]

[a] Department of Economics, University of Washington, Seattle, WA 98195, United States
[b] International Monetary Fund, 700 19th Street, N.W., Washington, DC 20431, United States

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ABSTRACT

The literature measuring the effects of WTO membership on trade flows has produced remarkably diverse results. Rose (2004) reports a wide range of empirical specifications that produce no WTO effects. Tomz et al. (2007) use Rose's data but include *de facto* WTO membership, to find positive WTO trade effects. Rose (2005) also produced positive WTO trade effects after accounting for the diverse trade effects produced by individual preferential trade agreements (PTAs). When Subramanian and Wei (2007) emphasize general equilibrium trade effects by controlling for multilateral resistance, they find strong WTO trade effects only for industrialized countries. Subramanian and Wei (2007), however, account neither for unobserved heterogeneity among trading partners, nor for differences in trade effects across PTAs (which could inflate WTO estimates). We unify the Rose, Tomz et al., and Subramanian and Wei specifications in one comprehensive approach that minimizes omitted variable bias to show that all specifications produce one consistent result: WTO effects on trade flows are not statistically significant, while PTAs produce strong but uneven trade effects. Extending the gravity model to address specific avenues in which WTO may have affected trade flows, we find that WTO membership boosts trade prior to PTA formation and increases trade among proximate developing countries (at the expense of distant trade). An augmented gravity model that accounts for WTO terms-of-trade theory shows that countries with greater incentives to bargain for tariff reductions before WTO accession experience positive and significant subsequent WTO trade effects.

1. Introduction

Reductions in trade barriers have been the hallmark of the World Trade Organization (WTO) and its predecessor, the General Agreement on Tariff and Trade (GATT).[1] While trade theory holds that tariff reductions should increase trade flows, the empirical literature on the effects of WTO membership has produced surprisingly ambiguous results. Rose (2004) initially documented the absence of WTO effects on bilateral trade flows. When Tomz et al. (2007, henceforth TGR) updated Rose's dataset to include both *de jure* and *de facto* WTO membership, they found positive WTO trade effects. Rose (2005) also produced a positive WTO impact on trade flows, after accounting for distinct effects of individual preferential trade agreements (PTAs).[2] Subramanian and Wei (2007, henceforth SW) then split the global sample to highlight that WTO trade effects exist for industrialized but not developing nations. This diversity of results in the empirical WTO literature seems to suggest that econometric specifications or data-coding conventions crucially influence the magnitude of WTO trade effects. A clear understanding as to what drives the diversity of results is highly relevant for policy makers and economists alike. Policy makers need to understand if and when gains from WTO can be expected, while economists seek to resolve whether datasets, coding, or empirical specifications drive results.

This paper unifies the above approaches to accessing WTO trade effects in order to produce four important insights. First, we show that the literature encompassing Rose, SW, and TGR generates one consistent result. These specifications all produce no evidence of positive WTO trade effects once we control comprehensively for three sources of omitted variable bias: multilateral resistance, unobserved bilateral heterogeneity, and individual PTA trade effects. Second, our robustness analysis shows that once the Rose, SW, and TGR approaches are unified, and their results correctly interpreted, multilateral resistance controls suffice to negate WTO trade effects. Third, when extending the gravity model to a version more suited to disentangle overlapping WTO and PTA membership, we find that WTO membership boosts trade effects just before PTA accession and increases trade among proximate developing countries, albeit at the

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* Corresponding author. Box 353330, University of Washington, Seattle, WA 98195, United States. Tel.: +1 206 685 8082.
E-mail addresses: te@u.washington.edu (T.S. Eicher), chenn@imf.org (C. Henn).

[1] Henceforth we use WTO as a synonym for GATT/WTO.
[2] See Table A5 in Rose (2005).

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expense of distant trade. Fourth, we extend the gravity model to include terms-of-trade theory (Bagwell and Staiger, 1999, 2002), which is specifically designed to analyze the effect of WTO membership. We find that countries with greater incentives to bargain for tariff reductions during WTO accession negotiations exhibit positive WTO trade effects.

Our main initial result of no WTO trade effects remains unchanged even after we account for *de facto* WTO membership (as TGR suggest), or when we code the WTO dummy using either the Rose or SW data-coding conventions.[3] In addition, the results are robust across time periods, WTO accession dynamics, and alternative types of bilateral heterogeneity controls. Our paper is the first to combine all three controls (multilateral resistance, unobserved bilateral heterogeneity, and individual PTA trade effects) in a large, bilateral trade dataset to examine WTO trade effects.[4]

The three sources of omitted variable bias that we address are shown to have exerted profound influence on estimates in the previous literature. The first omitted variable bias ensues when econometric specifications include only one average PTA control. With the exception of Rose (2005), this has been the case in previous WTO literature. Individual PTA trade effects matter, since preferential tariff reductions differ vastly across PTAs. When these individual trade effects are omitted from the empirical approach, the WTO coefficient may be biased upward if it assumes part of a positive, but omitted, PTA effect. The second omitted variables bias results when general equilibrium trade effects are not properly accounted for by comprehensive multilateral resistance controls as outlined in Anderson and van Wincoop (2003). Subramanian and Wei (2007) suggest that the absence of multilateral resistance controls in Rose (2004, 2005) biased his WTO results downward. The third potential omitted variable bias involves unobserved bilateral heterogeneity. In their illustrative derivation of the gravity model, Baldwin and Taglioni (2006) label the omission of country-pair fixed effects the "gold medal of classic gravity model mistakes." The authors derive the associated bias for coefficients of interest when two countries exhibit unobserved affinities for bilateral trade before joining a trade agreement (PTA or the WTO). The omission of country-pair fixed effects then renders WTO and/or PTA estimates biased upwards.[5]

The absence of WTO trade effects raises the question of whether WTO may foster trade in more subtle ways that cannot be identified by our basic framework unifying the Rose, SW, and TGR approaches. We extend the gravity framework in two dimensions to allow for specific trade effects that are unique to WTO members. The first extension disentangles WTO and PTA trade effects and explores a possible regional dimension of WTO trade creation. PTA accession is found to generate positive trade effects for WTO members and non-members alike. As theory predicts, the magnitude of these PTA trade effects is stronger for WTO non-members. Meanwhile trade flows between existing PTA members are hardly affected by WTO accession. There is evidence, however, that WTO membership increased trade flows prior to the formation of PTAs. In addition, WTO membership did foster regional trade integration among developing countries at the expense of more distant trade.

Our second extension incorporates proxies for the terms-of-trade theory of WTO, which, unlike the gravity model, has been specifically designed to model benefits of WTO membership. The terms-of-trade theory has been expounded in a series of papers by Bagwell and Staiger, who suggest that negotiations through GATT/WTO solve the terms-of-trade externality. Following Johnson's (1953-4) optimal tariff/retaliation argument, nations may hesitate to implement unilateral tariff reductions in the absence of WTO. The WTO terms-of-trade theory has received substantial support from Bagwell and Staiger (forthcoming) and Broda et al. (2008) in smaller datasets. Examining the trade gains due to WTO in 177 nations over 50 years, we find evidence in support of the terms-of-trade theory, even after controlling for the three sources of omitted variable bias that we discussed above. Specifically, those countries that had substantial incentives to negotiate tariff reductions during their WTO accession negotiations also exhibit significantly larger and positive WTO trade effects than other members, which are found to exhibit no WTO effects.

The paper is organized as follows. After presenting the data in Section 2, we unify the Rose, SW, and TGR approaches to WTO trade effect estimation in a unified baseline framework and demonstrate that all these approaches fail to find positive WTO trade effects (Section 3). Section 3 also provides a detailed discussion of the impacts of PTAs on trade, which are strong but uneven across individual agreements. Section 4 presents extensive robustness analysis for our unified baseline framework. Section 5 extends our unified framework in two directions to (i) further disentangle the effects of WTO and PTA membership and explore regional dimensions of WTO trade creation and (ii) to proxy for the terms-of-trade theory of WTO in the gravity model. Section 6 concludes.

2. Data

Our data is based on an updated version of SW's unbalanced panel.[6] Their bilateral trade values are derived from the IMF's *Direction of Trade Statistics,* deflated by the U.S. consumer price index. The dataset features not only a WTO dummy, but also a dummy that represents industrialized countries' unilateral trade concessions to developing trading partners under the GATT/WTO's Generalized System of Preferences (GSP) from 1979 onwards. We adjust the SW dataset to attribute a value of zero to GSP country-pairs that represent an industrialized country *exporting* to a developing country. The reasoning is that GSP is granted as a *unilateral* preference (for industrialized countries' *imports* from developing countries only). We also identify Luxembourg as a member of the European Union (EU) in 2000, and correct other minor coding errors identified by TGR. These changes do not affect our or SW's results qualitatively.

SW employ Rose's definition of *de jure* WTO membership. However, TGR indicated that *de facto* WTO members should also be considered. To illustrate that WTO trade effects vanish *even* when accounting for *de facto* membership, we use TGR's WTO membership definition throughout and refer the interested reader to the working paper version of this study which features analog results based on *de jure* WTO membership (Eicher and Henn, 2008). The conclusions are unaffected by the WTO membership definition. A single aggregate PTA indicator dummy has been prominent in a number of empirical trade flow studies (see e.g., Rose, 2000, 2004, 2005; Glick and Rose, 2002, SW, TGR), to capture the average effect of PTAs on trade flows. We extend the SW dataset and introduce a more extensive set of PTAs used by Rose (2005) and Eicher et al. (forthcoming) to properly account for trade effects of a large set of individual PTAs. Subsequent sections further modify the SW dataset. In Section 3, we include country-pair fixed effects to control for unobserved bilateral heterogeneity and introduce Rose's (2004, 2005)

[3] Rose/TGR coding of trade agreement memberships is *mutually inclusive* (dummies identify PTA and WTO memberships), while SW coding is *hierarchical, mutually exclusive* (dummies identify either PTA or WTO membership) as discussed in Section 3. The section also highlights that SW's coding convention is susceptible to producing biased WTO estimates when (a) industrialized and developing PTAs differ considerably in their trade effects and (b) these PTA effects are constrained to one average PTA coefficient. We show that in SW's dataset and WTO coding convention, their significantly positive WTO effect for industrialized countries is actually an industrialized country PTA trade effect.

[4] Baltagi et al. (2003) and Baier and Bergstrand (2007) motivated and included both multilateral resistance and unobserved heterogeneity controls. They did not examine WTO membership effects, however.

[5] See e.g., Egger (2000), Cheng and Wall (2005), Baltagi et al. (2003) and Baier and Bergstrand (2007).

[6] We use SW's preferred dataset, which excludes observations with import values of less than $500,000. A list of countries and their year of WTO accession is provided by Eicher and Henn (2008, Table A3).

mutually inclusive coding of trade agreements. Finally in Section 5 we extend the gravity model to account for WTO terms-of-trade theory.

The dimensionality of the dataset remains constant throughout, with 55,831 observations for 177 countries and 11,797 bilateral trade pairs in five year intervals from 1950 to 2000. The power of the regressions below depends on the number of observations in the dataset that change WTO membership status. Table A1 provides an overview of the changes in WTO membership status across country pairs in the panel. About 3834 changes in WTO status are observed at the country-pair level, which provides substantial power to the regressions. The chances of not rejecting the null when the alternative hypothesis is true are thus low.

3. A unified baseline framework to minimize omitted variable bias

In this section, we construct a unified framework from the Rose, SW, and TGR approaches to WTO trade effect estimation. We commence by extending SW to account for individual PTA trade effects to show that their industrialized country WTO effect is actually an industrialized PTA effect. We then further extend the framework to incorporate unobserved bilateral heterogeneity controls and allow for the alternative WTO coding convention employed by Rose.

3.1. Accounting for individual PTA trade effects

We begin by extending the gravity framework of SW to fully account for the impacts of all trade agreements (WTO, GSP, and individual PTAs). The SW setup has two important characteristics: first, in addition to time fixed effects, D_t, time-varying fixed effects are introduced for importers, D_{mt}, and exporters, D_{xt}, to capture multilateral resistance (see Baldwin and Taglioni, 2006). Multilateral resistance can be accounted for with these fixed effects, since any nation faces only one import/export price index at any point in time. The inclusion of time-varying importer and exporter effects requires, however, the dependent variable to be the log of bilateral imports, I_{mxt}, instead of the commonly used average trade flow variable.

The second important characteristic of SW's approach is their coding convention. SW code the trade agreement indicator dummies *mutually exclusively* to quantify "pure" GSP and WTO trade effects. SW's key assumption is that a PTA membership "represents the culmination of trade integration." Thus SW code trade agreement indicators such that *all* trade creation is *exclusively* attributed to PTAs, even if both trading partners are currently (or were previously) WTO/GSP members. For example, if trading partners are members of the WTO, GSP, and the same PTA, only the PTA dummy takes the value "1" in SW's coding convention. Coding is hierarchical throughout, so that when both WTO and GSP dummies could display a "1," only the GSP variable takes that value. The "*" superscripts below indicate mutually exclusive coding.

Our baseline specification replicates SW's preferred specification (regression 4 in Table 4 in SW, with the slight differences resulting from the TGR coding corrections discussed in Section 2).

$$I_{mxt} = a_0 + a_1 D_t + a_2 D_{mt} + a_3 D_{xt} + \beta_1 PTA_{mxt} + \beta_2 WTO\_Industrial^*_{mxt} + \beta_3 WTO\_Developing^*_{mxt} + \beta_4 GSP^*_{mxt} + \delta Z_{mxt} + \varepsilon_{mxt} \tag{1}$$

SW also disaggregate the WTO dummy to identify membership effects for industrialized and developed nations separately, $WTO\_Industrial_{mxt}$ and $WTO\_Developing_{mxt}$, respectively. Basic OLS regressions including only time fixed effects generate a WTO coefficient that represents the average difference in imports for country pairs that are WTO members vs. country pairs that are not WTO members. The controls for multilateral resistance also strip the WTO effect of time-varying effects common to any one importer or exporter.

In addition, the row vector, Z_{mxt}, is included, which represents a list of common gravity controls and proxies for transport costs and geographic/cultural proximity that are not absorbed by the fixed effects. The list includes the natural log of bilateral distance, $Distance_{mx}$, and dummies for common currency union ($CurrencyUnion_{mxt}$), contemporaneous or historical colonial relationships ($CurrentColony_{mxt}$ and $EverColony_{mx}$), common colonizer relationships post-1945 ($CommonColonizer_{mx}$), shared official languages ($CommonLanguage_{mx}$), and territorial dependency/contingency ($CommonNation_{mx}/Border_{mx}$). Some of the country-year specific regressors (for example, importer/exporter GDP) that can be found in canonical gravity equations are absorbed into the time-varying importer/exporter fixed effects. [7]

Our first extension of SW is to modify Eq. (1) by replacing the aggregate PTA vector, PTA_{mxt}, with dummies that allow each PTA to account for its own individual effect on bilateral imports. This converts PTA_{mxt} into a row vector and β_1 into a corresponding vector of regression coefficients that captures the trade impacts of individual PTAs. We disaggregate the PTA trade effects in two stages. First, we introduce only PTAs that are already contained in SW's and Rose's (2005) aggregate PTA dummy.[8] Then we enlarge the set of PTAs to include those suggested by the recent PTA literature (see Eicher et al., forthcoming).[9] Note that SW's mutually exclusive coding of trade agreements implies that the introduction of additional PTAs in this second stage diminishes the number of "1" entries in the WTO variables. Therefore, the introduction of additional PTAs may influence the WTO/GSP estimates for two reasons: a) by allowing individual PTAs to correct for omitted variable bias, and b) by reducing the WTO entries of PTA observations to "zero." If omitted PTAs are strongly trade creating and PTA members have also joined the WTO, the WTO coefficient in SW would be expected to be biased upward.

Before we discuss results, it is important to recall that SW's hierarchical coding convention assumes that PTA membership "represents the culmination of trade integration." Thus the SW PTA regression coefficients *include* the WTO effect. To compare results to the standard Rose coding method, we report SW PTA coefficients *net of* WTO trade effects in Table 1a.[10] Regressions 2 and 3 report results for SW's setup with the addition of individual PTA effects that we have added. The data corrections discussed in Section 2 increase the WTO and PTA coefficients slightly, but SW's original results are robust. Only industrialized countries are shown to benefit from WTO membership in regression 1 through a trade increase of 187% ($= e^{1.053} - 1$), while developing countries experience no WTO effect. With the exception of a common border and same nation status, all regressors that control for observable bilateral heterogeneity are highly significant, and they remain so throughout.

Regression 1 also indicates a highly significant coefficient associated with the aggregate PTA dummy in the original SW specification. The "pure" PTA effect, net of any WTO trade effects, is provided in Table 1a, regression 1 for industrialized and developing countries. Here we find that bilateral trade increases at a dramatically different rate for industrialized and developing country PTA members. Industrialized

[7] To obtain coefficients for the absorbed regressors, see Hsiao (1986, p. 50 f.) or Hausman and Taylor (1981).

[8] These are: ASEAN Free Trade Area, $AFTA_{mxt}$, the Australia-New Zealand Closer Economic Relations Trade Agreement, $ANZCERTA_{mxt}$, the Central American Common Market, $CACM_{mxt}$, the Caribbean Community/Carifta, $CARICOM_{mxt}$, the European Union (and its predecessor agreements), EU_{mxt}, the Southern Cone Common Market, $MERCOSUR_{mxt}$, the North America Free Trade Agreement, $NAFTA_{mxt}$, the South Pacific Regional Trade and Economic Cooperation Agreement, $SPARTECA_{mxt}$, and bilateral PTAs, $BilateralPTA_{mxt}$.

[9] Additional PTAs are: Asia-Pacific Economic Community, $APEC_{mxt}$, the Andean Pact, AP_{mxt}, the European Economic Area, EEA_{mxt}, the European Free Trade Association, $EFTA_{mxt}$, and the Latin America Integration Agreement/Lafta, $LAIA_{mxt}$. For a list of trade agreements and when countries entered, see Eicher and Henn (2008).

[10] Net PTA effects in Table 1a can be compared to Rose's Table 2 by subtracting WTO coefficients in Table 1b from the gross PTA coefficients in Table 1b using the Delta Method (e.g., Greene, 2003).

Table 1a
WTO and PTA Trade Effects. (Hierarchical, mutually exclusive coding.)

	Dependent variable: bilateral imports						
Regression #	1	1a	2	3	4	5	6
Estimation method	MLR	MLR	MLR	MLR	MLR & CPFE	MLR & CPFE	MLR & CPFE
WTO dummy coding	SW	SW	SW	SW	SW	SW	SW
WTO membership definition	TGR	TGR	TGR	TGR	TGR	TGR	TGR
Adj R^2	0.7411	0.7411	0.7415	0.7430	0.8747	0.8751	0.8760
F stat vs. regression #			# 1		# 1	# 2 # 4	# 3
Prob>F:			0.00		0.00	0.00 0.00	0.00
GSP_{mt} (ind. importer grants)	−0.209 ***	−0.209 ***	−0.202 ***	−0.202 ***	−0.242***	−0.233***	−0.181 ***
	(0.041)	(0.041)	(0.041)	(0.040)	(0.049)	(0.049)	(0.047)
$WTO\_Developing_{mt}$	0.062	0.062	−0.045	−0.235***	0.210	−0.207	−0.254***
	(0.134)	(0.134)	(0.131)	(0.075)	(0.147)	(0.148)	(0.082)
$WTO\_Industrial_{mt}$	1.053 ***	−0.152 ***	0.588 ***	−0.035	0.393 **	0.124	−0.068
	(0.141)	(0.056)	(0.195)	(0.092)	(0.165)	(0.189)	(0.092)
PTA_Ind_{mxt} (based on aggregate in Table 1b)	0.152 ***				0.361 ***		
	(0.056)				(0.055)		
PTA_Dev_{mxt} (based on aggregate in Table 1b)	1.143 ***				0.545 ***		
	(0.073)				(0.082)		
PTA_Ind_{mxt}		dropped					
PTA_Dev_{mxt}		1.143 ***					
		(0.073)					
$Bilateral\_PTA\_Ind_{mxt}$			0.072	0.351***		−0.169 *	−0.054
			(0.109)	(0.085)		(0.101)	(0.084)
$Bilateral\_PTA\_Dev_{mxt}$			0.705 ***	0.551 ***		0.161 *	0.132 *
			(0.085)	(0.080)		(0.086)	(0.180)
$NAFTA\_Ind_{mxt}$			1.529 ***	1.276 ***		0.155	0.257
			(0.188)	(0.161)		(0.139)	(0.169)
$NAFTA\_Dev_{mxt}$			2.162 ***	1.476 ***		0.485 ***	0.443 **
			(0.216)	(0.172)		(0.158)	(0.176)
EU_{mxt}			0.065	−0.139 **		0.473 ***	0.306 ***
			(0.056)	(0.067)		(0.062)	(0.068)
$CACM_{mxt}$			1.346 ***	1.509 ***		1.907 ***	1.938 ***
			(0.191)	(0.173)		(0.253)	(0.242)
$CARICOM_{mxt}$			1.437 ***	1.446 ***		1.071 ***	1.070 ***
			(0.143)	(0.143)		(0.253)	(0.253)
$MERCOSUR_{mxt}$			1.530 ***	1.552 ***		0.982 ***	1.068 ***
			(0.197)	(0.202)		(0.227)	(0.243)
$AFTA_{mxt}$			0.568***	0.391**		−0.035	0.049
			(0.159)	(0.172)		(0.196)	(0.207)
$ANZCERTA_{mxt}$			1.875 ***	1.647 ***		0.747 **	0.770 ***
			(0.123)	(0.135)		(0.293)	(0.297)
$SPARTECA\_Ind_{mxt}$			0.824 ***	1.238 ***		0.587 ***	0.742 ***
			(0.236)	(0.224)		(0.207)	(0.178)
$SPARTECA\_Dev_{mxt}$			1.457 ***	1.438 ***		0.917 ***	0.927 ***
			(0.210)	(0.217)		(0.176)	(0.170)
$EFTA_{mxt}$				0.491 ***			0.200 ***
				(0.091)			(0.075)
EEA_{mxt}				0.508 ***			0.342 ***
				(0.096)			(0.105)
AP_{mxt}				0.638 ***			0.677 ***
				(0.183)			(0.223)
$LAIA_{mxt}$				0.267 **			1.277 ***
				(0.105)			(0.191)
$APEC\_Ind_{mxt}$				0.651***			0.107
				(0.086)			(0.092)
$APEC\_Dev_{mxt}$				0.851 ***			0.293 ***
				(0.070)			(0.079)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses. GSP and PTA coefficients are net of WTO effects, see footnote 10. All regressions include time fixed effects. Fixed effects results are not reported. MLR indicates multilateral resistance controls, i.e. time-varying importer and exporter fixed effects. CPFE indicates unobserved bilateral heterogeneity controls, i.e. country-pair fixed effects. Additional regressors not reported include all common gravity variables discussed in the text (Eq. (1)). All these regressors are significant at the 1% level throughout, except common nation and border. Full results are available from the authors upon request.

countries see their trade increase by a meager 16% ($=e^{0.152}-1$) while trade in developing countries increases 214% ($=e^{1.143}-1$).

Regression 2 allows for individual PTA trade effects. We find that all multilateral trade dummies are heavily impacted, confirming the suspicion of potential upward bias in regression 1. Most notably, we find that the economic and statistical significance of WTO membership for industrialized countries is reduced by an order of magnitude. WTO-induced trade creation for industrialized countries falls from 187% in regression 1 to 80% ($=e^{0.588}-1$) in regression 2. This suggests that the sizable trade creation that was attributed to WTO membership in regression 1 is more accurately associated with individual PTAs. The precision of the individual PTA estimates together with the F-Statistic (which rejects regression 1 in favor of regression 2) suggests that regression 1 suffers from omitted variable bias.

The individual PTAs are all trade creating and highly statistically significant (with the exception of the EU, a case we discuss at length below). PTAs promote trade strongly, but unevenly. The magnitudes of trade creation implied by the estimates varies dramatically from 770% ($=e^{2.162}-1$) for NAFTA to 76% ($=\exp^{0.568}-1$) for the ASEAN free trade agreement (AFTA). Regression 3 controls for additional PTAs that were not included in SW (or Rose, 2005), but feature prominently in the empirical PTA literature.[11] EFTA, EEA, LAIA, and APEC are all shown to be highly trade creating. Once we account for the individual PTA trade effects with the most comprehensive set of PTAs, we find that SW's result of positive WTO trade effects in industrial countries vanishes.

Allowing for individual PTA trade effects also generates seemingly bleak insights for developing countries. The effects of WTO membership are estimated to be either nonexistent (regressions 1 and 2) or negative and statistically significant ($-21\%=e^{-0.235}-1$ in regression 3). On the upside, however, PTA coefficients indicate strong trade creation for *all* PTAs involving developing countries. It may well be the case that developing countries reoriented their import activity considerably towards PTA partners after joining PTAs; we will examine the robustness of this result below. The largest levels of trade creation in regression 3 are observed for PTAs that consist of developing nations (CACM, CARICOM, MERCOSUR, NAFTA-Developing and SPARTECA-Developing) where PTA-internal trade is estimated to be roughly 350% ($=e^{1.5}-1$) greater than PTA-external trade.

3.2. "Industrialized WTO" and "industrialized PTA" trade effects

Our results above suggest that PTA trade creation is not homogeneous across agreements. To the contrary, bilateral PTA trade is estimated to be significantly larger for PTAs that consist of developing countries than for industrialized trading partners that belong to the same PTA. Ignoring such differentials in trade creation introduces a specific type of omitted variable bias into SW style regressions. This section discusses this bias and outlines its impact on SW's key parameter: the industrialized country WTO coefficient. SW's results are easily misinterpreted due to their coding convention. We will show that their coding convention, together with their fixed effects specification generates an *"implicit industrialized PTA"* dummy. This dummy is the result of the interaction between a) SW's mutually exclusive coding convention, b) multilateral resistance controls (time-varying importer dummies), and c) the nature of industrialized countries' WTO/PTA accession in this particular dataset.

Since *all* industrialized importers in the dataset joined the WTO *before* joining a PTA, SW code either the PTA dummy or the industrialized WTO/GSP dummy as "1" for years after industrialized importers joined a PTA. As a result, the linear combination of these two dummies is perfectly collinear with industrialized countries' time-varying importer dummies. This led SW to mislabel the industrial countries' PTA effect as a industrial countries' WTO effect. To highlight the issue, we perform a quick experiment. Regression 1a partitions SW's aggregate PTA dummy into two separate effects: one for industrialized importers and another for developing importers featuring PTA relationships. The experiment results in perfect multicollinearity between the industrialized WTO dummy and the industrialized PTA dummy, forcing one or the other to be dropped from the regression. A comparison of regressions 1 and 1a (Table 1a) then reveals that the "*industrialized WTO*" dummy in SW's regression 1 is actually an *"industrialized PTA"* effect, with only the signs reversed because of SW's coding convention.

3.3. Accounting for unobserved bilateral heterogeneity

The PTA literature has long considered various PTA estimates (such as AP, LAIA and APEC) as suspiciously high, relative to the small tariff reductions associated with these agreements (e.g. Frankel, 1992; Frankel and Wei, 1993; Frankel et al., 1995; Frankel, 1997). There exists substantial evidence, however, that trade agreements tend to form between trading partners whose bilateral trade has been "naturally" elevated all along, due to unobserved characteristics (see e.g. Baier and Bergstrand, 2007). This raises the risk of incorrectly attributing "natural" trade-promoting characteristics to trade agreements. SW attempted to control for such country-pair specific characteristics with the inclusion of a number of control variables contained in the vector Z above. However, the resulting WTO and PTA coefficients are biased upwards, if the included controls do not account for *all* bilateral heterogeneity, for instance because some is unobserved.

Egger and Pfaffermayr (2003) and Cheng and Wall (2005) previously confirmed that implausibly large estimates can be lowered by accounting for unobserved bilateral heterogeneity (neither study includes GSP or WTO trade effects). To fully account for all time-invariant bilateral heterogeneity in Eq. (1), we follow the previous literature and replace the intercept in Eq. (1) with a country-pair specific dummy, D_{mx}:

$$\begin{aligned} I_{mxt} = {} & a_0 D_{mx} + a_1 D_t + a_2 D_{mt} + a_3 D_{xt} + \beta_1 PTA_{mxt} + \beta_2 WTO\_Industrial^*_{mxt} \\ & + \beta_3 WTO\_Developing^*_{mxt} + \beta_4 GSP^*_{mxt} + \delta_1 CurrencyUnion_{mxt} \\ & + \delta_2 CurrentColony_{mxt} + \varepsilon_{mxt}. \end{aligned} \tag{2}$$

Eq. (2) features fewer explanatory variables than Eq. (1) because all time-invariant regressors are now absorbed into the pair-specific fixed effects. When country-pair fixed effects are added, the WTO trade effect is also stripped of any average, time-invariant effect between trading partners.

Regressions 4–6 in Table 1a present analogs of regressions 1–3 above. The only differences are the added country-pair specific fixed effects that constitute comprehensive controls for unobserved bilateral heterogeneity. Two areas are the focus of our interest: a) whether WTO trade effects are influenced by unobserved bilateral heterogeneity, and b) whether PTA trade effects are reduced to plausible ranges. Regression 4 replicates the original SW specification (regression 4 in SW Table 4) with the addition of country-pair fixed effects.

The results show a substantial reduction of the WTO's economic and statistical significance for industrialized countries. WTO trade creation falls from 178% in regression 1 to 48% ($=e^{0.393}-1$) for industrialized countries in regression 4, and its statistical significance is reduced to the 5% level. The inclusion of specific PTA trade effects in regressions 5 and 6 again negate all WTO trade effects for industrialized countries. Regression 5 even shows that the simple disaggregation of SW's own PTA dummy is sufficient to neutralize any industrialized WTO trade effects, once we account for heterogeneity in bilateral relationships. For developing countries, in sharp contrast, regressions 4–6 illustrate that the GSP and WTO trade effects are hardly impacted by unobservable heterogeneity. Their estimates are closely aligned with those in regressions 1–3.

The F statistics confirm the importance of the inclusion of comprehensive country-pair fixed effects, and individual (in lieu of aggregate) PTA trade effects at significance levels exceeding 0.001%. The added controls also generate reduced (and more plausible) trade impacts for individual PTAs. This provides evidence that unobserved bilateral heterogeneity is generally trade-enhancing. Average trade creation across PTAs drops from 234% in regression 1 to 123% in regression 4 (Table 1b). Most individual trade agreements see their trade effects at least halved in regressions 5 and 6. In contrast, trade creation in the EU increases, and the EU coefficient is now estimated

[11] For our broadest PTA set, 3253 of the 55,813 observations are country pairs that are also members of a common PTA. Of these 3253 observations, 2700 of the importers are also contemporaneous WTO members.

Table 1b
Background results for Table 1a. This table reports the raw regression output for SW's mutually exclusive coding convention. Mutually exclusive coding implies that GSP and PTA coefficients *include* WTO Effects. The effects for all regressors *net of WTO effects* are reported in Table 1a.

	Dependent variable: bilateral imports						
Regression #	1	1a	2	3	4	5	6
Estimation method	MLR	MLR	MLR	MLR	MLR & CPFE	MLR & CPFE	MLR & CPFE
GSP_{mt} (*ind. importer grants*)	0.844 ***	−0.360***	0.385 **	−0.237 **	0.151	−0.110	−0.249 ***
	(0.142)	(0.115)	(0.194)	(0.093)	(0.166)	(0.187)	(0.095)
$WTO\_Developing_{mt}$	0.062	0.062	−0.045	−0.235 ***	0.210	−0.207	−0.254 ***
	(0.134)	(0.134)	(0.131)	(0.075)	(0.147)	(0.148)	(0.082)
$WTO\_Industrial_{mt}$	1.053 ***	−0.152 ***	0.588 ***	−0.035	0.393 **	0.124	−0.068
	(0.141)	(0.056)	(0.195)	(0.092)	(0.165)	(0.189)	(0.092)
PTA_{mxt}	1.205 ***				0.755 ***		
	(0.131)				(0.157)		
$PTA\_Industrial_{mxt}$		Dropped					
$PTA\_Developing_{mxt}$		1.205 ***					
		(0.131)					
$Bilateral\_PTA_{mxt}$			0.660 ***	0.316 ***		−0.045	−0.122
			(0.164)	(0.095)		(0.167)	(0.097)
$NAFTA_{mxt}$			2.117 ***	1.241 ***		0.279	0.189
			(0.258)	(0.148)		(0.215)	(0.152)
EU_{mxt}			0.652 ***	−0.174 *		0.596 ***	0.238 **
			(0.199)	(0.097)		(0.196)	(0.107)
$CACM_{mxt}$			1.301 ***	1.274 ***		1.700 ***	1.684 ***
			(0.168)	(0.166)		(0.244)	(0.239)
$CARICOM_{mxt}$			1.391 ***	1.211 ***		0.864 ***	0.817 ***
			(0.209)	(0.165)		(0.288)	(0.264)
$MERCOSUR_{mxt}$			1.485 ***	1.317 ***		0.776 ***	0.814 ***
			(0.246)	(0.198)		(0.271)	(0.233)
$AFTA_{mxt}$			0.523 ***	0.156		−0.241	−0.204
			(0.188)	(0.154)		(0.237)	(0.194)
$ANZCERTA_{mxt}$			2.463 ***	1.612 ***		0.870 **	0.702 **
			(0.227)	(0.132)		(0.347)	(0.295)
$SPARTECA_{mxt}$			1.412 ***	1.203 ***		0.710 ***	0.674 ***
			(0.199)	(0.207)		(0.178)	(0.164)
$EFTA_{mxt}$				0.455 ***			0.132
				(0.122)			(0.114)
EEA_{mxt}				0.472 ***			0.275 ***
				(0.069)			(0.065)
AP_{mxt}				0.404 **			0.423 **
				(0.166)			(0.214)
$LAIA_{mxt}$				0.032			1.023 ***
				(0.110)			(0.185)
$APEC_{mxt}$				0.616 ***			0.039
				(0.090)			(0.098)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses. All regressions include time fixed effects. Fixed effects results are not reported. MLR indicates multilateral resistance controls, i.e. time-varying importer and exporter fixed effects. CPFE indicates unobserved bilateral heterogeneity controls, i.e. country-pair fixed effects. Additional regressors not reported include all common gravity variables discussed in the text (Eq. (1)). All these regressors are significant at the 1% level throughout, except Common Nation and Border. Full results are available from the authors upon request.

with considerable precision. The case of the EU is discussed in Section 3.5, when we examine the individual PTA trade effects in detail.

The expanded SW results in regressions 1–6 give rise to the question whether the absence of WTO trade effects may be an artifact of the hierarchical, mutually exclusive coding in SW. As discussed above, when increases in trade flows are attributed to PTAs rather than to both, WTO and PTA memberships, one may suspect that the SW coding convention underestimates WTO trade effects. More problematically, the industrialized WTO effect is actually an implicit "industrialized PTA" effect, given the structure of the data. In the next section, we apply the more conventional, *mutually inclusive* WTO coding (as in Rose, 2004) to allow for separate identification of WTO and PTA trade effects. This analysis has two purposes. Not only will it settle whether the SW coding convention is driving the results, but it is also a substantive extension of Rose (2004, 2005), because we extend his specification to introduce both disaggregate PTAs as well as comprehensive multilateral resistance controls.

3.4. Accounting for Alternative WTO Coding Conventions

Rose (2004, 2005) controlled for unobserved bilateral heterogeneity through country-pair fixed effects, but SW noted that Rose did not include the most comprehensive multilateral resistance controls. By introducing comprehensive controls for multilateral resistance to Rose's dataset, we unify the SW and Rose approaches. Our approach controls for all three key determinants of trade under both coding conventions: unobserved bilateral heterogeneity, multilateral resistance, and individual PTA trade effects. The unified approach allows us to highlight whether any results are due to mutually inclusive (Rose) or mutually exclusive (SW) coding of WTO dummies. To allow for a comparison between Rose and SW coding results, we split Rose's inclusive WTO-dummy into SW-style indicators for industrialized and developing importers' WTO membership.

An exact comparison between Rose's standard coding convention and SW can be achieved by simply reproducing regressions 1–3 with mutually inclusive coding.

$$I_{mxt} = a_0 + a_1 D_t + a_2 D_{mt} + a_3 D_{xt} + \beta_1 PTA_{mxt} + \beta_2 WTO\_Industrial^{**}_{mxt} + \beta_3 WTO\_Developing^{**}_{mxt} + \beta_4 GSP^{**}_{mxt} + \delta Z_{mxt} + \varepsilon_{mxt} \quad (1')$$

Eq. (1′) is essentially Rose's specification *with* multilateral resistance controls. The only difference between Eqs. (1) and (1′) is

Table 2
WTO and PTA trade effects. (Inclusive coding.)

	Dependent variable: bilateral imports							
Regression #	7	8	9	10	11		12	
Estimation method	MLR	MLR	MLR	MLR & CPFE	MLR & CPFE		MLR & CPFE	
WTO dummy coding	Rose	Rose	Rose	Rose	Rose		Rose	
WTO membership definition	TGR	TGR	TGR	TGR	TGR		TGR	
Adj R^2	0.7401	0.7415	0.7431	0.8747	0.8751		0.8760	
F stat vs. regression #		# 7	# 8	# 7	# 8	# 10	# 9	# 11
Prob>F:		0.00	0.00	0.00	0.00	0.00	0.00	0.00
GSP_{mxt} (ind. importer grants)	−0.127 ***	−0.183 ***	−0.181 ***	−0.252 ***	−0.228 ***		−0.187 ***	
	(0.040)	(0.040)	(0.039)	(0.048)	(0.049)		(0.048)	
$WTO\_Developing_{mxt}$	−0.103	−0.118 *	−0.109 *	−0.051	−0.054		−0.035	
	(0.065)	(0.065)	(0.065)	(0.069)	(0.069)		(0.068)	
$WTO\_Industrial_{mxt}$	0.058	0.069	0.071	−0.028	−0.026		−0.002	
	(0.065)	(0.065)	(0.065)	(0.068)	(0.068)		(0.066)	
PTA_{mxt}	0.629 ***			0.473 ***				
	(0.052)			(0.045)				
$Bilateral\_PTA_{mxt}$		0.453 ***	0.470 ***		0.070		0.059	
		(0.069)	(0.070)		(0.068)		(0.068)	
$NAFTA_{mxt}$		1.755 ***	1.194 ***		0.323 **		0.150	
		(0.145)	(0.153)		(0.155)		(0.171)	
EU_{mxt}		0.072	−0.146 **		0.506 ***		0.312 ***	
		(0.056)	(0.064)		(0.060)		(0.068)	
$CACM_{mxt}$		1.316 ***	1.372 ***		1.788 ***		1.789 ***	
		(0.163)	(0.164)		(0.238)		(0.238)	
$CARICOM_{mxt}$		1.447 ***	1.454 ***		1.065 ***		1.061 ***	
		(0.144)	(0.143)		(0.254)		(0.254)	
$MERCOSUR_{mxt}$		1.540 ***	1.402 ***		0.984 ***		0.879 ***	
		(0.197)	(0.196)		(0.228)		(0.229)	
$AFTA_{mxt}$		0.551 ***	0.189		−0.043		−0.161	
		(0.156)	(0.154)		(0.195)		(0.194)	
$ANZCERTA_{mxt}$		1.877 ***	1.514 ***		0.747 **		0.623 **	
		(0.124)	(0.124)		(0.292)		(0.291)	
$SPARTECA_{mxt}$		1.310 ***	1.350 ***		0.814 ***		0.807 ***	
		(0.206)	(0.208)		(0.168)		(0.167)	
$EFTA_{mxt}$			0.473 ***				0.199 ***	
			(0.091)				(0.075)	
EEA_{mxt}			0.489 ***				0.294 ***	
			(0.067)				(0.069)	
AP_{mxt}			0.385 **				0.424 *	
			(0.167)				(0.217)	
$LAIA_{mxt}$			0.189 *				1.190 ***	
			(0.098)				(0.182)	
$APEC_{mxt}$			0.798 ***				0.244 ***	
			(0.063)				(0.071)	

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses. All regressions include time fixed effects. Fixed effects results are not reported. MLR indicates multilateral resistance controls, i.e. time-varying importer and exporter fixed effects. CPFE indicates unobserved bilateral heterogeneity controls, i.e. country-pair fixed effects. Additional regressors not reported include all common gravity variables discussed in the text (Eq. (1)). All these regressors are significant at the 1% level throughout, except common nation and border. Full results are available from the authors upon request.

the coding convention, where inclusive coding is now denoted by "**" superscripts.[12] The WTO dummy is again disaggregated to identify membership effects for industrialized and developed nations, $WTO\_Industrial^{**}_{mxt}$ and $WTO\_Developing^{**}_{mxt}$, as well as for GSP^{**}_{mxt} trading partners. Crucial is that under Rose's coding convention *both* WTO and GSP variables take on the value "1" when the two conditions are fulfilled. In addition, when the same trading partners are members in a common PTA, mutually inclusive coding assigns the value "1" to *all three* dummies. For comparison purposes, it is important to point out that inclusive coding delivers coefficient estimates that represent pure PTA trade effects. There is no need to produce net effects via the Delta method as in the SW case.

We proceed again in stages. First we provide results based on Eq. (1′), and then we add country-pair fixed effects to control for all time-invariant bilateral heterogeneity. Inclusion of country-pair effects again converts the constant a_0 to a pair-specific one, D_{mx}:

$$I_{mxt} = a_0 D_{mx} + a_1 D_t + a_2 D_{mt} + a_3 D_{xt} + \beta_1 PTA_{mxt} + \beta_2 WTO\_Industrial^{**}_{mxt} + \beta_3 WTO\_Developing^{**}_{mxt} + \beta_4 GSP^{**}_{mxt} + \delta_1 Currency\ Union_{mxt} + \delta_2 CurrentColony_{mxt} + \varepsilon_{mxt}. \quad (2')$$

Regression 7 in Table 2 establishes a baseline regression that represents our closest analog to Rose's preferred regressions (Table 1 in Rose, 2004). It represents a robustness test of Rose's findings that examines whether comprehensive accounting for multilateral resistance affects his original results. Regression 7 also provides a robustness

[12] Strictly speaking, there exists one additional discrepancy between mutually inclusive and exclusive coding of multilateral trade agreements. Mutually exclusive coding assigns a "1" to any WTO importer (without PTA or GSP relationship), while inclusive coding assigns a "1" only when *both* importer and exporter are WTO members. The reason is the collinearity between inclusive WTO-dummies and multilateral resistance controls. Specifically, the inclusive WTO dummy takes the value "1" for all observations that relate to a member countries' trade in a given year, which establishes collinearity with the importer-year dummy that controls for multilateral resistance. By construction, this collinearity is avoided in mutually exclusive coding since WTO importers are not considered WTO members when the WTO importer is in a PTA or GSP relationship with the exporter.

check of SW's preferred regression (Table 1a, regression 1) to examine whether results are affected by the coding convention.

Rose's preferred regressions report insignificant WTO trade effects throughout. The insertion of multilateral resistance controls does not change Rose's conclusions regarding WTO trade effects. Regression 7 shows that trade creation due to WTO membership for both industrialized and developing countries is insignificant. Even when the industrial and developing country WTO dummies in regression 7 are aggregated to one WTO dummy, the effect remains insignificant (see Table 5, regression 13). The only change is that Rose's GSP effect is eliminated.

A comparison of regressions 1a and 7 highlights that coding conventions do not drive our key conclusion. Both mutually inclusive and exclusive coding conventions render the WTO effect statistically and economically insignificant. In Rose-style mutually inclusive coding, this is more immediately apparent, because PTA coefficients provide *net effects*: hence the industrialized WTO dummy cannot function as an error-correction term. Under mutually inclusive coding it is not possible for a WTO dummy to implicitly split the aggregate PTA variable into North–North PTAs (with lower net trade creation) and South–South PTAs (with higher net trade creation) as in the case of mutually exclusive coding above. As a result, mutually inclusive coding in Table 2 can never deliver significant WTO coefficients.

The significant WTO coefficient in SW was only significant because mutually exclusive coding produced PTA coefficients that include *both* PTA and WTO trade effects. Given the structure of the data this implies that the industrialized WTO effect was actually an industrialized PTA effect. Hence we point to the important insight that the *net effect* generated by mutually inclusive coding significantly reduces the risk of omitted variable bias. Mutually exclusive coding, on the other hand, holds the danger that WTO dummies are biased when the following two conditions are met: (1) industrialized and developing PTAs differ considerably in their trade effects and (2) individual PTA trade effects are constrained to an average coefficient associated with one aggregate PTA dummy.

Regressions 8 and 9 (the inclusive-coding analogs of regressions 2 and 3) introduce individual PTA trade effects and represent two further robustness tests. The first test is whether Rose's (2005) results of a small, positive WTO effect are robust to controlling for multilateral resistance. At the same time, regressions 8 and 9 represent a second robustness test that examines whether SW's WTO effect vanishes only because of their coding convention. Regressions 8 and 9 overturn Rose's (2005) result of a statistically significant WTO effect when individual PTA trade effects are considered. Hence Rose's (2005) finding of small (but significant) WTO trade effects came about only because he did not control comprehensively for multilateral resistance. The regressions that allow for individual PTA trade effects also highlight that the vanishing WTO effect in SW – after we controlled for individual PTA trade effects – was indeed only due to SW's hierarchical and mutually exclusive coding.

Regressions 10–12 are the mutually inclusively coded analogs to regressions 4–6. With inclusive coding we find no WTO trade effects in these regressions. Regression 12 represents our preferred specification, since it features the more reliable coding convention and contains all three key controls: multilateral resistance, natural trading partner effects, and individual PTA trade effects. It confirms the absence of significant WTO trade gains. The validity of this regression is also established by the F-statistics: without exception we find that the inclusion of individual PTA trade effects *always* improves the estimation no matter which set of controls is selected.

3.5. Trade effects of individual PTAs

Examining WTO trade effects while controlling for individual PTA trade effects along with the most comprehensive set of fixed effects, raises the question how our PTA results compare to the voluminous literature on PTA effects on trade flows. We focus on our preferred specification, regression 12. The PTA coefficients in regression 12 indicate how much PTA-internal trade increased *relative to* trade with non-members for a typical importer/exporter pair. Hence, even insignificant PTA estimates do not imply that the PTA was necessarily ineffective in creating trade flows, only that PTA-internal trade grew at the rate of external trade. Note also that Anderson and van Wincoop's (2003) multilateral resistance renders trade creation and diversion indistinguishable, so that our results below refer to the sum of trade creation and diversion.[13]

The individual trade effects of PTAs in regression 12 are the most "reasonable" among all our regressions, in the sense that the net trade creation for most PTAs is estimated to range between 30% and 80%. Curiously, the Central and Latin American Trade Agreements (CACM, CARICOM, MERCOSUR and LAIA) show the largest increases in relative trade. These are also the only PTAs that report net increases in trade creation of over 100% (with the exception of SPARTECA, which reports an increase of $124\% = e^{0.807} - 1$). Note, however, that with the exception of LAIA, the implied trade effects are all lower than in the original Rose or SW specifications in Tables 1a and 2.

Most notable is the reduction in the estimated net trade creation for most PTAs, after we control for unobserved bilateral heterogeneity. Comparing regressions 9 and 12 reveals that, with the exception of the EU, CACM and LAIA, all PTA estimates are substantially reduced when we include country-pair fixed effects. In other words, our results suggest strongly that PTAs are formed between countries that have been sharing characteristics favorable to mutual trade all along. In this case, tariff reduction may simply be an afterthought. Controlling for unobserved bilateral heterogeneity also improves the precision of the estimates in all cases but NAFTA. The suspiciously large net trade creation of NAFTA (230%) in regression 9 is reduced to insignificance after we control for unobserved bilateral heterogeneity.

Controlling for unobserved bilateral heterogeneity also *increased* trade creation for three PTAs (EU, EFTA and CACM). Given these PTA member countries' characteristics, their actual trade flows are not *large enough* relative to the prediction of the gravity model. For example, in the case of CACM, all countries share a common language, colonizer, and proximity. The introduction of country-pair fixed effects resolves the systematic overprediction of the gravity model in this case and allows a better assessment of the impact of PTA accession. The fact that the EU trade impact is underestimated in both the traditional and multilateral resistance-augmented versions of the gravity equation is well known (e.g., Aitken, 1973; Rose, 2004). Our estimates show a statistically significant 37% ($= e^{0.312} - 1$) increase in trade due to EU accession, once we control for unobserved bilateral heterogeneity. In the case of the EU, it is likely that the large market and the strong harmonization efforts allowed firms to overcome trade fixed costs that subsequently led to strong trade creation between both member and non-member countries (e.g. Freund, 2000; Melitz, 2003). The increase in absolute trade volume among EU members that we observe seems then reasonably small compared to trade increases with non-members. It is also important to note that our estimates imply that EU members reaped another 34% trade benefit when they became EEA members in 1994.

Another trade agreement that has been the subject of great interest in the PTA literature is APEC. Highly significant and truly exorbitant APEC trade creation estimates (around 300%) have been common in the gravity literature, although APEC is only a forum without implications for tariffs (see e.g. Frankel and Wei, 1993; Frankel et al., 1995; Frankel, 1997). Regression 9 indicates that the inclusion of multilateral resistance lowers values for the APEC coefficient substantially to 123% ($= e^{0.802} - 1$). The inclusion of unobserved bilateral heterogeneity controls shows that much of the trade creation originally attributed to

[13] Multicollinearity does not allow for separate trade creation/diversion effects in the presence of multilateral resistance controls. For a given year, a typical PTA member country's import observations are partitioned into imports originating from (a) fellow PTA members, and (b) non-members. The linear combination of these two dummy variables is perfectly collinear with the time-varying importer dummies that control for multilateral resistance.

Table 3
Robustness of trade effects. Alternative approaches to control for unobserved bilateral heterogeneity: first-differencing and AR(1) errors.

	Dependent variable: bilateral imports				
Regression #	13	14	15	16	17
Estimation method:	CPFE	First-differenced	First-differenced MLR	AR(1)	AR(1) & MLR
WTO dummy coding	Rose	Rose	Rose	Rose	Rose
WTO membership definition	TGR	TGR	TGR	TGR	TGR
Number of observations	54,389	40,066	40,935	54,389	54,389
Adj R^2	0.7983	0.1087	0.2770	0.6380	0.7446
Estimated Autocorrelation Coef.				0.5026	0.4576
PTA_{mxt}	0.713***	0.244***	0.178***	0.569***	0.435***
	(0.050)	(0.030)	(0.033)	(0.035)	(0.035)
GSP_{mxt} (ind. importer grants)	−0.189***	−0.039	−0.052	−0.030	−0.153***
	(0.033)	(0.029)	(0.040)	(0.022)	(0.030)
$WTO\_Industrial_{mxt}$	0.381***	0.102***	0.009	0.387***	0.075*
	(0.036)	(0.029)	(0.060)	(0.022)	(0.040)
$WTO\_Developing_{mxt}$	0.102***	0.120***	0.124*	0.088***	−0.005
	(0.029)	(0.024)	(0.063)	(0.018)	(0.039)
GDP_{mt} (log of importer GDP)	0.455***	0.677***		0.650***	
	(0.049)	(0.060)		(0.009)	
GDP_{xt} (log of exporter GDP)	0.489***	0.313***		0.616***	
	(0.049)	(0.073)		(0.009)	
$GDPpc_{mt}$ (log of importer GDP per capita)	0.318***	0.031		0.163***	
	(0.050)	(0.058)		(0.014)	
$GDPpc_{xt}$ (log of exporter GDP per capita)	0.413***	0.441***		0.354***	
	(0.049)	(0.074)		(0.013)	
$Landlocked_m$ (importer is landlocked)				−0.477***	
				(0.034)	
$Landlocked_x$ (exporter is landlocked)				−0.320***	
				(0.034)	
$Island_m$ (importer is island)				0.152***	
				(0.034)	
$Island_x$ (exporter is island)				0.149***	
				(0.035)	
$Area_m$ (log of importer land mass)				−0.060***	
				(0.007)	
$Area_x$ (log of exporter land mass)				0.001	
				(0.007)	
CU_{mxt} (currency union)	0.695***	0.244**	0.690	0.518***	0.370***
	(0.153)	(0.112)	(0.630)	(0.062)	(0.059)
$CurrentColony_{mxt}$ (current colony)	0.517***	0.249***	0.950***	0.539***	1.219***
	(0.156)	(0.070)	(0.103)	(0.105)	(0.066)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses. All regressions include time fixed effects. Fixed effects results are not reported. MLR indicates multilateral resistance controls, i.e. time-varying importer and exporter fixed effects. CPFE indicates unobserved bilateral heterogeneity controls, i.e. country-pair fixed effects.

APEC was due to bilateral unobservables. While this had been the suspicion of Frankel and coauthors all along, their quest to identify these unobservable drivers has largely been unsuccessful. In our preferred regression 12, which controls for multilateral resistance, unobserved bilateral heterogeneity, and individual PTA trade effects, the APEC's trade creation estimate drops to 28% ($=e^{0.244}-1$).

4. Robustness

In this section we examine whether WTO trade effects are sensitive to alternative econometric approaches that account for unobserved bilateral heterogeneity or accession dynamics. Since most industrialized countries joined the WTO early, we also examine whether WTO trade effects differ for early or late joiners, and whether WTO trade effects differ across decades.

4.1. Unobserved bilateral heterogeneity and WTO trade effects

Above we hold that a comprehensive approach to addressing both multilateral resistance and unobserved bilateral heterogeneity is crucial to obtaining unbiased WTO estimates. Instead of the country-pair fixed effects we employed above, we follow Baier and Bergstrand (2007) in this section and first-difference the data. In this specification the time-varying importer/exporter dummies can then be interpreted as the change in multilateral resistance. Regression 13 (Table 3) establishes a benchmark by adding only country-pair fixed effects to a simple OLS specification with time fixed effects. As expected this produces qualitatively identical results as the analog first-differenced setup in regression 14. If anything, the country-pair fixed effect specification produces *larger* WTO trade effects.[14]

This finding has three important implications. First, regression 13 shows that there is sufficient power in the data to produce significant WTO trade effects (if such an effect exists), even after 7138 country-pair fixed effects have been added. Second, either econometric approach produces highly significant and positive WTO membership effects. This is essentially replicating the TGR result, which showed that, given Rose's own coding convention, WTO trade effects were indeed significant in the presence of country-pair fixed effects. This implies our third result, namely that controlling for bilateral heterogeneity with country-pair fixed effects *does not* drive our baseline results in Tables 1a and 2. Instead it is the addition of proper multilateral resistance controls that negate the influence of WTO on bilateral trade flows.

[14] This is because country-pair fixed effects compare all pre-accession periods to all post-accession periods. From regression 24 (Table 4), we know that country-pair fixed effects pick up increases in trade long after accession. First-differences focus only on the difference in trade in the period right after accession compared to the period right before accession.

This important third implication can be confirmed by adding multilateral resistance to the first-differenced specification to obtain regression 15. This is the first-differenced analog of regression 10 (Table 2) and again, WTO trade effects disappear only when multilateral resistance is added. Only a weak, marginally significant WTO effect remains for developing countries. An alternative approach is to allow for AR(1) error terms to control for bilateral heterogeneity. Once multilateral resistance is added, WTO trade effects largely vanish here as well (regression 17). This time, it is a weak industrialized country WTO effect that remains marginally significant at the 10% level. Our other robustness tables below confirm that the WTO effect always vanishes largely because of the introduction of multilateral resistance controls.

4.2. WTO accession dynamics

Table 4 examines whether much of the growth in bilateral trade occurred prior to actual WTO accession, perhaps because countries with high trade growth were more likely to enter the WTO, or because the simple announcement of future WTO accession caused an increase in bilateral trade. The accession dynamics allow us to pick up the pre and post accession changes in bilateral trade, although our results in Table 4 cannot address causality. Again, while country-pair fixed effects leave the WTO effect largely intact, it is eliminated by the introduction of multilateral resistance controls. Once again, our preferred three-way regression 21 shows no significant effect of WTO membership on trade growth either pre-accession or during WTO membership.

4.3. Variations in WTO trade effects over time

There is a possibility that the WTO coefficient's lack of significance is due to large variations in WTO accession experiences over time, where some effects may have been offsetting. Tang and Wei (2009) attempt to finesse this issue by focusing only on specific time periods characterized by a flurry of WTO accessions, for example the 1990s. Table 5 addresses the issue by reporting separate WTO coefficients for each decade. Results are provided for each estimation strategy: OLS, multilateral resistance, country-pair fixed effects, and our preferred three-way fixed effects approach. Our baseline results are confirmed for each decade. For each time period, the three-way fixed effect approach (column 4) eliminates any significant trade effect of WTO membership. Columns 2 and 3 demonstrate that country-pair fixed effects do not suffice to eliminate the WTO effect in most instances. Instead, the WTO effect is again negated by the multilateral resistance controls.

Table 4
Accession dynamics and WTO trade effects. (Only WTO coefficients reported.)

	Dependent variable: bilateral imports			
Regression #	18	19	20	21
Estimation method	OLS	MLR	CPFE	MLR & CPFE
WTO dummy coding	Rose	Rose	Rose	Rose
WTO membership definition	TGR	TGR	TGR	TGR
N	54,389	55,831	54,389	55,831
Adj R^2	0.6419	0.7400	0.7977	0.8370
PTA_{mxt}	0.860***	0.658***	0.735***	0.533***
	(0.051)	(0.062)	(0.051)	(0.049)
WTO/GSP_{mx}, $t-1$	0.038	0.057*	0.054**	0.014
	(0.026)	(0.032)	(0.024)	(0.033)
WTO/GSP_{mx}, t	0.224***	0.058	0.125***	−0.009
	(0.032)	(0.040)	(0.027)	(0.035)
WTO/GSP_{mx}, $[t+1, n)$	0.281***	0.002	0.306***	−0.048
	(0.031)	(0.055)	(0.037)	(0.053)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses. Results are based on Table 2 regression specifications with a single WTO/GSP dummy. WTO/GSP is then disaggregated as follows: "$t-1$" denotes the panel 5 years prior to WTO accession and "$t+1$" indicates 5 years post WTO accession. Full results can be obtained from the authors upon request.

4.4. WTO trade effects for industrialized early joiners

Another potential problem that is associated with the particular timing of WTO accessions relates to industrialized countries. The WTO trade effects of 15 of 23 industrialized countries are omitted from the analysis when country-pair fixed effects are introduced, because these countries joined GATT/WTO before the start of our dataset in 1950.[15] Trade between the United States and the United Kingdom might have been higher on average over the period 1950–2000, because both countries were GATT/WTO members from 1948 onward, but this effect is absorbed by the country-fixed effect between them. Omission of industrialized early joiners in the estimation of the aggregate WTO effect is problematic in view of the terms-of-trade theory, which will be described in more detail in Section 5.3. The theory holds that larger countries with market power have the largest incentives to negotiate tariff reductions upon WTO accession and thus may also reap the largest trade gains.

Regression 22a (Table 5) provides evidence in favor of the theory's prediction. It shows substantial and significant increases in trade for the subset of industrialized country pairs that joined the WTO prior to 1950. The OLS coefficient for these pre-1950 WTO members is significantly larger and the standard error remarkably lower than for the full sample. Once multilateral resistance is introduced, the full sample shows no WTO trade effects, but these pre-1950 WTO members show statistically and economically significant trade creation. This is a notable feat given that multilateral resistance controls eliminated WTO trade effects in virtually all other specifications. Thus, there may be reason to believe that the multilateral trading system indeed boosted trade among these industrialized early joiners. Note, however, that since regression 22a cannot include both country-pair fixed effects and a WTO dummy, a caveat necessarily remains. We cannot identify whether the higher trade among these 15 industrialized nations is indeed due to WTO membership or due to an unobserved trade-enhancing characteristic among them.

5. In search of WTO trade effects: extending the unified framework

Taken together with the strong PTA trade creation from before, these results on industrialized early joiners raise the suspicion that WTO membership may raise trade in more subtle ways than identified by our basic unified framework. In search of more subtle WTO trade effects, this section presents two extensions to our framework. The first extension sets out a gravity model specifically suited to disentangling overlapping PTA and WTO membership. It is then used to investigate whether WTO may have fostered trade regionally, along lines of future PTAs or more generally. We find evidence that WTO membership may underpin regional trade integration among developing countries and in the run-up to PTA formation. The second extension incorporates proxies for the terms-of-trade theory, for which we find support in the data. Our results imply that those countries that had substantial incentives to negotiate tariff reductions during their WTO accession negotiations also exhibit significantly larger and positive WTO trade effects than other members.

5.1. Disentangling WTO and PTA trade effects

Above we assumed that the choice of being in a PTA and/or the WTO is independent. In this section we explore how PTA and WTO membership interact to influence bilateral trade flows.[16] Given that PTA and WTO membership overlaps substantially, it may perhaps be the

[15] See Table A3 in Eicher and Henn (2008).

[16] The idea for this section was suggested to us by Robert Staiger.

Table 5
WTO trade effect across time. (Only WTO coefficients reported.)

Regr.#	Sample period	WTO membership	Estimation method			
			OLS	CPFE	MLR	MLR & CPFE
22	1950–2000	All	0.213***	0.172***	−0.058	−0.067
			(0.270)	(0.027)	(0.049)	(0.048)
22a	1950–2000	Early joiners 1950[a]	0.467***	na	0.240***	na
			(0.081)		(0.073)	
22b	1950–2000	No early joiners 1950	0.197***	0.190***	−0.051	−0.065
			(0.270)	(0.027)	(0.498)	(0.049)
23	1950–60	All	0.401***	0.253***	0.015	−0.020
			(0.050)	(0.076)	(0.088)	(0.110)
24	1960–70	All	0.236***	0.086*	−0.002	0.024
			(0.040)	(0.051)	(0.082)	(0.119)
25	1970–80	All	0.102***	0.078	−0.059	−0.036
			(0.038)	(0.058)	(0.075)	(0.090)
26	1980–90	All	0.019	0.012	−0.137*	−0.087
			(0.039)	(0.057)	(0.078)	(0.094)
27	1990–2000	All	0.319***	0.056	0.077	−0.003
			(0.035)	(0.051)	(0.080)	(0.104)
28	1950–75	All	0.252***	0.096***	−0.034	−0.020
			(0.035)	(0.036)	(0.061)	(0.061)
29	1975–2000	All	0.177***	0.150***	−0.046	−0.021
			(0.031)	(0.036)	(0.063)	(0.067)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses. Results are based on Table 2 regression specifications with a single WTO/GSP dummy. Full results are available from the authors. The dependent variable in all regressions is bilateral imports. WTO membership definition is TGR.
[a] This dummy variable is one for pairs of industrialized countries that joined the WTO in or before 1950 and zero otherwise.

case that we were not able to find WTO trade effects because of the basic gravity model's inability to disentangle the different impacts. To address this concern, this section's extended gravity model explicitly allows bilateral trade flows to be determined by interactions of PTA and WTO membership.

Suppose country-pair, (m, x), consists of WTO members that decide to join a common PTA. We would expect the impact of PTA membership on country m's imports from country x to be smaller than for two non-WTO members that join a common PTA. This statement is true, *ceteris paribus*, if the additional margin of preference implied by the PTA is lower for WTO members (whose MFN tariffs have presumably been negotiated to low levels) than for non-WTO members. At the same time, we would expect that the WTO impact on country m's imports from country x is positive only if the countries are not partners in a free-trade agreement. This statement is true except in the unusual case where the PTA involves a smaller margin of preference than WTO membership.[17] We test these two hypotheses by simplifying the WTO dummies in Eq. (2′) into one aggregate term and by augmenting the empirical framework with a WTO/PTA interaction term.[18]

$$I_{mxt} = a_0 D_{mx} + a_1 D_t + a_2 D_{mt} + a_3 D_{xt} + \beta_1 PTA_{mxt} + \beta_2 PTA_{mxt} \times WTO/GSP^{**}_{mxt} + \beta_3 WTO/GSP^{**}_{mxt} + \beta_5 CurrencyUnion_{mxt} + \beta_6 CurrencyColony_{mxt} + \varepsilon_{mxt}. \quad (3)$$

The interpretation of the parameters in Eq. (3) is as follows: β_1 is the impact of PTAs on trade flows for non-WTO members, β_3 is the WTO impact on trade flows for trading partners that do not belong to the same PTA, and $\beta_1+\beta_2$ represent the impact of PTA membership on trade for WTO members. We expect $\beta_1+\beta_2>0$ as per the discussion above, and by the same reasoning we expect $\beta_1+\beta_2$ to be smaller than β_1, because PTA-induced liberalization should be smaller if WTO liberalization has already been undertaken. Likewise, $\beta_2+\beta_3$ is the effect of WTO membership for trading partners that belong to the same PTA. In this case we expect $\beta_2+\beta_3<0$ since the new WTO bindings create incentives to import from non-PTA partners.

Table 6 reports the results from Eq. (3). While it delivers more structured insights into the mechanics of the WTO trade effects, it nevertheless broadly confirms our previous results. To economize on space, only the preferred three-way fixed effects results are reported, which include controls for multilateral resistance, unobserved country-pair heterogeneity, and time fixed effects. Regression 30 indicates that joining a PTA for WTO non-members doubles trade flows among trading partners ($e^{0.694}-1=100\%$). As expected, a smaller, but highly statistically significant 70% ($=e^{0.522}-1$) increase in trade is generated when WTO members join a PTA. In contrast, joining the WTO for PTA members or non-members has no statistically significant effect on trade flows. This confirms the general pattern we have observed throughout the paper, namely that large trade creation is generated by PTAs while WTO membership seems to have a weak impact on trade flows.

SW, Rose, and our results above lead us to suspect that PTA and WTO trade effects may be heterogeneous, depending on the *types* of countries that join PTAs or the WTO. This hypothesis is examined in regression 31, which disaggregates the WTO variable into industrialized, developing, and GSP-granting importers. While the lack of WTO trade effects is robust for industrial and developing importers, regression 31 does produce the expected decline in trade ($e^{-0.483}-1=-38\%$) for a small subset of trading partners that a) are members in a PTA, b) enter the WTO, and c) contain an industrialized importer extending GSP. The result is intuitive in a world of factor endowment driven trade (relevant exactly for industrialized/developing trading relationships), where the margin of preference implied by the PTA is higher for non-WTO members than for WTO members. WTO accession would lead to a greater reallocation of trade to non-PTA partners. The trade-enhancing effect of entering a PTA is confirmed and statistically significant for industrialized, developing, and industrialized GSP-granting nations. However, the effect is twice as large for developing countries ($110\%=e^{0.740}-1$) as for industrialized ($52\%=e^{0.420}-1$), or GSP granting importers ($55\%=e^{0.438}-1$).

[17] APEC might be a case in point, because it never actually instituted tariff reductions among member countries.
[18] We report results only for the most comprehensive fixed effects and Rose's conventional coding methodology. Results are similar for alternative fixed effects specifications; they can be obtained from the authors upon request.

Table 6
Does PTA membership or distance modify WTO trade effects?

	Dependent variable: bilateral imports					
Regression #	30	31	30a	31a	32	33
N	55,831	55,831	55,831	55,831	55,831	55,831
Adj R^2	0.8370	0.8373	0.8371	0.8374	0.8374	0.8380
Impact of PTAs on trade flows for trading partners that are non-WTO members (β_1)						
PTA_{mxt}	0.694***	0.714***	0.722***	0.753***	0.861***	
	(0.152)	(0.161)	(0.159)	(0.166)	(0.161)	
$PTA\_Industrial_{mxt}$						0.266
						(0.312)
$PTA\_Developing_{mxt}$						1.008***
						(0.177)
Impact of PTA membership on trade flows for trading partners in WTO ($\beta_1+\beta_2$)[a]						
All countries	0.522***		0.573***		0.441***	
	(0.048)		(0.080)		(0.046)	
Industrialized importers		0.420***		0.386***		0.372***
		(0.063)		(0.097)		(0.062)
Developing importers		0.740***		0.823***		0.581***
		(0.082)		(0.102)		(0.080)
Industrialized importers granting GSP		0.438**		0.525**		0.025
		(0.198)		(0.210)		(0.319)
Impact of WTO membership on trade flows for trading partners that share the same PTA ($\beta_2+\beta_3$)[a]						
All countries	−0.239		−0.222		1.081***	
	(0.152)		(0.152)		(0.237)	
Industrialized importers		−0.259		−0.326*		0.771*
		(0.187)		(0.191)		(0.421)
Developing importers		−0.059		−0.028		1.725***
		(0.165)		(0.167)		(0.273)
Industrialized importers granting GSP		−0.483***		−0.436***		0.041
		(0.117)		(0.121)		(0.347)
Impact of WTO on trade flows for trading partners that do not share same PTA (β_3)						
All countries	−0.067		−0.073		1.501***	
	(0.048)		(0.048)		(0.232)	
Industrialized importers		0.036		0.041		0.665**
		(0.072)		(0.074)		(0.300)
Developing importers		−0.085		−0.098		2.152***
		(0.071)		(0.072)		(0.272)
Industrialized importers granting GSP		−0.208***		−0.209***		0.282
		(0.045)		(0.047)		(0.354)
WTO interactions with contemporaneous PTA dummy (β_2)						
$WTO/GSP_{mxt} * PTA_{mxt}$	−0.172		−0.149		−0.420***	
	(0.143)		(0.144)		(0.155)	
$WTO\_Industrial_{mxt} * PTA_{mxt}$		−0.294*		−0.367**		0.106
		(0.168)		(0.177)		(0.312)
$WTO\_Developing_{mxt} * PTA_{mxt}$		0.026		0.070		−0.427**
		(0.146)		(0.148)		(0.168)
$GSP_{mxt} * PTA_{mxt}$		−0.276**		−0.228*		−0.241**
		(0.119)		(0.127)		(0.118)
Impact of WTO membership on trade flows for trading partners that share a PTA in the future ($\beta_3+\beta_4$)[a]						
All countries			−0.003			
			(0.077)			
Industrialized importers				−0.012		
				(0.095)		
Developing importers				0.121		
				(0.109)		
Industrialized importers granting GSP				−0.212**		
				(0.090)		
WTO interaction with distance						
$WTO/GSP_{mxt} * Distance_{mxt}$					−0.191***	
					(0.028)	
$WTO\_Industrial_{mxt} * Distance_{mxt}$						−0.075**
						(0.036)
$WTO\_Developing_{mxt} * Distance_{mxt}$						−0.271***
						(0.032)
$GSP_{mxt} * Distance_{mxt}$						−0.059
						(0.043)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses. All regressions include all appropriate covariates (as in Table 2) and higher order interactions. WTO interactions with the future PTA membership dummy, β_4, are omitted for space reasons. Full results are available upon request from the authors. All regressions feature country-pair fixed effects (CPFE), multilateral resistance (MLR), and TGR WTO membership definition.

[a] Indicates composite coefficients calculated using Delta method (Greene, 2003).

5.2. Are WTO trade effects regional?

The steady reduction of transport costs over the time period covered by our dataset suggests less regionalization of trade. In contrast, market size effects can lower trade costs sufficiently to boost regional but not distant trade (see Baldwin, 2008). WTO accession may thus exert asymmetric effects on proximate/distant trade. Many of these effects are already addressed by multilateral resistance and country-pair fixed effects. However, country-pair fixed effects account only for average bilateral effects over the entire sample period and they might not capture time-varying effects, especially after trading partners enter the WTO. In addition, it may also be the case that WTO membership increased regional trade particularly for countries that eventually form PTAs. In this case, some WTO trade effects may be falsely picked up by PTA coefficients.

To investigate whether WTO trade effects are regional, we add a dummy to Eq. (3) to identify trading partners that are a) currently in the WTO, and b) join a common PTA in the future. The results are presented in regressions 30a and 31a (Table 6). Again, we do not find a significant effect: WTO members' trade with future PTA partners did not increase soon after WTO accession. Alternatively, we also investigate whether WTO membership increases trade among PTA partners-to-be over time. To do so, we split the PTA regressor into dummies that indicate the pre-PTA accession period, $(t-1)$, the PTA accession period, (t), and all subsequent periods, $[t+1, n)$. Results are reported in Table 7, where specification 30b initially omits WTO/PTA interactions. The purpose of regression 30b is to show that the overall PTA effect can be dissected into an 11% ($=e^{0.112}-1$) increase for the pre-PTA period, which rises to 43% ($=e^{0.360}-1$) in the accession period, and 121% ($=e^{0.794}-1$) post-accession. Not only the economic significance, but also the statistical significance increases over time. When WTO/PTA interactions are included again (regression 31b), we find that *only* WTO members experience a significant PTA pre-accession effect of 11% ($=e^{0.106}-1$). The corresponding effect for non-WTO members, albeit similar in magnitude, remains insignificant. Therefore, joint WTO membership may facilitate the formation of PTAs. As expected, post-accession trade effects in PTAs between WTO members remain lower than in PTAs between WTO non-members.[19]

Finally, WTO trade creation may also have fostered regional trade integration irrespective of current or future PTA membership. To explore this possibility, regression 32 (Table 6) examines the effects of distance on the magnitude of WTO trade creation. The basic results are largely similar to those in regression 30, where the impact of PTAs on trade flows was strong for WTO member or non-member countries. Once we allow WTO trade effects to vary with distance, we find two important new effects: the WTO coefficient is now statistically significant at 0.861 and 1.501 for PTA and Non-PTA members, respectively, but these effects are moderated by distance. The interaction between distance and WTO produces a negative significant coefficient of −0.191, indicating that the positive effect of WTO membership on trade declines with distance. We can trace the WTO trade effect for PTA members and non-members in Figs. 1 and 2. For PTA members, Fig. 1a indicates an insignificant WTO effect throughout, while WTO members that do not belong to a PTA exhibit positive and statistically significant WTO trade effects up to 1500 miles. These effects turn negative and statistically significant beyond 4000 miles. Thus there is strong evidence that WTO membership creates proximate trade incentives for non-PTA countries, but at least partially at the expense of distant trade. Our insignificant coefficients

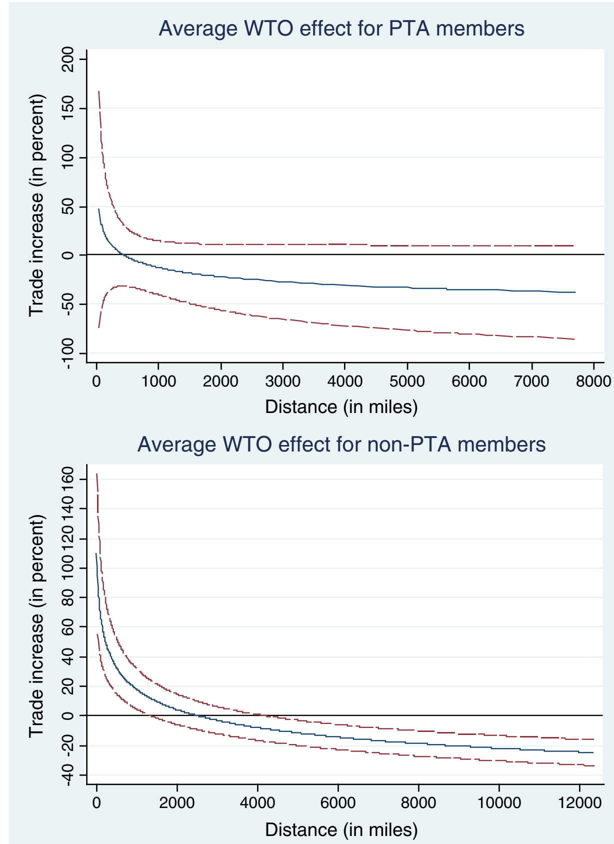

Fig. 1. Variation with distance of overall WTO effect. The figures are based on composite coefficient estimates derived from Table 6. Dashed lines represent 95% confidence bands.

Table 7
WTO influence on PTA accession dynamics.

	Dependent variable: bilateral imports			
Regression #	30b	31b	30c	31c
Estimation method	MLR & CPFE	MLR & CPFE	First-differenced MLR	First-differenced MLR
N	55,831	55,831	40,925	40,925
Adj R^2	0.8373	0.8373	0.2773	0.2774
Impact of PTAs on trade flows for trading partners that are not WTO members (β_1)				
PTA_{mx}, $t-1$	0.112**	0.193	0.099***	0.179
	(0.046)	(0.151)	(0.037)	(0.139)
PTA_{mx}, t	0.360***	0.309*	0.289***	0.012
	(0.053)	(0.184)	(0.049)	(0.181)
PTA_{mx}, $[t+1, n)$	0.794***	1.118***	0.515***	0.403***
	(0.067)	(0.150)	(0.061)	(0.157)
Impact of WTO on trade flows for trading partners that do not share same PTA (β_3)				
WTO/GSP_{mxt}	−0.061	−0.060	0.077*	0.078*
	(0.048)	(0.048)	(0.040)	(0.040)
Impact of PTA membership on trade flows for trading partners that are also WTO members ($\beta_1+\beta_2$)[a]				
$t-1$		0.106**		0.094**
		(0.045)		(0.037)
t		0.369***		0.313***
		(0.054)		(0.049)
$[t+1, n)$		0.771***		0.527***
		(0.065)		(0.060)
Impact of WTO membership on trade flows for trading partners that share the same PTA ($\beta_2+\beta_3$)[a]				
$t-1$		−0.147		−0.006
		(0.159)		(0.140)
t		0.001		0.379**
		(0.189)		(0.177)
$[t+1, n)$		−0.407***		0.202
		(0.143)		(0.147)
WTO interactions with contemporaneous PTA dynamics (β_2)				
$WTO/GSP_{mx}*PTA_{mx}$, $t-1$		−0.087		−0.085
		(0.149)		(0.135)
$WTO/GSP_{mx}*PTA_{mx}$, t		0.060		0.301*
		(0.180)		(0.171)
$WTO/GSP_{mx}*PTA_{mx}$, $[t+1, n)$		−0.347***		0.124
		(0.134)		(0.141)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parentheses.
[a] Indicates composite coefficients were calculated using the Delta method (Greene, 2003). Regressions based on specifications in Table 2. The regressions also include all appropriate higher order interactions involving PTA/WTO&GSP/distance. Full results are available upon request from the authors. All regressions feature country-pair fixed effects (CPFE), multilateral resistance (MLR), and TGR WTO membership definition.

[19] In regression 31b, the 206 percent ($=e^{1.118}-1$) post-accession PTA effect for WTO non-members is about twice as large as the 116 % ($=e^{0.771}-1$) increase in PTA bilateral trade flows for WTO members.

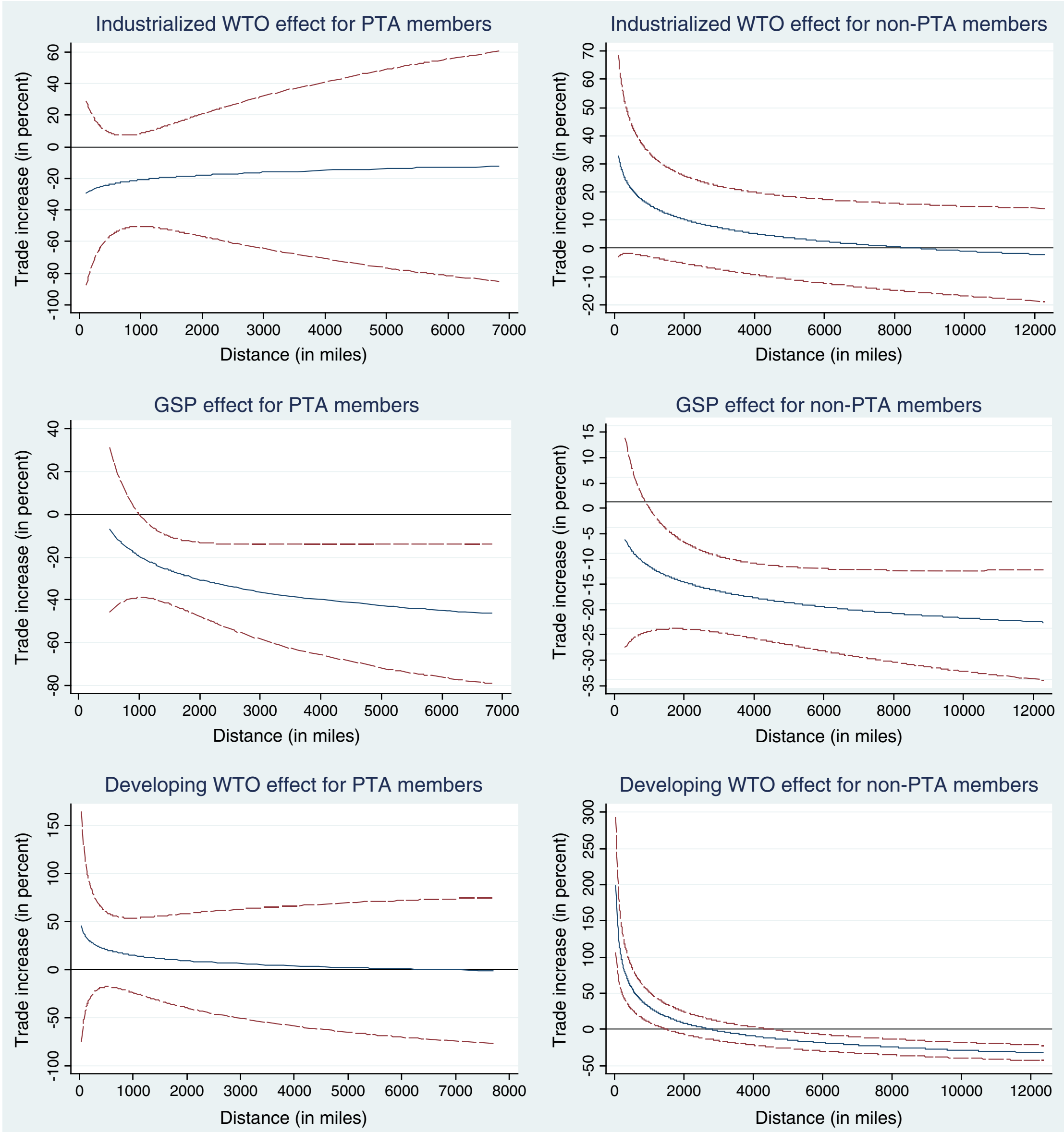

Fig. 2. Variation with distance of country group specific WTO effects. The figures are based on composite coefficient estimates derived from Table 6. Dashed lines represent 95% confidence bands.

in previous regressions may then be explained by offsetting positive (negative) effects for proximate (distant) trading partners.

In regression 33 we again disaggregate the WTO variable into individual WTO trade effects for industrialized GSP and developing importers (Fig. 2). The figures generally do not show positive and statistically significant WTO trade effects. The exception is the case of developing country importers that are not PTA members (Fig. 2, bottom right). For these countries, WTO accession generates positive and significant trade benefits with proximate partners (<1750 miles) at the cost of significant negative effects with distant trading partners (>4250 miles).

5.3. Variations in WTO trade effects according to terms-of-trade theory

Strict economic interpretations of Rose's, SW's, or our findings can be difficult at times because the basic gravity model does not provide a specific theoretical framework to analyze WTO trade effects. Bagwell and Staiger (2010) suggest that the absence of theoretical guidance, which specifically addresses WTO effects on trade, calls into question whether the Rose/SW gravity approach can claim to provide a comprehensive assessment of WTO trade effects. In this section we augment the gravity model to proxy for effects suggested by the terms-of-trade theory, which has been specifically designed to analyze the effects of WTO membership.

Bagwell and Staiger (1999) put forth a terms-of-trade-theory of GATT which finds WTO membership particularly useful for governments that seek to escape a terms-of-trade-driven prisoners' dilemma. The notion is that large countries with market power and the ability to influence world prices will do so through trade barriers that move the terms of trade in their favor. The resulting retaliation from other large countries then generates the terms-of-trade prisoner's dilemma. Since larger countries have greater incentives to attempt to change the terms of trade in their favor, terms-of-trade theory suggests that the magnitude of negotiated tariff reductions prior to WTO accession is larger for such countries. Larger tariff reductions then imply greater post-accession trade gains.

Table 8
Using terms of trade theory to identify WTO trade effects. WTO induced trade gains by country-import rank in WTO accession year.

	Dependent variable: bilateral imports						
Regression #	34	35	36	37	38	39	40
Estimation method	MLR & CPFE	MLR & CPFE	MLR & CPFE	MLR & CPFE	MLR & CPFE	MLR & CPFE	MLR & CPFE
WTO dummy coding	Rose	Rose	Rose	Rose	Rose	Rose	Rose
WTO membership definition	*De jure*	*De jure*	*De jure*	*De jure*	*De jure*	*De jure*	*De jure*
N	55,831	55,831	55,831	55,831	55,831	55,831	55,831
Adj R^2	0.8383	0.8383	0.8383	0.8383	0.8383	0.8383	0.8383
WTO/GSP_{mxt}	−0.001	−0.017	−0.019	−0.018	−0.028	−0.014	−0.007
	(0.053)	(0.048)	(0.046)	(0.045)	(0.043)	(0.041)	(0.040)
WTO/GSP_{mxt} * imports_66%[a]	0.012						
	(0.058)						
WTO/GSP_{mxt} * imports_70%[a]		0.051					
		(0.055)					
WTO/GSP_{mxt} * imports_75%[a]			0.061				
			(0.055)				
WTO/GSP_{mxt} * imports_80%[a]				0.067			
				(0.055)			
WTO/GSP_{mxt} * imports_85%[a]					0.111**		
					(0.056)		
WTO/GSP_{mxt} * imports_90%[a]						0.121**	
						(0.061)	
WTO/GSP_{mxt} * imports_95%[a]							0.162**
							(0.074)
$PTAs_{mxt}$ *(individual PTAs?)*	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Composite of both WTO effects[b]	0.014	0.034	0.042	0.049	0.084	0.107*	0.155**
	(0.044)	(0.047)	(0.048)	(.050)	(0.052)	(0.060)	(0.074)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parenthesis. Regressions based on Table 2 specifications. Full results are available upon request from the authors.
[a] *Imports_x%* is a dummy that identifies countries whose $imports_{jt}/(world\_imports_t)$ exceed the x^{th} percentile (where *t* is the year of country *j*'s accession to the GATT/WTO). The percentile rankings are generated as follows: First, we obtain the ratio of country $imports_{jt}$ over $world\_imports_t$ from the IMF Direction of Trade Statistics. Second, countries acceding WTO in year *t* are then percentile-ranked relative to all other countries in year *t*. (Results are just about identical when establish a simple percentile ranking of all accession year ratios.) Data availability required that some WTO accession countries had to be ranked based on a post WTO accession year data: Bangladesh 1973 (1972), Bermuda 1958 (1948), Comoros 1969 (1948), Dem. Rep. of Congo 1972 (1971), Kuwait 1973 (1963), and Seychelles 1970 (1963), where the formal accession year is provided in the parentheses. Antigua and Barbuda and South Africa had to be omitted due to missing data.
[b] Composite coefficients calculated using the Delta method (Greene, 2003).

The terms of trade approach has been taken to the data by Broda et al. (2008), who focus on market power, and by Bagwell and Staiger (forthcoming) import volumes. Bagwell and Staiger show that the terms-of-trade theory implies that negotiated tariff reductions at WTO accession increase (i) the larger the country's ability to alter foreign exporter prices, (ii) the larger the country's pre-negotiation import volume, and (iii) the smaller the rate at which the costs of protection-induced domestic distortions rise as tariffs rise. Using data on WTO accession negotiations in a panel of 16 countries from 1995 to 2005, Bagwell and Staiger (forthcoming) show that terms-of-trade theory is consistent with observed patterns of negotiated tariff concessions. Specifically, the authors derive an econometric model that suggests the international cost-shifting incentives increase with a country's import volume. Accordingly, the larger a country's Nash import volume, the greater should be its incentive to negotiate tariff cut at WTO accession.

The key insight from Bagwell and Staiger (forthcoming) is that country characteristics affect accession negotiations, tariff concessions, and hence the subsequent trade gains that can be generated by WTO membership. Their data clearly show that for recent WTO accession countries, greater import volumes were associated with larger tariff cuts, which should then generate larger trade gains. Following Bagwell and Staiger (2010), we focus on import volumes as a proxy for the gains a country may reap from liberalization. Specifically, we attempt to discern whether countries with higher import volumes in their WTO accession year possessed greater terms-of-trade incentives to negotiate tariff reductions, which then generated larger subsequent trade gains. Table 8 reports that WTO trade effects indeed increase with trade volumes at accession, suggesting that countries with larger import volumes (relative to world trade) negotiated larger tariff reductions at accession.

Results in Table 8 are obtained by augmenting a regression specified by Eq. (2′) with an additional dummy that indicates whether a country ranks above a specific import volume threshold at the time of WTO accession. As we vary the threshold from the 66th to the 95th percentiles, we find that those countries which rank below the 66th percentile in import volumes never experience positive WTO trade effects. WTO trade effects turn positive and significantly different from the rest of the sample for countries that rank above the 85th percentile in imports relative to world trade. The trade effects increase in magnitude and significance until the 90th and 95th percentiles, at which point not only the marginal, but also the aggregate WTO trade effects are positive and significant. The results imply a 17% ($=e^{0.162}-1$) trade increase due to WTO membership for countries in the highest import percentiles. This WTO trade effect is not as large as some of the PTA trade effects we found in Table 2. When comparing magnitudes it must be considered, however, that PTA trade effects *include* industrialized countries, while the WTO trade effects in Table 8 *exclude* trade effects generated by industrialized country pairs that joined prior to 1950.[20] Since exactly this set of countries had presumably the greatest terms of trade externalities, it is all the more remarkable that we can establish the pattern of WTO trade gains suggested by terms-of-trade theory among the set of remaining countries.

Our result complements Bagwell and Staiger (2010, Fig. 1b), who find in a sample of 16 countries (from 1995 to 2000) that those countries whose import quantities exceeded the 80th percentile agreed to greater than average tariff concessions in their WTO accession negotiations, and that the tariff concessions increased dramatically for countries whose import quantities exceeded the 90th percentile. We observe a similar effect in Table 8 in terms of magnitudes of WTO trade gains in our sample of 177 countries over 50 years. However, our effect relates to post-WTO accession import gains rather than tariff concessions at WTO accession. Presumably these import gains were generated by correspondingly larger tariff concessions at accession.

[20] See the discussion in Section 4.4 of Table 5, regressions 22a and 22b.

6. Conclusion

This paper reexamines the effects of WTO membership on bilateral trade flows. First we show that a number of previous approaches can be combined into one unified framework. This framework controls comprehensively for omitted variable bias in three dimensions: individual PTA effects, multilateral resistance, and unobserved bilateral heterogeneity. Our results show that all previous approaches (Rose, 2004, 2005; Subramanian and Wei, 2007; and Tomz et al., 2007) produce the result that WTO membership does not generate statistically significant trade effects. The analysis highlights that the diverging and conflicting results regarding WTO trade effects in the literature were generated by omitted variable bias.

In contrast, we find that PTAs create trade strongly, but unevenly across individual agreements. The magnitude of our individual PTA estimates resolves a number of empirical puzzles. Most notably, the non-tariff reducing APEC is shown to exert comparatively little trade impact, and the strongly tariff-reducing EU is shown to be trade creating. Trade theory motivates the inclusion of comprehensive multilateral resistance controls to pick up variations in relative trade costs. These controls are shown to be insufficient to generate unbiased estimates of trade agreements' impacts on trade flows. Of crucial importance are also country-pair fixed effects that control for unobserved bilateral characteristics.

In two extensions of the gravity model that account for specific ways in which theory suggests WTO trade creation, we find positive and significant trade effects. Our first extension disentangles overlapping WTO and PTA membership effects. We find that WTO membership increases trade effects just before PTA accession. In addition, WTO membership fosters regional trade integration among developing country members at the expense of more distant trade. Our second extension augments the gravity model with proxies for the WTO terms-of-trade theory. Here we find that countries with greater incentives to bargain for tariff reductions during WTO accession negotiations exhibit positive and significant WTO trade effects.

Appendix A

Table A1
Observations with changes in WTO status.

Year	Observations with changes in WTO membership	Total number of observations
1950	na	1502
1955	244	1989
1960	109	3166
1965	428	3995
1970	128	4738
1975	717	5688
1980	372	5968
1985	425	6316
1990	530	6715
1995	400	7674
2000	481	8080
Total	3834	55,831

Table A2
OLS results.

Regression #	A1	A2	A3	A4	A5	A6	A7	A8	A9
Dependent variable	Avg. trade	Imports	Imports	Avg. trade	Imports	Imports	Imports	Imports	Imports
WTO dummy coding	Rose	Rose	Rose	SW	SW	SW	Rose	Rose	SW
WTO membership definition	*De jure*	*De jure*	TGR	*De jure*	*De jure*	TGR	*De jure*	TGR	TGR
N	54,389	54,389	54,389	54,389	54,389	54,389	54,389	54,389	54,389
Adj R^2	0.6946	0.6398	0.6411	0.6946	0.6399	0.6410	0.6422	0.6431	0.6436
PTA_{mxt}	0.801***	0.877***	0.865***	0.801***	0.863***	0.855***	0.840***	0.833***	1.169***
	(0.052)	(0.050)	(0.051)	(0.052)	(0.050)	(0.050)	(0.051)	(0.051)	(0.056)
WTO/GSP	0.044*	0.103***	0.213***	0.045*	0.129***	0.243***			
	(0.026)	(0.025)	(0.027)	(0.023)	(0.024)	(0.027)			
GSP_{mxt}							0.065*	0.053	0.661***
							(0.036)	(0.036)	(0.053)
WTO_Ind_{mxt}							0.278***	0.379***	1.034***
							(0.036)	(0.038)	(0.054)
WTO_Dev_{mxt}							−0.089***	0.044	0.006
							(0.027)	(0.029)	(0.036)

Notes: *, **, *** are 10%, 5%, 1% significance levels. Standard errors in parenthesis. Fixed effect coefficients are suppressed. All regressions include time fixed effects. Due to SW's coding convention, PTA and GSP effects are calculated according to footnote 14. Average Trade data is obtained from Rose (2004). As specified in the text, these regressions also include Log of importer GDP, Log of exporter GDP, Log of importer GDP per capita, Log of exporter GDP per capita, importer is Landlocked, exporter is Landlocked, importer is Island, exporter is Island, log of importer land mass, log of exporter land mass, currency union, current colony, ever colony, common colonizer, common language, same nation, common border, log of distance. Full results are available upon request from the authors.

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