On the Robustness of the Trade-Inducing Effects of Trade Agreements and Currency Unions

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Abstract
Regional trade agreements (RTAs) and currency unions (CUs) share the characteristic of being potentially endogenous trade cost proxies in gravity equations. In both cases, this problem is magnified by the paucity of reliable instruments. Instead of resorting to the oft-employed alternative of panel data in order to address selection on just the time-invariant unobservables, this paper provides the first empirical analysis of the extent to which the positive association between CU or RTA membership and bilateral trade can be considered causal. Despite not identifying point estimates, striking results are obtained. Although most cross-sections find both RTAs and CUs to be associated with increased bilateral trade, the evidence in favor of a causal effect is strong only for CUs, and is also witnessed at the extensive margin of trade.

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1 Introduction

Although Anderson and van Wincoop (2004, p. 710) allude to a "list of observable arguments ... which have been used in the trade cost function" of gravity equations, analyses pertaining to the effects of regional trade agreements (RTAs) and currency unions (CUs) have perhaps received the most attention.[1] While Baier and Bergstrand (2009b, p. 78) consider "the most common usage" of the gravity model in analyzing the role of trade agreements, the number of studies estimating the effects of CUs on members' bilateral trade are not few either. In fact, Rose and Stanley (2005, p. 359) consider the "empirical literature on the trade consequences of currency unions" to be "rich" despite its "youth."

Regardless of the extant literature, the relevance of analyzing RTAs and CUs cannot be emphasized enough. While the notification of more than fifty RTAs to the World Trade Organization (WTO) between January 2005 and December 2006 suggests their recent proliferation, ongoing negotiations of such agreements indicate their unabated rise in years to come. In fact, if all such agreements currently under negotiation and proposal are implemented, then one would be looking at over four hundred RTAs by 2010 (Fiorentino et al., 2007). Hence, policy issues associated with trade agreements are relevant for some time to come. However, the importance of trade agreements does not undermine the significance of CUs in the current trading climate. Recent episodes of dollarization in Ecuador, El Salvador and Guatemala, adoption of the euro, contemplation by a number of West African states to form a CU, and the intention of "six oil-producing countries ... to form a currency union by 2010," urge Barro and Tenreyro (2007, p. 2) to regard studies pertaining to the "economic effects" of CUs as "imperative."

Apart from being relevant from a trade policy perspective, RTAs and CUs also share the common characteristic of being potentially endogenous regressors in a gravity equation. In fact, while discussing trade cost proxies which may not be exogenous, Anderson and van Wincoop (2004, p. 706) provide "membership in a currency union or regional trade agreement" as examples. In other words, the point estimates of RTA and CU coefficients are clearly susceptible to bias from selection on unobservables. To make matters worse, the sign of the bias is also ambiguous, as discussed by Barro and Tenreyro (2007) in the context of CUs, and Baier and Bergstrand (2009a) with respect to trade agreements. Although instrumental variables (IV) could be used in theory to estimate the causal effects of RTAs or CUs, the selection issue is further plagued by the lack of appropriate instruments for both. While Frankel and Rose (2002, p. 459) suggest that "plausible instrumental variables for currency union membership do not appear to exist in practice," Baier and Bergstrand (2007, 2009a) also rue the lack of reliable instruments

[1] The trade agreements database of the World Trade Organization, a source consulted extensively for this study, classifies RTAs according to their scale of integration. Since this paper does not attempt to distinguish between such regimes, the generic terminology of RTAs is used.

in the context of trade agreements. As a result, point estimates from gravity equations seem all the more questionable. Incidentally, both Rose (2000, p. 17) and Frankel and Rose (2002, p. 461) recommend against taking the point estimates of the CU effect "too literally." In the context of trade agreements, Ghosh and Yamarik (2004, p. 389) take an even stronger stand by referring to a point estimate as a "quantitative magnitude for ignorance."

In light of this, the current study contributes to the literature by acknowledging the unreliability of point estimates, and instead assessing the sensitivity of RTA and CU coefficient estimates, obtained under exogeneity, to non-random selection. While most authors in the literature have resorted to panel data in order to control for selection on just the time-invariant unobservables, the approach in this paper is novel.[2] Given recent concerns over the evolution of trade at two margins, i.e., intensive and extensive, the selection issue is also examined at the latter. In addition, this paper also provides the first study to examine, in relative detail, the combined effect of RTAs and CUs by considering country pairs party to both as the treatment group. Intriguingly, Baur and Winschel (2009) is the only study, to the author's knowledge, to consider a combined effect of CU and free trade agreement membership. Using a Bayes network approach, Baur and Winschel (2009) conclude that the effect of any joint policy should be examined by modeling the policy combination as a separate regime. In other words, the combined policy effect should not be deduced as the sum of the individual policies' partial effects.

Using theoretically consistent gravity models and data that include country pairs with zero trade, striking results are obtained. Although most cross-sections find both RTAs and CUs to be associated with increased bilateral trade, the evidence in favor of a causal effect is strong only for CUs. On the contrary, most positive and significant RTA coefficients, estimated under the assumption of exogeneity, can be explained by even modest levels of positive selection under a set of reasonable assumptions. Accordingly, concerns over selection bias are well-founded in the context of trade agreements. However, on the basis of obtained evidence, Rose's (2001, p. 456) conjecture that "non-random selection" into CUs is of "academic interest, but unimportant in practice," is difficult to deny except as it relates to the *magnitude* of the CU effect. Interestingly, a similar pattern of selection is also witnessed at the extensive margin. Moreover, results from the paper suggest that in order to determine the trade-inducing effect of membership into both CUs and RTAs, the policy combination should be modeled as a separate regime.

The remainder of the paper is organized as follows. Section 2 reviews the literature. Section 3 describes the empirical methodology. Section 4 discusses the data. Section 5 presents the results, while Section 6 concludes.

[2] While Altonji et al. (2005) and Millimet et al. (forthcoming) adopt a similar strategy in other contexts, Chang and Lee (2008) is the only study in the trade literature to pursue a related approach.

2 Literature Review

Regional Trade Agreements Before proceeding to review even part of the existing gravity literature on the effects of trade agreements, it is important to recognize the types of such agreements. While a preferential trade agreement is essentially an arrangement among countries whereby members engage in trade at reduced tariff rates, they can be classified as partial or total with respect to the extent of duty reduction or commodity coverage. In fact, the total agreements can be further categorized on the basis of their level of integration.[3] Accordingly, the gravity literature has addressed a host of policy issues by analyzing various kinds of trade agreements. While Baier and Bergstrand (2007, 2009a) focus on free trade agreements, others such as Magee (2003) and Ghosh and Yamarik (2004) analyze the effects of RTAs in general. Adopting an entirely different stance, Carrère (2006) estimates the partial effects of seven specific RTAs from a single gravity equation. More recently, Roy (forthcoming) and Vicard (2009, p. 182) assess whether the "depth" or form of agreements matter with respect to their impact on members' bilateral trade by including separate controls for RTA types. Thus, given such differences in specification across studies, it is probably best to refrain from comparing the various point estimates. Perhaps, in order to circumvent this issue for conducting a meta-analysis, Cipollina and Salvatici (2006, p. 2) classify RTAs only into "reciprocal ... and nonreciprocal" agreements while focusing on the former.

However, irrespective of the differences in defining the RTA variable(s), the overriding concern in the literature is that of selection into trade agreements and potential bias stemming from it. This concern attains greater relevance, given the lack of consensus regarding the direction of bias. While Baier and Bergstrand (2009a, p. 65) opine that omitted variables might bias the trade agreement "coefficient estimates up or down," Magee (2003, p. 1) suggests the possibility of positive selection by alluding to the "natural trading partner hypothesis." Intriguingly, despite the large volume of the RTA literature dating back to Tinbergen (1962), attempts to address this issue have only begun recently. In fact, Magee (2003, p. 14) uses a simultaneous equations model in providing "one of the first estimates" of the effect of preferential trade agreements while attempting to address the issue of endogeneity. Although Magee (2003) relies on IV and finds countries' volume of bilateral trade to increase the possibility of their entering into a trade agreement, the evidence on the impact of such agreements on trade is less clear. Moreover, the quality of instruments used, such as GDP similarities between two countries, or differences in their relative factor endowments, is clearly suspect. Interestingly, despite being questionable, the instruments in Magee (2003) can still be credited for highlighting the fact that most political and economic variables fail to satisfy the

[3] In this context, it should be noted that Frankel (1997) also categorizes partial agreements as reciprocal and nonreciprocal. Frankel (1997, p. 13) considers one-way concessions to have been "widely tolerated" by the General Agreement on Tariffs and Trade (GATT).

required exclusion restriction, and hence, should not be used as instruments.

Given the issue of non-random selection and the paucity of reliable instruments, Baier and Bergstrand (2007) recommend the use of panel data in order to at least control for selection on the basis of time-invariant unobservables. Accordingly, Kandogan (2008) and Magee (2008), among others, have resorted to the use of panel fixed effects. However, such results should also be interpreted with caution given the tension between the time dimension of the data and the assumption of time-invariant unobservables. Moreover, even the use of panel data does not quell the ambiguity regarding the direction of bias. While Baier and Bergstrand (2007) motivate the possibility of negative selection on the basis of time-invariant unobservables, according to Magee (2008, p. 361), the bilateral fixed effects "solve the problem" of positive selection.

More recently, although Henderson and Millimet (2008) find concerns over the gravity model's parametric form unwarranted, Egger et al. (2008) and Baier and Bergstrand (2009a) employ matching techniques. While Egger et al. (2008) continue to avoid potential selection due to time-invariant unobservables by using difference-in-differences matching, Baier and Bergstrand (2009a, p. 64) revert to the world of selection on observables by alluding to "many" who "have argued that the selection bias on observables may well dominate that on unobservables." However, the fact that the gravity model does not specify which observables to include in the trade cost function is well-known. In such a scenario, this study contributes to the literature by controlling for a fairly representative set of observables, and then assessing the robustness of RTA coefficient estimates to selection on unobservables.

Currency Unions Unlike the case of trade agreements, the literature analyzing the effects of CUs on members' trade is relatively recent and can be traced only as far back as Rose (2000). Since Rose's (2000) analysis found CUs to more than triple bilateral trade, according to Rose and Stanley (2005, p. 348), "nearly everyone" considered the effect to be "implausibly large." In fact, Rose and Stanley (2005, p. 348) go on to hold the magnitude of the point estimate responsible for having motivated nearly "all the subsequent research in this area." Interestingly, the subsequent studies have mostly confirmed a substantial CU effect. In other words, the positive and sizeable CU effect has proven to be robust to the use of matching techniques in studies such as Rose (2001), panel fixed effects as in Glick and Rose (2002), and IV as in Rose (2000).[4] Using pre-World War I data, Flandreau and Maurel (2001) and López-Córdova and Meissner (2003) also find the trade promoting effect of monetary unions to be non-negligible. In short, the CU effect has mostly been described as enormous or some synonym thereof (see, e.g., Rose and van Wincoop, 2001, Barr et al., 2003, and Rose and Engel, 2002).

[4] Persson (2001, p. 446) employs matching techniques to obtain a "very sizeable" CU effect. However, the level of significance of the CU coefficient estimate is sensitive to the matching algorithm in Persson (2001).

In light of these findings and given the fact that it is "unclear" why a CU raises "trade levels so much," concerns over endogeneity seem warranted (Anderson and van Wincoop, 2004, p. 710). In fact, according to Frankel (2008), attempts to address the selection issue should be considered priority. While Barro and Tenreyro (2007) discuss the possibilities of both positive and negative selection, the positive and statistically significant CU effects across most studies has mainly raised concerns over a positive selection bias. Perhaps, Baldwin (2006, p. 35) best expresses this by suspecting that "reverse causality ... must be biasing the Rose effect upward." However, while a number of authors including Baldwin (2006) rue the lack of appropriate instruments for the CU dummy, others such as Frankel and Rose (2002) and Glick and Rose (2002) simply believe the results to be robust to endogeneity. In order to quell the debate, Barro and Tenreyro (2007) attempt a new IV approach. The IV estimates suggest that if anything, the "OLS results underestimate" the CU effect (Barro and Tenreyro, 2007, p. 12).[5]

However, in keeping with Rose (2000), Rose and van Wincoop (2001), and Frankel and Rose (2002), Barro and Tenreyro (2007, p. 16) also consider the point estimates to be "extremely large." In other words, while the literature seems to concur on a substantial trade promoting effect of CUs, there appears to be a general reluctance to precisely quantify the effects. Again, in this light, the sensitivity analysis undertaken in this paper is relevant.

Before discussing the empirical methodology, it is essential to note that all studies examining the effects of CUs in general, have focused on large datasets with a number of small and developing countries. As a result, most studies in this literature including Rose (2000), Rose and Engel (2002), and Barro and Tenreyro (2007) caution against generalizing or extrapolating the findings to more developed countries, or specifically the European Economic and Monetary Union (EMU). In fact, there exists a separate strand of the CU literature focusing solely on the euro's effect on members' trade.[6] A characteristic feature of this branch is the finding of a typically lower trade promoting effect of the euro as witnessed by Chintrakarn (2008) and Millimet and Tchernis (2009), among others. Perhaps Frankel (2008, p. 3) best summarizes this by stating that "the central tendencies" of the euro estimates "seems to be an effect in the first few years on the order of 10-15%." The meta-analysis in Havránek (2009, p. 7) lends further support to this by finding the effects of the euro and other CUs to be "immensely different." Accordingly, for the relevant

[5] According to Barro and Tenreyro (2007), two countries may share a common currency due to their independent decisions to maintain parity with a third anchor currency. As a result, the instrument for the CU dummy is obtained as the joint probability that a country pair adopts the same anchor currency. However, unobservable historical and political ties between two countries may not only affect their bilateral trade, but may also lead to their choice of a common anchor thereby rendering the validity of the instrument doubtful. In addition, the lack of theoretically motivated multilateral resistance terms make the results unreliable.

[6] Chintrakarn (2008) provides a concise review of this literature.

cross-section, the CU effect is analyzed in this study with and without the euro countries as part of the treatment group.

Although the euro's role in encouraging bilateral trade is examined, the focus is essentially on assessing the extent to which CUs and RTAs cause trade, if any. This paper also contributes to the literature by providing the first study, to the author's knowledge, to examine the combined effect of CU and RTA membership in relative detail.

3 Empirical Methodology

3.1 Baseline Approach

In keeping with most of the empirical literature on RTAs and CUs, gravity models are estimated in logs.[7] However, in order to examine selection issues at the extensive margin, a bivariate probit model is also employed. In all cases, controls for Anderson and van Wincoop's (2003) theoretically motivated multilateral resistance (MR) terms are included. Although country-year-specific dummy variables control for the MR terms, Baier and Bergstrand (2009b, p. 78) suggest an alternative which provides "virtually identical coefficient estimates" to the fixed effects approach. It involves a first-order Taylor expansion of the MR terms.[8] In fact, referring to it as "'bonus vetus'" ordinary least squares (BVOLS), Baier and Bergstrand (2009b, p. 78) also recommend the approach for probit models. Hence, given its computational ease, this method is adopted for the bivariate probit. The use of BVOLS instead of the country fixed effects also avoids the potential incidental parameters problem (see, however, Egger et al., 2009). Moreover, regardless of the incidental parameters problem, convergence of the bivariate probits with country fixed effects proved very difficult for majority of the cross-sections. For completeness, the log specification is subjected to both methods.

Log Model Using Country Fixed Effects The (cross-section) specification is given by

$$\ln(T_{ij}+1) = \tau D_{ij} + X_{ij}\beta + \eta_i + \eta_j + \varepsilon_{ij}. \tag{1}$$

Here, T_{ij} is the value of imports of country i from country j; D_{ij} is a dummy variable taking the value one if i and j belong to the same RTA, or CU, or both, depending on the treatment under consideration,

[7] Note that Santos Silva and Tenreyro (2006) caution against the log model because heteroskedasticity in levels can induce correlation between the covariates and the error term in logs even if the covariates are exogenous in the levels model. Since the objective here is to assess sensitivity to correlation between covariates and the error term in the log model, this issue is not particularly relevant. In addition, estimation in logs enables comparison to the literature.

[8] In Baier and Bergstrand (2009b), the Taylor expansion is centred around a world of symmetric trade costs, i.e., $t_{ij} = t$ for all country pairs ij. Here t denotes trade costs.

and zero otherwise; and X_{ij} is a vector of observable attributes of country-pair ij (including an intercept). The following covariates are included in X: (log) distance between i and j, a binary variable assuming the value unity if i and j share a land border, a dummy variable taking the value one if i and j share a common language, a dummy variable taking the value one if i and j share a common religion, a binary variable taking the value unity if i has ever been a colony of j, a binary variable taking the value unity if i has ever been a colonizer of j, a dummy variable taking the value one if i is considered to be a colony of j, a dummy variable taking the value one if i is considered to be a colonizer of j, a binary variable assuming the value unity if i and j were ever colonized by the same colonizer, a measure of the intensity of military conflict between i and j, a dummy variable taking the value one if i and j are in a formal alliance, a dummy variable taking the value one if both i and j are WTO members, a dummy variable taking the value one if either i or j is a WTO member, a dummy variable taking the value 1 if i offers preferences to j under the Generalized System of Preferences, and a dummy variable taking the value 1 if j offers preferences to i under the Generalized System of Preferences. η_i and η_j are country-specific dummies.[9] The pairwise unobservables are denoted by ε_{ij} and capture all remaining factors affecting bilateral trade.

Log Model Using BVOLS In this case, the (cross-section) specification is given by

$$\ln(T_{ij}+1) = \tau D_{ij} + \pi_1 \ln GDP_i + \pi_2 \ln GDP_j + X_{ij}\beta + \widetilde{\tau}\mathrm{MR}D_{ij} + \mathrm{MR}X_{ij}\widetilde{\beta} + \varepsilon_{ij}. \tag{2}$$

Due to the use of BVOLS instead of country-specific dummies, the variables $\ln GDP_i$ and $\ln GDP_j$ appear in (2). Moreover, X now includes (log) geographical areas of i and j in addition to the variables listed previously. Here, GDP_i (GDP_j) is the real gross domestic product (GDP) of country i (j).[10]

Before proceeding, it is important to note that all variables in (2) except the GDPs depict trade costs. As a result, all such variables enter the MR terms. However, for brevity, the MR term from Baier and Bergstrand's (2009b) BVOLS method is only defined for a representative trade cost variable, t_{ij}, faced by country pair ij. In other words, each trade cost variable (t_{ij}) in (2) has a corresponding MR term, given by

$$\mathrm{MR}t_{ij} = \sum_{k=1}^{N}\theta_k t_{ik} + \sum_{m=1}^{N}\theta_m t_{mj} - \sum_{k=1}^{N}\sum_{m=1}^{N}\theta_k\theta_m t_{km} \tag{3}$$

where N is the number of countries, and $\theta_k\,(\theta_m)$ is the GDP share of country $k\,(m)$.[11] Also, the theory in Baier and Bergstrand (2009b) restricts the coefficients on each t_{ij} and the corresponding $\mathrm{MR}t_{ij}$ to be

[9] The country dummies are usually used to control for country-specific unobservables that do not vary across trading partners as well as the MR terms. In this case, they also capture the impact of GDP and area.

[10] The US consumer price index is used to express GDP in 1995 dollars.

[11] Here, area of a country is also considered to reflect trade costs. In fact, Melitz (2008, p. 676) considers area to be "a proxy for internal distance." If $area_i\,(area_j)$ represents the area of country $i\,(j)$, the corresponding MR terms are given

identical but of opposite signs. Henceforth, these restrictions are referred to as BV constraints. The BV constraints imply that $\tau = -\widetilde{\tau}$ and $\beta = -\widetilde{\beta}$. Note that with the BV constraints, the trade cost variables can be conceptualized as $t_{ij} - \mathrm{MR}t_{ij}$.

3.2 Sensitivity to Selection on Unobservables

In light of the endogeneity issues exacerbated by the lack of suitable instruments, two methodologies for assessing the extent of selection bias are discussed.

Extent of Selection on Unobservables Altonji et al. (2005, p. 153) propose a method of assessing the extent of selection on unobservables by using "the degree of selection on observables as a guide." In order to employ it in this context, the set of non-treatment covariates needs to be conceived of as a random draw from the (large) set of all factors affecting bilateral trade with no factor (observed or unobserved) having an overriding influence.[12] In such a scenario, the extent of selection on unobservables is expected to equal the amount of selection on observables. As a result, the robustness of the causal effect, obtained under exogeneity, can be determined by asking how the amount of selection on unobservables must compare to the amount of selection on observables to fully explain the estimated effect. Intuitively, a large value (i.e., greater than one) of the estimated factor or ratio, suggests robustness of the obtained treatment effect to selection on unobservables. Contrarily, a small value (i.e., at least less than one) of the ratio indicates susceptibility to selection bias.

Interestingly, according to Baldwin (2006, p. 17), "bilateral trade costs are determined by many factors, ranging from personal relationships ... to convenient flight schedules." Barr et al. (2003, p. 581) concur that the "range of variables that might influence trade is myriad." In fact, the view that the "world is not so generous as to provide observable measures" of trade costs is also shared by Baier and Bergstrand (2009b, p. 78). More succinctly, most authors including Anderson and van Wincoop (2004) agree that gravity models include an arbitrary list of trade cost proxies. Hence, despite the theoretical background of gravity models, the assumptions invoked by Altonji et al. (2005) are not tenuous in this context. However, due to the solid theoretical underpinning of the gravity model, as well as the out-of-sample forecast ability of gravity models as documented in Henderson and Millimet (2008), one might think that the assumptions of equal amounts of selection on observables and unobservables is unrealistic in this application. Thus, one

by $\mathrm{MR}\ln area_i = \sum_{i=1}^{N}\theta_i \ln area_i + \sum_{j=1}^{N}\theta_j \ln area_j - \sum_{i=1}^{N}\sum_{j=1}^{N}\theta_i\theta_j \ln area_i$, and $\mathrm{MR}\ln area_j = \sum_{j=1}^{N}\theta_j \ln area_j + \sum_{i=1}^{N}\theta_i \ln area_i - \sum_{j=1}^{N}\sum_{i=1}^{N}\theta_j\theta_i \ln area_j$.

[12] There is also an additional requirement that is weaker than independence between the observed (non-treatment) covariates and the remaining determinants of bilateral trade.

might interpret a value of anything above, say, 0.75 as indicative of a robust causal effect. Nonetheless, the author sticks to a value of one to be conservative and errs on the side of rejecting the null of a causal effect of the treatment. However, the reader should be cautious in interpreting ratios below one, but still fairly sizeable.

Before proceeding, it is essential to note that the main objective of this study is to assess the extent to which any (estimated) positive and significant association between bilateral trade and membership into RTAs, or CUs, or both, can be considered causal. As a result, unlike studies which aim at point identification, absence of a positive association, under exogeneity, does not merit attention. In other words, if the estimated treatment effect is insignificant, or even negative, then the issue of sensitivity to positive selection is moot (unless one believes in the possibility that the policy variables under consideration might actually have a negative, causal effect on bilateral trade).

To proceed, the (normalized) amount of selection on unobservables is represented as

$$\frac{\mathrm{E}\left[\varepsilon|D=1\right]-\mathrm{E}\left[\varepsilon|D=0\right]}{\mathrm{Var}\left(\varepsilon\right)} \tag{4}$$

where D denotes the treatment under consideration and ε depicts the unobservables in (1) and (2). Similarly, the amount of selection on observables, adjusted for variance, is formalized by

$$\frac{\mathrm{E}\left[Z\delta|D=1\right]-\mathrm{E}\left[Z\delta|D=0\right]}{\mathrm{Var}\left(Z\delta\right)} \tag{5}$$

where Z refers to the set of non-treatment regressors in the (outcome) equations, (1) and (2), and δ is the corresponding coefficient vector.[13] Under the assumption that the amount of selection on unobservables is equal to the amount of selection on observables, the ratios in (4) and (5) are expected to be equal.

Next, in the wake of evidence suggesting a trade promoting effect of any of the treatments, the idea is to quantify the amount of selection on unobservables, relative to the amount of selection on observables, that would be necessary to attribute the entire effect to selection bias. In order to compute this implied ratio, first express treatment participation as

$$D_{ij} = Z_{ij}\lambda + \upsilon_{ij}. \tag{6}$$

Substituting (6) into (1) or (2) results in

$$\ln\left(T_{ij}+1\right) = Z_{ij}\left(\delta+\tau\lambda\right)+\tau\upsilon_{ij}+\varepsilon_{ij}. \tag{7}$$

[13] Since Z represents the set of all regressors except D, it includes X. While it also includes the country fixed effects in (1), in case of (2) it denotes the GDPs and the MR terms from BVOLS, in addition to X.

Now, the probability limit of the OLS estimate of τ can be decomposed into the true treatment effect and bias as

$$\begin{aligned} \text{plim}\widehat{\tau} &= \tau + \frac{\text{Cov}(\upsilon, \varepsilon)}{\text{Var}(\upsilon)} \\ &= \tau + \frac{\text{Var}(D)}{\text{Var}(\upsilon)} \{\text{E}[\varepsilon|D=1] - \text{E}[\varepsilon|D=0]\}. \end{aligned} \tag{8}$$

Given the assumptions invoked above, the amount of selection on unobservables is expected to be equal to the extent of selection on observables. Accordingly, the bias term in (8) is expressed as

$$\frac{\text{Cov}(\upsilon, \varepsilon)}{\text{Var}(\upsilon)} = \frac{\text{Var}(D)}{\text{Var}(\upsilon)} \left\{ \frac{\text{E}[Z\delta|D=1] - \text{E}[Z\delta|D=0]}{\text{Var}(Z\delta)} \text{Var}(\varepsilon) \right\}. \tag{9}$$

From (9), computation of the bias term requires identification of δ. However, under the null of no treatment effect, δ can be estimated from (7) with τ constrained to be zero. After this, only the sample values of Var(D) and Var(υ) are required in order to compute the bias.

Finally, the implied ratio is computed by dividing the estimate of τ, obtained under exogeneity, by the bias calculated from (9).

Since the MR terms are controlled for by using country fixed effects or the BVOLS terms, both approaches are employed while assessing the robustness of treatment effects. In fact, although the BVOLS approach is not needed in the log models, results from both approaches are presented nonetheless to convince the reader that each is valid. This is useful since later, in the bivariate probit model, we rely exclusively on the BVOLS approach. In this context, it should be noted that the although the Taylor expansion in Baier and Bergstrand (2009b) recommends imposition of the BV constraints, the BVOLS method is also employed without the constraints imposed. Also, since the bias term in (9) is computed under the null of no treatment effect, the MR term for the treatment variable, MRD, is not included as a regressor in (6) or (7) when τ is constrained to be zero.

Bivariate Probit Model Given recent concerns over the evolution of trade at the intensive and extensive margins, the selection issue is also analyzed at the extensive margin of trade by conducting a bivariate probit analysis along the lines of Altonji et al. (2005). To the author's knowledge, Egger et al. (2009) is the only study in the trade agreements literature to assess the issue of endogeneity by estimating a bivariate probit model. However, instead of relying on exclusion restrictions such as GDP similarities or differences in relative factor endowments between two countries, this paper contributes to the literature by examining the robustness of any evidence suggesting a trade promoting role of CUs or RTAs at the extensive margin. The model is represented as

$$T_{ij} = \mathrm{I}\left(\tau D_{ij} + Z_{ij}\delta + \varepsilon_{ij} > 0\right) \tag{10}$$
$$D_{ij} = \mathrm{I}\left(Z_{ij}\lambda + \upsilon_{ij} > 0\right)$$

where T_{ij} is a binary variable assuming the value unity in case of positive bilateral trade between i and j, $\mathrm{I}(\cdot)$ is the indicator function, D_{ij} continues to represent the treatment under consideration, and $\varepsilon, \upsilon \sim N_2(0, 0, 1, 1, \rho)$. The correlation between unobservables determining a country-pair's decision to engage in bilateral trade and unobservables that affect their likelihood of entering into a trade agreement, or currency union, or both, is denoted by ρ. As a result, while $\rho > 0$ denotes positive selection on unobservables, negative selection is depicted by $\rho < 0$. In this context, it is important to note that Z_{ij} includes the set of all regressors in (2), except D_{ij}, since the MR terms are controlled for by adopting the BVOLS approach. Also, estimation is performed with and without the imposition of the BV constraints in the outcome equation.[14]

Now, the lack of reliable instruments in the context of the log model also raises concerns over the availability of suitable exclusion restrictions here. While the model can still be identified off the assumption of bivariate normality, results obtained solely from such parametric assumptions are not viewed as reliable. Again, since the ultimate focus lies in assessing the robustness of any apparent trade promoting effect obtained under exogeneity, an alternative methodology is adopted. First, the parametric restriction is made less severe by constraining ρ and estimating the conditional likelihood function (i.e., the model conditional on the constrained value of ρ). Next, sensitivity to increasing amounts of positive selection is gauged by constraining ρ to 0, 0.1, ..., 0.5. In addition, the unconstrained model is also estimated, relying solely on the parametric assumption, to provide some indication of the likely value of ρ.

4 Data

The majority of the data come from Liu (2009); thus, only limited details are provided.[15] More than 80% of the bilateral imports data are obtained from the International Monetary Fund's Direction of Trade Statistics (DOT). Any missing import data is replaced by the corresponding export data. If neither are

[14] Note that the theory in Baier and Bergstrand (2009b) only discusses the BV constraints in the context of the outcome equation. Hence, the selection equation includes the relevant terms without imposing the constraints. However, if one conceives of the trade cost variables as $t_{ij} - \mathrm{MR}t_{ij}$, then Altonji et al.'s (2005) proposed method requires the imposition of the BV constraints in (6).

[15] Since the data includes zero trade observations and more countries than used in most previous analyses, Liu (2009) considers it to be relatively more complete.

available from the DOT, the World Trade Flows dataset (developed by Robert C. Feenstra) and the World Export Dataset (developed by Jan Faber and Tom Nierop) are relied on. However, the two sources except DOT, are used for only 5% of the observations in Liu (2009). The GDP data mainly come from the Penn World Table 6.1, or other sources such as Penn World Table 5.6, World Development Indicators 2003, the International Monetary Fund's International Financial Statistics (IFS), and the United Nations (UN) Statistical Yearbooks. The UN publication: Operation and Effects of the Generalized System of Preferences, and an additional UN source, the Generalized System of Preferences List of Beneficiaries (2001), provide the GSP data. While a measure of hostility is obtained from the Militarized Interstate Dispute Dataset (Ghosn and Palmer, 2003), the information on formal alliances is acquired from the Formal Alliance dataset (Gibler and Sarkees, 2004). However, it is important to note that the RTA and CU variables have been modified.

Since most analyses pertaining to the effects of trade agreements rely on the WTO's RTA database, (see, e.g., Baier and Bergstrand, 2007, Egger and Larch, 2008, Magee, 2008, Liu, 2009), it is important to note that a new trade agreements database was launched in January 2009.[16] Apart from providing more updated information on existing agreements, it also allows for additional information on trade agreements which are no longer active. This provides an opportunity to correctly include a number of additional agreements that were previously omitted in the literature. For example, although a free trade agreement existed between European Free Trade Association (EFTA) and Spain from 1980 to 1986, it has mostly been neglected in the literature.[17] As a result, the new database has been extensively consulted in order to incorporate such agreements. In addition, sources such as Frankel (1997), Jovanovic (1998), World Development Indicators 2008, and RTA secretariat webpages have also been used. The additional sources not only provide a useful check, but are also essential in the context of agreements which have not been notified to the WTO. For example, Frankel (1997) treats Group of Three, formed between Colombia, Mexico, and Venezuela, as a free trade agreement. However, it does not find mention on the WTO's RTA database. In light of all this, the RTA dummy has been redefined for the purpose of this study. Although a number of sources were consulted for creating the trade agreement variable, Frankel (1997) and the WTO's database have been relied on the most. For further clarity, Table A1, in the appendix, lists the RTAs considered. The RTA variable was created using the agreements' dates of entry into force.

The CU dummy in Liu (2009) is obtained from Glick and Rose (2002, p. 1128), who consider a country-pair to be in a CU if "money was interchangeable" between them "at a 1:1 par for an extended period of

[16] See http://www.wto.org/english/news_e/pres09_e/pr548_e.htm.

[17] Note that the agreement for the mentioned years, did not feature in Liu's (2009) original data as well. See http://www.efta.int/content/about-efta/history/history-of-efta/?searchterm=spain%20agreement%201979.

time." However, in this context, a few points are noteworthy. First, Glick and Rose (2002) only include observations upto 1997. As a result, the euro, which was adopted in 1999, did not feature in the original dataset. In addition, the CU dummy in Liu's (2009) data involved errors which have been corrected.[18] While most corrections are based on the appendix and data from Glick and Rose (2002), a number of additional sources have been consulted for the zero trade observations, which did not feature in Glick and Rose (2002). Again, for further clarity, Table A2, in the appendix, lists the CUs considered along with the relevant years. In keeping with Glick and Rose (2002), specific names of CU regimes have not been provided. However, instead of listing country-pairs, for brevity, groups of CU members have been listed in separate panels. The sources consulted find mention in the table footnotes.[19]

The 2003 CIA Fact Book is relied on for data on the other trade cost variables including latitudes and longitudes for constructing great-circle distances.

5 Results

Log Results Even a cursory glance at the summary statistics presented in Table 1 suggests selection (on observables) into RTAs and CUs. For example, country pairs which are RTA (CU) members not only seem to engage in more bilateral trade than non-RTA (non-CU) countries, but are also more likely to be characterized by proximity and adjacency. As a result, the summary statistics provide further motivation for examining the selection issue.

Accordingly, Tables 2, 3, and 4 utilize cross-section data for the years 1950, 1960, ..., 2000, in providing coefficient estimates from the log model. The results in table 2 correspond to the case where the treatment dummy, D, in Section 3 is defined as one if countries i and j share a RTA. Similarly, the results in Table 3 correspond to CU as the treatment. For results pertaining to Table 4, the treatment variable, D, assumes the value one if i and j share a CU and RTA. Apart from considering country pairs belonging to both CUs and RTAs as the treatment group, Table 4 also differs from Tables 2 and 3 by providing estimates from only 1960 onwards. This is necessitated by the presence of extremely few country pairs belonging to both CUs and RTAs during 1950. Tables 3 and 4 also distinguish themselves from Table 2 by providing two sets of estimates for 2000 - with and without the euro countries as members of a CU. Given the infancy of euro, and the fact that the euro countries differ from members of other CUs in terms of economic size, this is relevant.[20]

[18] The author would like to thank Xuepeng Liu and Andrew Rose for their help in this.

[19] A few minor corrections, were also made to the WTO membership variables. As a result of the corrections, Czechoslovakia (for 1950 to 1990), Lebanon (for 1950), Liberia (for 1950), and Syria (for 1950) are treated as members.

[20] Note that Frankel's (2008) findings do not support either hypothesis as an explanation for the typically observed discrep-

Before proceeding, it is useful to note that for all specifications and cross-sections across Tables 2, 3, and 4, the columns first report the treatment effect estimated under exogeneity followed by the bias computed from (9). Next, the implied ratio, obtained from dividing the estimated treatment effect by the bias term, is displayed. Since results from the use of both fixed effects and BVOLS are reported, it is also important to note that while both approaches lead to similar coefficient estimates, coverage rates from the first Monte Carlo analysis in Baier and Bergstrand (2009b) find the fixed effects approach to be slightly more reliable.

Now, the fixed effects estimates in Table 2 find RTA members to be associated with significantly greater amounts of bilateral trade from 1970 onwards.[21] In fact, the positive and significant estimates, also witnessed for 1950, are significant at the 1% level (of significance). Surprisingly, this positive association is similar in magnitude for all cross-sections except 1950. However, instead of attaching too much weight on the point estimates obtained under exogeneity, their robustness to selection on unobservables needs to be examined. In this light, the results obtained are striking. Apart from 1950, the positive RTA estimates are accompanied by ratios varying between 0.2 and 0.3. In other words, even if the amount of (normalized) selection on unobservables is thirty percent the amount of (normalized) selection on observables, the RTA effects estimated under exogeneity can be completely explained. Since such a modest amount of selection on unobservables is sufficient to entirely attribute the positive and significant RTA estimates to a selection effect, there is hardly any evidence in support of RTAs causing an increase in bilateral trade.

Although there is fairly strong evidence in favor of RTAs to have caused bilateral trade in 1950, it is more interesting to note that for all positive and significant RTA coefficients, the estimated bias term is always greater than zero suggesting positive selection into trade agreements. This is consistent with the "natural trading partner hypothesis" discussed in Magee (2008, p. 350), i.e., "countries ... tend to form regional agreements if they already have significant bilateral trade."

Interestingly, regardless of the imposition of BV constraints, the BVOLS approach presents a very similar picture. Not only are the RTA estimates similar in sign and magnitude to those obtained using country fixed effects, but so are the bias terms and implied ratios. In other words, there is strong evidence suggesting that any RTA effect, observed for years more recent than 1950, is not causal but essentially a reflection of positive selection.

Turning to CUs, Frankel (2008, p. 6) opines that "endogeneity ... is perhaps the most intractable

ancy in trade promoting effects of the euro and other CUs. Although Frankel (2008) suspects the typically small sample sizes in most euro studies as a possible explanation, the suspicion is driven by gravity equations which do not control for the MR terms.

[21] The (fixed effects) coefficient estimates of the other regressors are similar to those found in the literature, but are not the focus of the paper. As a result, they do not find mention, but are available upon request.

problem with ... Rose-style estimates." However, Glick and Rose (2002) consider endogeneity to be a non-issue as far as currency unions are considered. In this light, it is interesting to examine the CU results. From Table 3, the use of country fixed effects yields positive, large and statistically significant estimates of CU membership on bilateral trade for most years except 1960. Even before embarking on a discussion of causality, it is interesting to note that the CU estimates are the largest for 1970, 1980, and 1990. In fact, it is even more interesting to recall that Rose's (2000) estimates, which are among the largest to have been reported in the CU literature, rely on data from 1970 to 1990 only. In further consonance with the literature, the CU effect is considerably smaller for 2000 when countries belonging to the EMU are treated as CU members. While Rose and van Wincoop (2001, p. 388) use data for 1980 and 1990 to suggest "a smaller effect of the EMU ... than most other currency unions," here the data for 2000 find the effect to be small enough to render the CU effect insignificant.

Strikingly, the positive and significant CU coefficient estimates are found to be quite robust to selection on unobservables. All corresponding implied ratios are negative indicating that in order to fully explain the estimated effects obtained under exogeneity, selection on unobservables must necessarily work in the direction opposite to that of selection on observables. To be clear, the negative bias term implies that country pairs that share a currency possess worse observables, in terms of raising bilateral trade, than country pairs in the control group. Thus, while the treatment group possesses worse observables, they must possess better unobservables, in terms of raising bilateral trade, to explain the estimates obtained under exogeneity. Since this is unlikely to be the case, the evidence suggests a highly robust and causal CU effect. In other words, CUs are found to have "a genuine positive trade effect" as evidenced by Rose and Stanley's (2005, p. 347) meta-analysis. In fact, the causal effect of CU membership is found to be even stronger for 1970, 1980, and 1990 as the implied ratios are not only negative, but also greater than one in magnitude. Moreover, the negative bias terms corresponding to the positive CU effects further vindicate the results in Barro and Tenreyro (2007), who rely on IV only to find the positive and significant OLS coefficients to be downward biased. It is interesting to note that the evidence in favor of negative selection in Barro and Tenreyro (2007) is consistent with the current finding of the treatment group having worse observables with respect to trade promotion, under the assumption of equal amounts of selection on observables and unobservables.

The BVOLS estimates continue to support the findings from the fixed effects approach. Although the coefficient estimate for 2000 (Panel VI) is positive only when the BV constraints are not imposed, the corresponding values of bias and ratio hardly suggest a robust causal effect.[22]

[22] Before proceeding, note that some gravity studies include both RTA and CU as regressors in the same equation. However, the estimated RTA (CU) effects were hardly found to change upon the inclusion of CU (RTA) as a regressor. In fact, this was

Since RTA members and countries using a common currency are on average engaged in higher levels of trade, but only the CU effect is robust to selection on unobservables, Table 4 presents an interesting scenario. Now, the treatment is defined as sharing both a currency union and belonging to a RTA. Surprisingly, the fixed effects estimates are no longer significant for 1970 despite the individual partial effects of RTAs and CUs being positive and significant for that year. Thus, even under exogeneity, it is potentially misleading to determine the association between a joint policy and trade by "summing" the associations between each policy considered in isolation. However, the fixed effects estimates for 1980 and 1990 are not only positive and significant, but also indicate a causal relationship. In other words, the amount of selection on unobservables relative to the extent of selection on observables required to completely relegate the positive finding to a selection effect, is much greater than one or negative. It is also worth noting that the coefficient estimates for 2000 are sensitive to the inclusion of the euro countries in the treatment group.[23] However, all positive coefficient estimates appear strikingly robust to selection on unobservables. This finding of a robust trade promoting effect of simultaneous membership in both a CU and RTA could not have been deduced from the causal CU and mostly non-causal RTA effects. Hence, the necessity of modeling policy combinations separately is further justified.

Again, the BVOLS estimates provide ample support to the fixed effects results. The only apparent discrepancy appears in case of 2000 (Panel V) when the coefficient estimate is positive and significant. However, modest selection on unobservables to the tune of even thirteen percent the amount of selection on observables is sufficient to render it non-causal.[24]

Bivariate Probit Results According to Egger et al. (2009) and Liu (2009), among others, it is important to distinguish between an increase in the volume of trade among existing trading partners (intensive margin) and the establishment of new bilateral trading relationships (extensive margin). In fact, Felbermayr and Kohler (2006, p. 643) go on to describe gravity studies that neglect the extensive margin as "inadequate." Hence, in light of such concerns and given the potential selection issue, a (cross-section) bivariate probit analysis for the three treatment variables is performed. The results find mention in Tables 5, 6, and 7.[25] Across all three tables, the MR terms are controlled for using the BVOLS approach, which

true regardless of the BVOLS or fixed effects approach. For the fixed effects approach, the (cross-section) correlations between CU and RTA, after conditioning on all regressors, were found to be less than 0.075 in the majority of cases, and reached a maximum of 0.1 in 1970.

[23] Incidentally, countries which adopted the euro are also RTA members.

[24] Although the dataset contains a large (about fifty percent) number of zero trade observations, an additional check was performed before estimating the probit model. The log model was estimated after dropping observations for a country prior to its independence if all pairwise observations containing that country had zero trade, prior to its independence. The results remained very similar.

[25] A relevant concern in the context of extensive margin is the issue of birth of nations (e.g., due to decolonization). However,

is employed with and without the BV constraints in the outcome equation. To proceed, the estimation is first performed without constraining ρ. In this case, the values of τ and ρ are identified from the non-linearity arising from the assumption of bivariate normality. Next, the coefficient estimates of the treatment variables are obtained after constraining the degree of selection from zero to positive amounts. If the estimated effect of the treatment is positive and statistically significant when ρ is constrained to zero, but quickly becomes statistically insignificant as ρ deviates from zero, then this is evidence of a non-robust causal effect. Moreover, the unconstrained ρ gives some indication of the likely value of ρ in reality.

Table 5 finds the RTA coefficient estimates, obtained under exogeneity (i.e., $\rho = 0$), to be positive and statistically significant for all years except 1960.[26] In other words, RTA members are associated with a higher probability of forming a new trading relationship. In fact, the 1950 estimates are also robust to the varying amounts of positive selection considered. Even if there is positive selection to the tune of $\rho = 0.5$, the treatment effect under exogeneity cannot be explained away. Moreover, given the negative estimates of ρ in the unconstrained model, a value greater than 0.5 is unlikely. As a result, regardless of the BV constraints, there is strong evidence suggesting that the RTA effect on the extensive margin for 1950 is causal with the marginal effects (unreported) varying between 0.78 and 0.56.[27]

However, for 1970, 1990, and 2000, there is hardly any evidence of causality at the extensive margin. For each of these years, irrespective of the BV constraints, the estimated ρ is mostly greater than the amount of selection required to render the RTA effect insignificant or negative. For example, when $\rho = 0$, the 1970 estimates obtained after imposing the BV constraints find RTA members to be significantly more likely to establish a new trading relationship. However, if ρ is increased to even 0.2, the magnitude of selection is sufficient to reduce the marginal effect from 0.15 to 0.04. In fact, if ρ increases to 0.3, the coefficient estimates are rendered negative. Since the unconstrained estimate of ρ is greater than 0.4, a value of ρ around 0.3 is certainly plausible. Accordingly, the evidence against a robust trade promoting effect of RTA membership is strong. Similarly, for 1990, a likely value of ρ around 0.3 renders any positive association between bilateral trade (at the extensive margin) and RTA membership negative. In case of 2000, the amount of selection required to attribute the positive and significant RTA coefficient estimates, obtained under exogeneity, to a selection effect is even smaller.[28] Intriguingly, hitherto, the effects of RTAs

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Liu (2009) does not consider the pattern of extensive margin in this dataset to be driven by the emergence of new countries. Moreover, Felbermayr and Kohler (2009) consider the issue to be less significant with cross-section data.

[26] The author reports non-robust standard errors to be more conservative (i.e., the author does not want to find the statistically significant effect obtained under exogeneity to disappear quickly due to large standard errors). Nonetheless, the results are virtually unchanged if one uses robust standard errors.

[27] Using the notation in (10), the marginal effects are computed as $\Phi\left(\widehat{\tau} + \overline{Z}\widehat{\delta}\right) - \Phi\left(\overline{Z}\widehat{\delta}\right)$, where $\overline{Z}$ depicts the sample mean of all covariates except D, and $\widehat{\tau}$ and $\widehat{\delta}$ are the coefficient estimates from the bivariate probit.

[28] For 2000, the evidence against a causal effect is less clear. Although the estimates obtained without imposing the BV

at the extensive margin and on the volume of bilateral trade appear similar. On the other hand, the results pertaining to 1980 are different. A line of reasoning similar to that adopted for the other cross-sections finds strong evidence in favor of some sort of a causal effect of RTAs on trade during 1980.[29] The results for 1980 further highlight the importance of the sensitivity analysis undertaken. Although the corresponding RTA coefficients, under exogeneity, are positive and significant across the log and probit models, the evidence in favor of a causal effect is only obtained at the extensive margin.

Turning to the CU results in Table 6, perhaps the most striking aspect is the fact that even when the euro countries are considered as part of the treatment group, CU members (during 2000) are associated with a significantly greater probability of establishing a new trading relationship. However, a closer look at the estimates obtained after imposing the BV constraints dispel any thoughts of a robust causal effect. To be more clear, the positive and significant CU effect obtained under exogeneity turns insignificant at the 95% confidence level if the amount of positive selection is extremely modest with $\rho = 0.1$. Moreover, a value of ρ around 0.2 is sufficient to rule out a significant CU effect even at the 90% confidence level. Since the unconstrained estimate of ρ, obtained solely from the parametric assumptions of the model, exceeds 0.3, the evidence against a robust CU effect is strong.[30]

On the other hand, the coefficient estimates for 1970 to 1990, and for 2000 without the euro countries considered as CU members, are found to be robust to selection on unobservables. In other words, for most of these years, there is strong evidence suggesting CUs have a causal effect on trade at the extensive margin, with the marginal effects between 0.05 and 0.19. More precisely, in the absence of any non-random selection, the marginal effect of CU membership is found to be at least 0.15 during 1970. While the marginal effects are hardly sensitive to the imposition of the BV constraints, they are also similar when identified solely from the assumption of bivariate normality. Although the positive and significant 1970 CU estimates turn insignificant if positive selection is to the tune of $\rho = 0.3$, the unconstrained estimates of ρ highlight the implausibility of such values. Following a similar logic, there is strong evidence suggestive of a robust causal CU effect during 1990 and 2000 with no euro members in the treatment group. The evidence for 1980 is unclear.

Importantly, while there exists strong indication of some sort of a positive CU effect, its magnitude

constraints continue to find the effect of RTAs to be quite sensitive to non-random selection, given Baier and Bergstrand's (2009b) theory, the estimates obtained after imposing the BV constraints are considered more reliable. However, the fact that the RTA estimate changes from positive and significant to negative when ρ changes from 0.1 to 0.2, and given that the estimated ρ lies between these values, an unambiguous conclusion on robustness is extremely difficult.

[29] The marginal effects for the positive and significant estimates range from 0.05 to 0.18.

[30] Although the estimates without the BV constraints suggest a causal effect, in keeping with the theory in Baier and Bergstrand (2009b), the estimates obtained after imposing the constraints are relied upon.

is clearly sensitive to selection. For example, the 1990 marginal effect of CU, obtained after imposing the BV constraints, declines from a value of 0.17 to 0.01 when ρ is increased from 0 to 0.4.[31] If the euro members are relegated to the control group and the BV constraints are also imposed, the marginal effect of CU membership during 2000 decreases from 0.14 to 0.06 when ρ is increased from 0 to 0.3. In fact, across all years, an amount of selection to the tune of $\rho = 0.3$ is sufficient to render any positive association inconsequential. Thus, although the evidence in favor of CUs promoting bilateral trade is strong, policymakers interested in *quantifying* the CU effect should continue to be wary of the selection issue.

For country pairs which form trade agreements in addition to adopting a common currency, the results in Table 7 find some evidence of a robust effect on trade promotion only for 1990 and 2000 if none of the common currencies happen to be euro. While the value of ρ is required to be 0.4 to explain the positive and significant association during 1990, a more modest amount of selection is necessary to classify the estimates pertaining to 2000 as a selection effect. However, such positive values of ρ seem highly unlikely given the negative unconstrained estimates.

To summarize, in the wake of evidence obtained here, concerns over selection into trade agreements are warranted. However, with respect to CUs, one is inclined to agree with Rose's (2001, p. 456) conjecture that "non-random selection" is of "academic interest, but unimportant in practice," only if the magnitude of the CU effect is unimportant.

6 Conclusion

RTAs and CUs share the characteristics of being trade cost proxies whose coefficient estimates are not only of considerable interest, but are also conceived to be susceptible to selection bias. In both cases, this problem is magnified by the paucity of reliable instruments. As a result, studies estimating the effects of both policy regimes on members' bilateral trade have mostly resorted to the use of panel data, but as is well known this only controls for selection on time-invariant unobservables.

Here, instead of seeking point estimates of the causal effect of RTAs and CUs under perhaps implausible assumptions, this paper adopts a novel approach by assessing the extent to which a positive association between bilateral trade and CU or RTA membership can be considered causal. The results obtained are striking. Although most cross-sections find both RTAs and CUs to be associated with increased bilateral trade, the evidence in favor of a causal effect is strong only for CUs. Contrarily, most positive and significant RTA coefficients, estimated under the assumption of exogeneity, can be explained by even

[31] The marginal effect corresponding to the unconstrained estimate of ρ is 0.21.

modest levels of positive selection. Accordingly, concerns over selection bias are well-founded in the context of trade agreements. However, with respect to CUs, in light of evidence from the novel method adopted here, Rose's (2001, p. 457) recommendation of adjusting "priors," seems a wiser strategy than trying yet another "radically new approach." Interestingly, a similar pattern of selection is also witnessed at the extensive margin. Moreover, results from the paper suggest that in order to determine the trade-inducing effect of membership into both CUs and RTAs, the policy combination should be modeled as a separate regime. Finally, while the evidence here finds CUs to have at least some sort of causal effect, the *magnitude* of the effect is sensitive to the level of selection on unobservables. Thus, from a policy perspective, future work should still look into identifying point estimates of the causal effect of CUs.

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Table 1. Summary Statistics

Variable	Full Sample Mean	Full Sample SD	RTA Mean	RTA SD	Non-RTA Mean	Non-RTA SD	CU Mean	CU SD	Non-CU Mean	Non-CU SD
RTA (1=Yes)	0.101	0.302	1.000	0.000	0.000	0.000	0.185	0.388	0.099	0.299
CU (1=Yes)	0.024	0.152	0.043	0.202	0.021	0.145	1.000	0.000	0.000	0.000
Bilateral Imports	1.010e+08	1.490e+09	4.170e+08	3.680e+09	6.480e+07	9.660e+08	2.640e+08	2.340e+09	9.660e+07	1.460e+09
Distance	7721.259	4449.068	5895.604	4615.478	7927.492	4382.341	3989.992	3667.726	7811.100	4427.607
Border (1=Adjacent Countries)	0.019	0.137	0.062	0.242	0.014	0.118	0.104	0.305	0.017	0.129
Common Language (1=Yes)	0.118	0.322	0.175	0.380	0.111	0.315	0.632	0.482	0.105	0.307
Common Religion (1=Yes)	0.517	0.500	0.648	0.478	0.503	0.500	0.621	0.485	0.515	0.500
Colony (1=Yes)	0.007	0.082	0.021	0.143	0.005	0.072	0.042	0.201	0.006	0.077
Colonizer (1=Yes)	0.007	0.082	0.021	0.143	0.005	0.071	0.042	0.201	0.006	0.076
Current Colony (1=Yes)	0.002	0.045	0.002	0.047	0.002	0.045	0.034	0.182	0.001	0.035
Current Colonizer (1=Yes)	0.002	0.045	0.002	0.047	0.002	0.045	0.034	0.182	0.001	0.035
Common Colonizer (1=Yes)	0.164	0.370	0.161	0.368	0.164	0.370	0.765	0.424	0.149	0.356
Hostility	0.010	0.092	0.018	0.131	0.010	0.086	0.019	0.112	0.010	0.091
Alliance (1=Countries in a Formal Alliance)	0.056	0.230	0.173	0.378	0.043	0.203	0.107	0.309	0.055	0.228
Both WTO Members (1=Yes)	0.287	0.452	0.604	0.489	0.251	0.434	0.245	0.430	0.288	0.453
One WTO Member (1=Yes)	0.464	0.499	0.346	0.476	0.477	0.499	0.319	0.466	0.467	0.499
Generalized System of Preferences, Offered by Importer (1=Yes)	0.071	0.257	0.195	0.396	0.057	0.232	0.012	0.111	0.072	0.259
Generalized System of Preferences, Offered by Exporter (1=Yes)	0.071	0.257	0.195	0.396	0.057	0.232	0.012	0.111	0.073	0.260
Importer GDP	161485.000	586181.900	249229.900	606487.000	151167.100	582883.100	107599.000	631936.100	162720.200	585035.400
Exporter GDP	162513.100	585779.200	249239.900	606493.900	152329.900	582454.700	100968.700	600340.400	163923.000	585367.800
Importer Area	868129.900	2389365.000	623054.000	1710846.000	895814.700	2452716.000	506250.300	1215196.000	876843.100	2409928.000
Exporter Area	882677.600	2424206.000	623929.800	1710930.000	911906.800	2490295.000	494521.300	1164244.000	892023.500	2445799.000

Notes: N = 146,948 (full sample); 14,915 (RTA); 132,033 (Non-RTA); 3455 (CU); 143,493 (Non-CU). For all samples, missing GDP values are less than 5%, but around 9% for the CU sample. Most data are from Liu (2009), except the RTA and CU dummies. Observations from 1950 to 2000, at ten year intervals, are pooled.

Table 2. Regional Trade Agreement: Amount of Selection on Unobservables Relative to Selection on Observables Required to Attribute the Entire Treatment Effect to Selection Bias

	τ	Cov(ε,ν)÷ Var(ν)	Implied Ratio	Number of Observations	Number of Treated
I. 1950					
Fixed effects	6.545***	3.929	1.666	20435	58
	(1.008)				
Unconstrained BV	9.282***	9.464	0.981	15801	44
	(1.096)				
Constrained BV	12.273***	6.473	1.896	15801	44
	(0.941)				
II. 1960					
Fixed effects	-2.647***	7.372	-0.359	21361	188
	(0.391)				
Unconstrained BV	-3.116***	11.829	-0.263	17662	174
	(0.487)				
Constrained BV	-4.021***	7.952	-0.506	17662	174
	(0.585)				
III. 1970					
Fixed effects	1.457***	5.815	0.251	22592	705
	(0.220)				
Unconstrained BV	1.794***	9.032	0.199	21743	695
	(0.236)				
Constrained BV	2.968***	6.199	0.479	21743	695
	(0.221)				
IV. 1980					
Fixed effects	1.245***	7.316	0.170	24951	2249
	(0.145)				
Unconstrained BV	1.767***	7.852	0.225	23997	2201
	(0.139)				
Constrained BV	2.440***	5.759	0.424	23997	2201
	(0.141)				
V. 1990					
Fixed effects	1.117***	3.961	0.282	25898	4628
	(0.094)				
Unconstrained BV	1.710***	3.896	0.439	24568	4555
	(0.095)				
Constrained BV	1.935***	3.368	0.575	24568	4555
	(0.097)				
VI. 2000					
Fixed effects	1.093***	3.980	0.275	31711	7087
	(0.075)				
Unconstrained BV	1.368***	3.903	0.351	30224	6929
	(0.076)				
Constrained BV	1.634***	3.666	0.446	30224	6929
	(0.076)				

Notes: *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Robust standard errors in parentheses. Cov(ε,ν)/Var(ν) refers to the asymptotic bias of the unconstrained estimate under the assumption of equal (normalized) selection on observables and unobservables. τ refers to the unconstrained estimate of the effect of RTA membership. The implied ratio is the latter divided by the former. See text for details. The number of observations for the BVOLS are smaller due to missing GDP values.

Table 3. Currency Union: Amount of Selection on Unobservables Relative to Selection on Observables Required to Attribute the Entire Treatment Effect to Selection Bias

	τ	Cov(ε,ν)÷ Var(ν)	Implied Ratio	Number of Observations	Number of Treated
I. 1950					
Fixed effects	0.503***	-1.644	-0.306	20435	1017
	(0.158)				
Unconstrained BV	0.444**	-1.710	-0.259	15801	735
	(0.207)				
Constrained BV	0.622***	-2.761	-0.225	15801	735
	(0.206)				
II. 1960					
Fixed effects	0.129	-0.685	-0.188	21361	1007
	(0.193)				
Unconstrained BV	-0.527**	-0.983	0.537	17662	838
	(0.224)				
Constrained BV	-0.354	-1.266	0.280	17662	838
	(0.226)				
III. 1970					
Fixed effects	2.172***	-0.855	-2.542	22592	469
	(0.282)				
Unconstrained BV	1.643***	-0.888	-1.852	21743	436
	(0.343)				
Constrained BV	2.709***	-2.452	-1.105	21743	436
	(0.321)				
IV. 1980					
Fixed effects	2.771***	-0.373	-7.433	24951	271
	(0.369)				
Unconstrained BV	2.282***	-0.520	-4.386	23997	252
	(0.405)				
Constrained BV	2.220***	-1.389	-1.598	23997	252
	(0.414)				
V. 1990					
Fixed effects	2.583***	-0.583	-4.431	25898	280
	(0.342)				
Unconstrained BV	2.536***	-1.111	-2.283	24568	255
	(0.378)				
Constrained BV	2.801***	-2.241	-1.250	24568	255
	(0.372)				
VI. 2000					
Fixed effects	0.354	2.637	0.134	31711	410
	(0.237)				
Unconstrained BV	1.069***	2.976	0.359	30224	366
	(0.227)				
Constrained BV	0.127	2.373	0.054	30224	366
	(0.236)				
VII. 2000 without euro					
Fixed effects	1.869***	-0.039	-48.256	31711	282
	(0.287)				
Unconstrained BV	2.034***	-0.467	-4.351	30224	258
	(0.284)				
Constrained BV	1.516***	-0.936	-1.621	30224	258
	(0.281)				

Notes: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors in parentheses. Cov(ε,ν)/Var(ν) refers to the asymptotic bias of the unconstrained estimate under the assumption of equal (normalized) selection on observables and unobservables. τ refers to the unconstrained estimate of the effect of CU membership. The implied ratio is the latter divided by the former. See text for details. The number of observations for the BVOLS are smaller due to missing GDP values.

Table 4. Currency Unions and Regional Trade Agreements: Amount of Selection on Unobservables Relative to Selection on Observables Required to Attribute the Entire Treatment Effect to Selection Bias

	τ	Cov(ε,ν)÷ Var(ν)	Implied Ratio	Number of Observations	Number of Treated
I. 1960					
Fixed effects	-2.225***	-1.773	1.255	21361	50
	(0.487)				
Unconstrained BV	-2.151***	-1.772	1.214	17662	50
	(0.476)				
Constrained BV	-3.420***	-1.801	1.899	17662	50
	(0.498)				
II. 1970					
Fixed effects	0.188	-1.223	-0.154	22592	123
	(0.569)				
Unconstrained BV	-2.080***	0.497	-4.185	21743	123
	(0.602)				
Constrained BV	1.086*	-4.880	-0.223	21743	123
	(0.569)				
III. 1980					
Fixed effects	4.740***	0.388	12.222	24951	72
	(0.690)				
Unconstrained BV	3.795***	0.421	9.026	23997	66
	(0.685)				
Constrained BV	6.342***	-2.845	-2.229	23997	66
	(0.843)				
IV. 1990					
Fixed effects	2.995***	0.754	3.973	25898	116
	(0.451)				
Unconstrained BV	3.220***	0.339	9.488	24568	110
	(0.490)				
Constrained BV	2.930***	-1.861	-1.575	24568	110
	(0.528)				
V. 2000					
Fixed effects	-0.087	4.468	-0.019	31711	275
	(0.274)				
Unconstrained BV	0.689***	5.404	0.127	30224	247
	(0.254)				
Constrained BV	-0.673**	4.534	-0.149	30224	247
	(0.290)				
VI. 2000 without euro					
Fixed effects	2.256***	1.165	1.937	31711	147
	(0.360)				
Unconstrained BV	2.309***	1.021	2.262	30224	139
	(0.346)				
Constrained BV	1.126***	0.289	3.896	30224	139
	(0.361)				

Notes: *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Robust standard errors in parentheses. Cov(ε,ν)/Var(ν) refers to the asymptotic bias of the unconstrained estimate under the assumption of equal (normalized) selection on observables and unobservables. τ refers to the unconstrained estimate of the effect of membership into both CUs and RTAs. The implied ratio is the latter divided by the former. See text for details. The number of observations for the BVOLS are smaller due to missing GDP values.

Table 5. Regional Trade Agreements: Bivariate Probit Results with Different Assumptions Concerning Correlation Among the Disturbances

		Unconstrained ρ	Constrained ρ					
I. 1950								
Unconstrained BV	τ	1.929***	1.849***	1.831***	1.813***	1.794***	1.773***	1.750***
		(0.266)	(0.261)	(0.262)	(0.262)	(0.262)	(0.262)	(0.261)
	ρ	-0.448	0	0.1	0.2	0.3	0.4	0.5
		(0.347)						
Constrained BV	τ	2.544***	2.481***	2.469***	2.455***	2.438***	2.418***	2.395***
		(0.277)	(0.275)	(0.275)	(0.275)	(0.275)	(0.275)	(0.274)
	ρ	-0.653	0	0.1	0.2	0.3	0.4	0.5
		(0.529)						
II. 1960								
Unconstrained BV	τ	-1.977***	-1.141***	-1.256***	-1.369***	-1.479***	-1.588***	-1.697***
		(0.199)	(0.145)	(0.146)	(0.146)	(0.146)	(0.145)	(0.145)
	ρ	0.746***	0	0.1	0.2	0.3	0.4	0.5
		(0.116)						
Constrained BV	τ	-2.155***	-1.232***	-1.339***	-1.441***	-1.540***	-1.637***	-1.732***
		(0.126)	(0.123)	(0.123)	(0.123)	(0.122)	(0.121)	(0.120)
	ρ	0.926***	0	0.1	0.2	0.3	0.4	0.5
		(0.049)						
III. 1970								
Unconstrained BV	τ	-0.102	0.360***	0.206***	0.050	-0.111	-0.277***	-0.451***
		(0.137)	(0.072)	(0.072)	(0.071)	(0.071)	(0.069)	(0.068)
	ρ	0.294***	0	0.1	0.2	0.3	0.4	0.5
		(0.072)						
Constrained BV	τ	-0.280**	0.369***	0.237***	0.103*	-0.034	-0.174***	-0.320***
		(0.116)	(0.061)	(0.061)	(0.061)	(0.061)	(0.060)	(0.059)
	ρ	0.473***	0	0.1	0.2	0.3	0.4	0.5
		(0.067)						
IV. 1980								
Unconstrained BV	τ	0.719***	0.475***	0.309***	0.138***	-0.037	-0.219***	-0.409***
		(0.102)	(0.045)	(0.044)	(0.044)	(0.043)	(0.042)	(0.041)
	ρ	-0.151**	0	0.1	0.2	0.3	0.4	0.5
		(0.058)						
Constrained BV	τ	0.710***	0.451***	0.303***	0.152***	-0.004	-0.165***	-0.334***
		(0.095)	(0.039)	(0.039)	(0.039)	(0.039)	(0.038)	(0.037)
	ρ	-0.177***	0	0.1	0.2	0.3	0.4	0.5
		(0.060)						
V. 1990								
Unconstrained BV	τ	-0.586***	0.388***	0.217***	0.043	-0.134***	-0.315***	-0.499***
		(0.116)	(0.030)	(0.030)	(0.030)	(0.029)	(0.029)	(0.028)
	ρ	0.546***	0	0.1	0.2	0.3	0.4	0.5
		(0.059)						
Constrained BV	τ	-0.853***	0.376***	0.212***	0.046*	-0.123***	-0.295***	-0.472***
		(0.070)	(0.028)	(0.028)	(0.028)	(0.028)	(0.027)	(0.026)
	ρ	0.706***	0	0.1	0.2	0.3	0.4	0.5
		(0.034)						
VI. 2000								
Unconstrained BV	τ	-0.572***	0.256***	0.084***	-0.091***	-0.268***	-0.449***	-0.634***
		(0.102)	(0.027)	(0.027)	(0.027)	(0.026)	(0.026)	(0.025)
	ρ	0.467***	0	0.1	0.2	0.3	0.4	0.5
		(0.053)						
Constrained BV	τ	0.075	0.273***	0.112***	-0.052**	-0.218***	-0.388***	-0.563***
		(0.110)	(0.025)	(0.025)	(0.025)	(0.025)	(0.024)	(0.024)
	ρ	0.122*	0	0.1	0.2	0.3	0.4	0.5
		(0.066)						

Notes: *** p<0.01, ** p<0.05, * p<0.1. Standard errors in parentheses. See text for details. For convergence, curcolony and curcolonizer had to be dropped 1990 onwards, only for the unconstrained BV models. Only curcolony had to be dropped for the 1980 unconstrained BV model. For each cross section, within each specification, the set of variables is the same.

Table 6. Currency Unions: Bivariate Probit Results with Different Assumptions Concerning Correlation Among the Disturbances

		Unconstrained ρ	Constrained ρ					
I. 1950								
Unconstrained BV	τ	0.511***	0.156*	0.008	-0.140	-0.291***	-0.446***	-0.605***
		(0.177)	(0.093)	(0.093)	(0.092)	(0.091)	(0.090)	(0.088)
	ρ	-0.240**	0	0.1	0.2	0.3	0.4	0.5
		(0.101)						
Constrained BV	τ	0.037	0.022	-0.107	-0.235***	-0.364***	-0.496***	-0.632***
		(0.139)	(0.084)	(0.084)	(0.083)	(0.083)	(0.082)	(0.080)
	ρ	-0.012	0	0.1	0.2	0.3	0.4	0.5
		(0.085)						
II. 1960								
Unconstrained BV	τ	-0.134	-0.050	-0.194***	-0.337***	-0.479***	-0.622***	-0.768***
		(0.123)	(0.071)	(0.071)	(0.071)	(0.070)	(0.069)	(0.067)
	ρ	0.058	0	0.1	0.2	0.3	0.4	0.5
		(0.070)						
Constrained BV	τ	-0.303***	-0.140**	-0.269***	-0.397***	-0.524***	-0.651***	-0.781***
		(0.103)	(0.066)	(0.066)	(0.065)	(0.065)	(0.064)	(0.062)
	ρ	0.127**	0	0.1	0.2	0.3	0.4	0.5
		(0.062)						
III. 1970								
Unconstrained BV	τ	0.348***	0.384***	0.258***	0.131*	0.001	-0.134*	-0.276***
		(0.112)	(0.076)	(0.076)	(0.076)	(0.075)	(0.074)	(0.072)
	ρ	0.028	0	0.1	0.2	0.3	0.4	0.5
		(0.066)						
Constrained BV	τ	0.430***	0.408***	0.296***	0.184***	0.067	-0.054	-0.183***
		(0.099)	(0.070)	(0.070)	(0.069)	(0.069)	(0.068)	(0.066)
	ρ	-0.020	0	0.1	0.2	0.3	0.4	0.5
		(0.063)						
IV. 1980								
Unconstrained BV	τ	0.457***	0.523***	0.404***	0.282***	0.157	0.028	-0.107
		(0.152)	(0.098)	(0.097)	(0.097)	(0.096)	(0.094)	(0.092)
	ρ	0.055	0	0.1	0.2	0.3	0.4	0.5
		(0.097)						
Constrained BV	τ	0.209*	0.424***	0.328***	0.231***	0.131	0.028	-0.080
		(0.121)	(0.085)	(0.085)	(0.084)	(0.083)	(0.082)	(0.081)
	ρ	0.222**	0	0.1	0.2	0.3	0.4	0.5
		(0.088)						
V. 1990								
Unconstrained BV	τ	0.713***	0.496***	0.359***	0.220**	0.077	-0.071	-0.226**
		(0.170)	(0.099)	(0.099)	(0.098)	(0.097)	(0.095)	(0.092)
	ρ	-0.163	0	0.1	0.2	0.3	0.4	0.5
		(0.104)						
Constrained BV	τ	0.628***	0.488***	0.376***	0.261***	0.143*	0.019	-0.113
		(0.141)	(0.087)	(0.086)	(0.086)	(0.085)	(0.084)	(0.082)
	ρ	-0.127	0	0.1	0.2	0.3	0.4	0.5
		(0.101)						
VI. 2000								
Unconstrained BV	τ	1.010***	0.628***	0.476***	0.321***	0.161	-0.005	-0.180*
		(0.184)	(0.110)	(0.109)	(0.108)	(0.106)	(0.104)	(0.100)
	ρ	-0.258**	0	0.1	0.2	0.3	0.4	0.5
		(0.101)						
Constrained BV	τ	-0.090	0.255***	0.143*	0.031	-0.083	-0.201**	-0.324***
		(0.154)	(0.083)	(0.083)	(0.083)	(0.082)	(0.081)	(0.080)
	ρ	0.306**	0	0.1	0.2	0.3	0.4	0.5
		(0.113)						
VII. 2000 without euro								
Unconstrained BV	τ	1.253***	0.645***	0.500***	0.351***	0.194*	0.029	-0.146
		(0.164)	(0.111)	(0.110)	(0.109)	(0.107)	(0.104)	(0.101)
	ρ	-0.454***	0	0.1	0.2	0.3	0.4	0.5
		(0.095)						
Constrained BV	τ	0.851***	0.544***	0.437***	0.326***	0.209**	0.085	-0.049
		(0.142)	(0.089)	(0.089)	(0.089)	(0.088)	(0.087)	(0.086)
	ρ	-0.297**	0	0.1	0.2	0.3	0.4	0.5
		(0.108)						

Notes: *** p<0.01, ** p<0.05, * p<0.1. Standard errors in parentheses. See text for details. For convergence, curcolony and curcolonizer had to be dropped 1990 onwards, only for the unconstrained BV models. Only curcolony had to be dropped for the 1980 unconstrained BV model. For 1970, the GSP variables had to dropped; curcolony and curcolonizer had to be dropped for 1960; additional variables including common language and religion had to dropped for 1950. For each cross section, within each specification, the set of variables is the same.

Table 7. Currency Unions and Regional Trade Agreements: Bivariate Probit Results with Different Assumptions Concerning Correlation Among the Disturbances

		Unconstrained ρ	Constrained ρ					
I. 1960								
Unconstrained BV	τ	-0.534	-0.615	-0.647	-0.679	-0.712*	-0.749*	-0.791*
		(0.485)	(0.431)	(0.431)	(0.429)	(0.426)	(0.423)	(0.419)
	ρ	-0.232	0	0.1	0.2	0.3	0.4	0.5
		(0.605)						
Constrained BV	τ	-1.074**	-1.021**	-1.048**	-1.075**	-1.101***	-1.129***	-1.159***
		(0.465)	(0.426)	(0.426)	(0.424)	(0.422)	(0.419)	(0.415)
	ρ	0.199	0	0.1	0.2	0.3	0.4	0.5
		(0.727)						
II. 1970								
Unconstrained BV	τ	-1.010***	-0.668***	-0.738***	-0.806***	-0.875***	-0.944***	-1.016***
		(0.183)	(0.144)	(0.145)	(0.145)	(0.144)	(0.143)	(0.142)
	ρ	0.492	0	0.1	0.2	0.3	0.4	0.5
		(0.156)						
Constrained BV	τ	-0.470**	-0.349**	-0.419***	-0.489***	-0.560***	-0.635***	-0.714***
		(0.191)	(0.145)	(0.146)	(0.146)	(0.145)	(0.145)	(0.143)
	ρ	0.173	0	0.1	0.2	0.3	0.4	0.5
		(0.176)						
III. 1980								
Unconstrained BV	τ	1.537***	0.931***	0.830***	0.722***	0.607***	0.483**	0.348*
		(0.377)	(0.200)	(0.199)	(0.197)	(0.194)	(0.191)	(0.186)
	ρ	-0.723	0	0.1	0.2	0.3	0.4	0.5
		(0.459)						
Constrained BV	τ		1.482***	1.386***	1.283***	1.169***	1.043***	0.902***
			(0.194)	(0.194)	(0.192)	(0.190)	(0.187)	(0.183)
	ρ		0	0.1	0.2	0.3	0.4	0.5
IV. 1990								
Unconstrained BV	τ	1.134***	0.690***	0.559***	0.423**	0.279*	0.128	-0.034
		(0.241)	(0.166)	(0.166)	(0.164)	(0.162)	(0.159)	(0.155)
	ρ	-0.370**	0	0.1	0.2	0.3	0.4	0.5
		(0.158)						
Constrained BV	τ	0.683***	0.554***	0.459***	0.362***	0.261**	0.155	0.042
		(0.213)	(0.127)	(0.127)	(0.127)	(0.126)	(0.125)	(0.123)
	ρ	-0.137	0	0.1	0.2	0.3	0.4	0.5
		(0.185)						
V. 2000								
Unconstrained BV	τ	1.090***	0.716***	0.535***	0.348**	0.153	-0.050	-0.264*
		(0.311)	(0.176)	(0.174)	(0.171)	(0.168)	(0.163)	(0.157)
	ρ	-0.215	0	0.1	0.2	0.3	0.4	0.5
		(0.151)						
Constrained BV	τ	-1.071***	-0.072	-0.175*	-0.276***	-0.376***	-0.478***	-0.582***
		(0.137)	(0.101)	(0.101)	(0.101)	(0.101)	(0.101)	(0.101)
	ρ	0.876***	0	0.1	0.2	0.3	0.4	0.5
		(0.056)						
VI. 2000 without euro								
Unconstrained BV	τ	1.273***	0.708***	0.566***	0.417**	0.258	0.090	-0.091
		(0.236)	(0.177)	(0.176)	(0.173)	(0.170)	(0.166)	(0.160)
	ρ	-0.445***	0	0.1	0.2	0.3	0.4	0.5
		(0.135)						
Constrained BV	τ	0.367*	0.341***	0.260**	0.176	0.090	-0.003	-0.103
		(0.200)	(0.116)	(0.116)	(0.117)	(0.117)	(0.116)	(0.115)
	ρ	-0.034	0	0.1	0.2	0.3	0.4	0.5
		(0.205)						

Notes: *** p<0.01, ** p<0.05, * p<0.1. Standard errors in parentheses. See text for details. For convergence, curcolony and curcolonizer had to be dropped 1990 onwards, only for the unconstrained BV models. For 2000, bothin had to be dropped as well. Convergence was extremely hard to achieve for 1980 with the BV constraints but otherwise curcolony, curcolonizer, and samelang had to be dropped. For 1960 and 1970, usually curcolony and curcolonizer had to be dropped. 1960 required additional variables to be dropped. For each cross section, within each specification, the set of variables is the same.

Table A1. List of Trade Agreements

WTO List of Active Agreements*
Andean Community (1989): Bolivia, Colombia, Ecuador, Peru (until 1992 and again from1997), Venezuela.
Association of South East Asian Nations Free Trade Area, or ASEAN FTA (1998): Brunei Darussalam, Cambodia, Indonesia, Lao People's Democratic Republic, Malaysia, Myanmar, Philippines, Singapore, Thailand, Vietnam.
Asia Pacific Trade Agreement, or APTA (1976): Bangladesh, India, Republic of Korea, Lao People's Democratic Republic, Sri Lanka.
Australia - New Zealand Closer Economic Relations (1983)
Australia - Papua New Guinea (1977)
Canada - Chile (1997)
Canada - Israel (1997)
Caribbean Community, or Caricom (1968): Antigua & Barbuda, Bahamas (1983), Barbados, Belize (1971), Dominica, Grenada, Guyana, Jamaica, St. Kitts & Nevis, St. Lucia, St. Vincent & the Grenadines, Suriname (1995), Trinidad & Tobago.
Central American Common Market (1959): Costa Rica (1962), El Salvador, Guatemala, Honduras, Nicaragua.
Chile - Mexico (1999)
Common Market for Eastern and Southern Africa, or Comesa (1994): Angola, Burundi, Comoros, Democratic Republic of Congo, Djibouti (2000), Egypt (2000), Eritrea, Ethiopia, Kenya, Lesotho (until 1997), Madagascar, Malawi, Mauritius, Rwanda, Seychelles (2000), Sudan, Swaziland, Tanzania (until 2000), Uganda, Zambia, Zimbabwe.
Commonwealth of Independent States, or CIS (1994): Armenia, Azerbaijan (1996), Belarus, Georgia (1999), Kazakhstan, Kyrgyz Republic (1995), Moldova, Russia, Tajikistan (1997), Turkmenistan, Ukraine, Uzbekistan.
Costa Rica - Mexico (1995)
East African Community (2000): Kenya, Tanzania, Uganda.
European Union, or EU (1958): Belgium, Luxembourg, France, Italy, Germany, Netherlands, Denmark (1973), Ireland (1973), United Kingdom (UK) (1973), Greece (1981), Portugal (1986), Spain (1986), Austria (1995), Finland (1995), Sweden (1995).
EU - Faroe Islands (1997)
EU - Israel (2000)
EU - Mexico (2000)
EU - Morocco (2000)
EU - Overseas Countries & Territories, or EU - OCT (1971): EU - Anguilla, Aruba, Cayman Islands, Comoros, Djibouti, Falkland Islands, French Polynesia, Greenland, Netherlands Antilles, New Caledonia, St. Helena, St. Pierre & Miquelon, Suriname, Wallis & Futuna Islands.
EU - Palestinian Authority (1997)
EU - South Africa (2000)
EU - Syria (1978)
EU - Tunisia (1998)
EU - Turkey (1996)
Economic and Monetary Community of Central Africa (1999): Cameroon, Central African Republic, Chad, Equatorial Guinea, Gabon, Republic of Congo.
Economic Community of West African States (1993): Benin, Burkina Faso, Cape Verde, C te d 'Ivorie, Gambia, Ghana, Guinea, Guinea Bissau, Liberia, Mali, Mauritania, Niger, Nigeria, Senegal, Sierra Leone, Togo.
Economic Cooperation Organization (1992): Afghanistan, Azerbaijan, Iran, Kazakhstan, Kyrgyz Republic, Pakistan, Tajikistan, Turkey, Turkmenistan, Uzbekistan.
European Free Trade Association, or EFTA (1960): Austria (until 1994), Denmark (until 1973), Finland (1986-1994), Iceland (1970), Norway, Portugal (until 1986), Sweden (until 1994), Switzerland, United Kingdom (until 1973).
EFTA - Israel (1993)
EFTA - Morocco (1999)
EFTA - Palestinian Authority (1999)
EFTA - Turkey (1992)
Eurasian Economic Community (1997): Belarus, Kazakhstan, Kyrgyz Republic, Russia, Tajikistan.
EU - EFTA (1973)
Faroe Islands - Norway (1993)
Faroe Islands - Switzerland (1995)
Georgia - Russia (1994)
Global System of Trade Preferences among Developing Countries (1989): Algeria, Argentina, Bangladesh, Benin, Venezuela, Bolivia, Brazil, Cameroon, Chile, Colombia, Cuba, Ecuador, Egypt, Macedonia, Ghana, Guinea, Guyana, India, Indonesia, Iran, Iraq, Democratic People's Republic of Korea, Republic of Korea, Libya, Malaysia, Mexico, Morocco, Mozambique, Myanmar, Nicaragua, Nigeria, Pakistan, Peru, Philippines, Romania, Singapore, Sri Lanka, Sudan, Tanzania, Thailand, Trinidad and Tobago, Tunisia, Vietnam, Zimbabwe.
Israel - Mexico (2000)
Lao People's Democratic Republic - Thailand (1991)

Notes: The author is unaware of changes made to the WTO database after June 2009. The parentheses contain an agreement's year of entry, except where noted otherwise. If an agreement entered into force prior to 1950, then the start date is mentioned as 1950. Although the European Union (EU) was preceded by the European Communities (EC), and the European Economic Community (EEC), it is referred to as EU throughout.

Overlapping agreements do not find mention. For example, the WTO database includes Armenia - Moldova (1995), but it is captured in CIS.

* See http://rtais.wto.org/UI/PublicMaintainRTAHome.aspx; Frankel (1997), Jovanovic (1998), and RTA secretariat webpages were also consulted.

† Primarily from Frankel (1997); also see http://www.sice.oas.org/agreements_e.asp.

In keeping with the table in Frankel (1997, p. 247) sectoral, partial, and nonreciprocal agreements have also been considered.

Table A1 (cont.). List of Trade Agreements

WTO List of Active Agreements (cont.)*
Latin American Integration Agreement (1981): Argentina, Bolivia, Brazil, Chile, Colombia, Cuba, Ecuador, Mexico, Paraguay, Peru, Uruguay, Venezuela.
Melanesian Spearhead Group (1994): Fiji, Papua New Guinea, Solomon Islands, Vanuatu.
Mercosur (1991): Argentina, Brazil, Paraguay, Uruguay.
Mexico - Nicaragua (1998).
North American Free Trade Agreement, or NAFTA (1994): Canada, Mexico, United States (US).
Pan Arab Free Trade Area (1998): Bahrain, Egypt, Iraq, Jordan, Kuwait, Lebanon, Libya, Morocco, Oman, Qatar, Saudi Arabia, Sudan, Syria, Tunisia, United Arab Emirates, Yemen.
Protocol on Trade Negotiations (1973): Bangladesh, Brazil, Chile, Egypt, Israel, Republic of Korea, Mexico, Pakistan, Paraguay, Peru, Philippines, Romania, Socialist Federal Republic of Yugoslavia, Serbia & Montenegro, Tunisia, Turkey, Uruguay.
South Asian Preferential Trade Arrangement (1995): Bangladesh, Bhutan, India, Maldives, Nepal, Pakistan, Sri Lanka.
South Pacific Regional Trade and Economic Cooperation Agreement (1981): Australia, Fiji, Kiribati, Marshall Islands, Federated States of Micronesia, Nauru, New Zealand, Papua New Guinea, Samoa, Solomon Islands, Tonga, Tuvalu, Vanuatu.
Southern African Development Community (2000): Angola, Botswana, Democratic Republic of Congo, Lesotho, Malawi, Mauritius, Mozambique, Namibia, Seychelles, South Africa, Swaziland, Tanzania, Zambia, Zimbabwe.
Turkey - Macedonia (2000)
Turkey - Israel (1998)
US - Israel (1985)
West African Economic and Monetary Union (2000): Benin, Burkina Faso, C te d 'Ivorie, Guinea Bissau, Mali, Niger, Senegal, Togo.

WTO List of Inactive Agreements*
Arusha Agreement (1971-1976): EU - Kenya, Tanzania, Uganda
Australia - New Zealand (1966-1983)
Bulgaria - Macedonia (2000)
Bulgaria - Turkey (1999)
Canada - US (1989-1994)
Central Europe Free Trade Agreement, or CEFTA (1993): Bulgaria (1998), Czech Republic, Hungary, Poland, Romania (1997), Slovak Republic, Slovenia (1996).
Croatia - Macedonia (1997)
Czech Republic - Estonia (1998)
Czech Republic - Israel (1997)
Czech Republic - Latvia (1997)
Czech Republic - Lithuania (1997)
Czech Republic - Romania (1995-1997)
Czech Republic - Slovak Republic (1993)
Czech Republic - Slovenia (1994-1995)
Czech Republic - Turkey (1998)
EU - Algeria (1978)
EU - Bulgaria (1993)
EU - Cyprus (1973)
EU - Czech Republic (1995)
EU - Egypt (1973)
EU - Estonia (1998)
EU - Faroe Islands (1992-1997)
EU - Finland (1974-1994)
EU - Greece (1962-1981)
EU - Hungary (1994)
EU - Israel (1970-2000)
EU - Jordan (1978)
EU - Latvia (1998)
EU - Lebanon (1973)
EU - Lithuania (1998)
EU - Malta (1971)
EU - Morocco (1969-2000)
EU - Overseas Countries & Territories, or EU - OCT1 (1964-1971): EU - Comoros, French Polynesia, Netherlands Antilles, New Caledonia, St. Pierre & Miquelon, Suriname, Wallis & Futuna Islands.
EU - Poland (1994)
EU - Romania (1995)
EU - Slovak Republic (1995)

Notes: The author is unaware of changes made to the WTO database after June 2009. The parentheses contain an agreement's year of entry, except where noted otherwise. If an agreement entered into force prior to 1950, then the start date is mentioned as 1950. Although the European Union (EU) was preceded by the European Communities (EC), and the European Economic Community (EEC), it is referred to as EU throughout.

Overlapping agreements do not find mention. For example, the WTO database includes Armenia - Moldova (1995), but it is captured in CIS.

* See http://rtais.wto.org/UI/PublicMaintainRTAHome.aspx; Frankel (1997), Jovanovic (1998), and RTA secretariat webpages were also consulted.

† Primarily from Frankel (1997); also see http://www.sice.oas.org/agreements_e.asp.

In keeping with the table in Frankel (1997, p. 247) sectoral, partial, and nonreciprocal agreements have also been considered.

Table A1 (cont.). List of Trade Agreements

WTO List of Inactive Agreements (cont.)*
EU - Slovenia (1993-1999)
EU - Spain (1970-1986)
EU - Tunisia (1969-1998)
EU - Turkey (1964-1996)
EFTA - Bulgaria (1994)
EFTA - Czech Republic (1992)
EFTA - Estonia (1996)
EFTA - Hungary (1994)
EFTA - Latvia (1996)
EFTA - Lithuania (1997)
EFTA - Poland (1994)
EFTA - Romania (1994)
EFTA - Slovak Republic (1992)
EFTA - Slovenia (1995)
EFTA - Spain (1980-1986)
El Salvador - Nicaragua (1951-1959)
Estonia - Faroe Islands (1998)
Baltic Free Trade Area (1994): Estonia, Latvia, Lithuania.
Estonia - Norway (1992-1996)
Estonia - Switzerland (1993-1996)
Estonia - Ukraine (1996)
Faroe Islands - Iceland (1993)
Finland - Bulgaria (1974-1993)
Finland - Czechoslovakia (1974-1992)
Finland - German Democratic Republic (GDR) (1974-1989)
Finland - Hungary Agreement (1974-1993)
Finland - Poland Agreement (1978-1993)
EFTA - Finland (1961-1986)
Gulf Cooperation Council (1983): Bahrain, Kuwait, Oman, Qatar, Saudi Arabia, United Arab Emirates.
Hungary - Israel (1998)
Hungary - Latvia (2000)
Hungary - Lithuania (2000)
Hungary - Slovenia (only 1995)
Hungary - Turkey (1998)
Ireland - United Kingdom (1966-1973)
Latvia - Norway (1992-1996)
Latvia - Switzerland (1993-1996)
Lithuania - Norway (1992-1996)
Lithuania - Switzerland (1993-1996)
Poland - Faroe Islands (1998)
Poland - Israel (1998)
Poland - Latvia (1999)
Poland - Lithuania (1997)
Romania - Moldova (1995)
Romania - Turkey (1998)
Slovak Republic - Estonia (1998)
Slovak Republic - Israel (1997)
Slovak Republic - Latvia (1997)
Slovak Republic - Lithuania (1997)
Slovak Republic - Romania (1995-1997)
Slovak Republic - Slovenia (1994-1995)
Slovak Republic - Turkey (1998)
Slovenia - Croatia (1998)
Slovenia - Estonia (1997)
Slovenia - Macedonia (1996)
Slovenia - Israel (1998)
Slovenia - Latvia (1996)
Slovenia - Lithuania (1997)
South Africa - Zimbabwe (1950-1954)
Trade Expansion and Cooperation Agreement (1968-1983): India, Egypt, Yugoslavia.
Turkey - Estonia (1998)
Turkey - Latvia (2000)

Notes: The author is unaware of changes made to the WTO database after June 2009. The parentheses contain an agreement's year of entry, except where noted otherwise. If an agreement entered into force prior to 1950, then the start date is mentioned as 1950. Although the European Union (EU) was preceded by the European Communities (EC), and the European Economic Community (EEC), it is referred to as EU throughout.

Overlapping agreements do not find mention. For example, the WTO database includes Armenia - Moldova (1995), but it is captured in CIS.

* See http://rtais.wto.org/UI/PublicMaintainRTAHome.aspx; Frankel (1997), Jovanovic (1998), and RTA secretariat webpages were also consulted.

† Primarily from Frankel (1997); also see http://www.sice.oas.org/agreements_e.asp.

In keeping with the table in Frankel (1997, p. 247) sectoral, partial, and nonreciprocal agreements have also been considered.

Table A1 (cont.). List of Trade Agreements

WTO List of Inactive Agreements (cont.)*

Turkey - Lithuania (1998)
Turkey - Poland (2000)
Turkey - Slovenia (2000)
Yaounde I (1964-1971): EU - Benin, Burkina Faso, Burundi, Cameroon, Central African Republic, Chad, Congo, Zaire, Gabon, C te d 'Ivorie, Madagascar, Mali, Mauritania, Niger, Rwanda, Senegal, Somalia, Togo.
Yaounde II (1971-1976): EU - Benin, Burkina Faso, Burundi, Cameroon, Central African Republic, Chad, Congo, Zaire, Gabon, C te d 'Ivorie, Madagascar, Mali, Mauritania, Mauritius (1972), Niger, Rwanda, Senegal, Somalia, Togo.
Lome I (1976-1981): EU - Bahamas, Barbados, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Cape Verde, Central African Republic, Congo, C te d 'Ivorie, Djibouti, Ethiopia, Fiji, Gabon, Equatorial Guinea, Ghana, Gambia, Guinea, Guinea Bissau, Guyana, Jamaica, Kenya, Kiribati, Comoros, Liberia, Lesotho, Madagascar, Mali, Mauritania, Mauritius, Malawi, Niger, Nigeria, Papua New Guinea, Rwanda, Solomon Islands, Seychelles, Sudan, Sierra Leone, Senegal, Somalia, Suriname, Sao Tome & Principe, Swaziland, Chad, Togo, Tonga, Trinidad & Tobago, Tuvalu, Tanzania, Uganda, Dominica, Grenada, St. Lucia, Samoa, St. Vincent & Grenadines, Zambia, Zaire.
Lome II (1981-1986): EU - Bahamas, Barbados, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Cape Verde, Central African Republic, Congo, C te d 'Ivorie, Djibouti, Ethiopia, Fiji, Gabon, Equatorial Guinea, Ghana, Gambia, Guinea, Guinea Bissau, Guyana, Jamaica, Kenya, Kiribati, Comoros, Liberia, Lesotho, Madagascar, Mali, Mauritania, Mauritius, Malawi, Niger, Nigeria, Papua New Guinea, Rwanda, Solomon Islands, Seychelles, Sudan, Sierra Leone, Senegal, Somalia, Suriname, Sao Tome & Principe, Swaziland, Chad, Togo, Tonga, Trinidad & Tobago, Tuvalu, Tanzania, Uganda, Dominica, Grenada, St. Lucia, Samoa, St. Vincent & Grenadines, Zambia, Zaire, Antigua & Barbuda, Belize, St. Kitts & Nevis, Vanuatu, Zimbabwe (1982).
Lome III (1986-1991): EU - Bahamas, Barbados, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Cape Verde, Central African Republic, Congo, C te d 'Ivorie, Djibouti, Ethiopia, Fiji, Gabon, Equatorial Guinea, Ghana, Gambia, Guinea, Guinea Bissau, Guyana, Jamaica, Kenya, Kiribati, Comoros, Liberia, Lesotho, Madagascar, Mali, Mauritania, Mauritius, Malawi, Niger, Nigeria, Papua New Guinea, Rwanda, Solomon Islands, Seychelles, Sudan, Sierra Leone, Senegal, Somalia, Suriname, Sao Tome & Principe, Swaziland, Chad, Togo, Tonga, Trinidad & Tobago, Tuvalu, Tanzania, Uganda, Dominica, Grenada, St. Lucia, Samoa, St. Vincent & Grenadines, Zambia, Zaire, Antigua & Barbuda, Belize, St. Kitts & Nevis, Vanuatu, Zimbabwe, Angola, Mozambique.

Agreements not listed on the WTO database †

Lome IV (1991): EU - Bahamas, Barbados, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Cape Verde, Central African Republic, Congo, C te d 'Ivorie, Djibouti, Ethiopia, Fiji, Gabon, Equatorial Guinea, Ghana, Gambia, Guinea, Guinea Bissau, Guyana, Jamaica, Kenya, Kiribati, Comoros, Liberia, Lesotho, Madagascar, Mali, Mauritania, Mauritius, Malawi, Niger, Nigeria, Papua New Guinea, Rwanda, Solomon Islands, Seychelles, Sudan, Sierra Leone, Senegal, Somalia, Suriname, Sao Tome & Principe, Swaziland, Chad, Togo, Tonga, Trinidad & Tobago, Tuvalu, Tanzania, Uganda, Dominica, Grenada, St. Lucia, Samoa, St. Vincent & Grenadines, Zambia, Zaire, Antigua & Barbuda, Belize, St. Kitts & Nevis, Vanuatu, Zimbabwe, Angola, Mozambique, Dominican Republic, Haiti, Namibia.
European Coal and Steel Community (1951-1957): Belgium, France, Germany, Italy, Luxembourg, Netherlands.
Canada - US Automotive Agreement (1965-1989)
Mercosur - Bolivia (1996)
Mercosur - Chile (1996)
Bolivia - Mexico (1995)
Customs Union of West African States (1959-1966): Benin, Burkina Faso, C te d 'Ivorie, Mali, Mauritania, Niger, Senegal.
West African Economic Community (1966-1994): Benin, Burkina Faso, C te d 'Ivorie, Mali, Mauritania, Niger, Senegal.
Equatorial Customs Union (1959-1966): Cameroon(1961), Central African Republic, Chad, Congo, Gabon.
Economic and Customs Union of the Central African States (1966-1985): Cameroon, Central African Republic, Chad, Congo, Equatorial Guinea(1985), Gabon.
Group of Three (1995): Colombia, Mexico, Venezuela.
Mano River Union (1973): Guinea, Liberia, Sierra Leone.
Arab Maghreb Region (1991): Algeria, Libya, Mauritania, Morocco, Tunisia.
Caricom - Colombia (1995)
Comecon (1950-1991): Bulgaria, Czechoslovakia, Hungary, Poland, Romania, Soviet Union, Albania, GDR (1950), Mongolia (1962), Cuba (1972), Vietnam (1978).
Caribbean Basin Initiative (1983): US - Antigua & Barbuda, Bahamas, Barbados, Belize, Costa Rica, Dominica, Dominican Republic, El Salvador, Grenada, Guatemala, Guyana, Haiti, Honduras, Jamaica, Netherlands Antilles, Nicaragua, Panama, St. Kitts & Nevis, St. Lucia, St. Vincent & Grenadines, Trinidad & Tobago.
Andean Trade Preference Act (1991): US - Bolivia, Colombia, Ecuador, Peru (1993).
Caribcan (1986): Canada - Anguilla, Antigua & Barbuda, Bahamas, Barbados, Bermuda, Cayman Islands, Dominica, Grenada, Jamaica, St. Kitts & Nevis, St. Lucia, St. Vincent & Grenadines, Trinidad & Tobago, Belize, Guyana.
Southern African Customs Union (1950*): Botswana, Lesotho, Namibia, South Africa, Swaziland.
Cross Border Initiative (1993): Burundi, Comoros, Kenya, Madagascar, Malawi, Mauritius, Namibia, Rwanda, Seychelles, Swaziland, Tanzania, Uganda, Zambia, Zimbabwe.
ASEAN (1978-1997): Brunei Darussalam (1984), Indonesia, Malaysia, Philippines, Singapore, Thailand, Vietnam (1995).
East African Common Market (1967-1979): Kenya, Tanzania, Uganda.
Indian Ocean Commission (1984): Comoros; Madagascar; Mauritius; Seychelles.

Notes: The author is unaware of changes made to the WTO database after June 2009. The parentheses contain an agreement's year of entry, except where noted otherwise. If an agreement entered into force prior to 1950, then the start date is mentioned as 1950. Although the European Union (EU) was preceded by the European Communities (EC), and the European Economic Community (EEC), it is referred to as EU throughout.
Overlapping agreements do not find mention. For example, the WTO database includes Armenia - Moldova (1995), but it is captured in CIS.
* See http://rtais.wto.org/UI/PublicMaintainRTAHome.aspx; Frankel (1997), Jovanovic (1998), and RTA secretariat webpages were also consulted.
† Primarily from Frankel (1997); also see http://www.sice.oas.org/agreements_e.asp.
In keeping with the table in Frankel (1997, p. 247) sectoral, partial, and nonreciprocal agreements have also been considered.

Table A2. List of Currency Unions

I.
Antigua & Barbuda, Barbados (1975), Dominica, Grenada, Guyana (1971), St. Kitts & Nevis, St. Lucia, St.Vincent & Grenadines, Trinidad & Tobago (1976).
II.
Aruba, Netherlands Antilles, Suriname (1994).
III.
Australia, Kiribati, Nauru, Solomon Islands (1979), Tonga (1991), Tuvalu.
IV.
Belgium, Luxembourg, Burundi (1964), Democratic Republic of Congo (1961), Rwanda (1966).
V.
Spain - Equatorial Guinea (1969)
VI.
Benin, Burkina Faso, C te d 'Ivorie, Equatorial Guinea (from 1985), Gabon, Guinea (1969), Guinea Bissau (from 1997), Madagascar (1981), Mali (upto 1961 and from 1984), Mauritania (1974), Niger, Reunion (1976), Senegal, Togo, Cameroon, Central African Republic, Chad, Comoros (1994), Republic of Congo, St. Pierre & Miquelon (1976).
VII.
India, Bangladesh (1974), Bhutan, Burma (1966), Pakistan (1966), Maldives (1966), Mauritius (1966), Mauritius (1966), Seychelles (1966), Kuwait (1961), Oman (1970), Qatar (1966), Sri Lanka (1966) People's Democratic Republic of Yemen (1951).
VIII.
Pakistan, Burma (1971), Maldives (1971), Mauritius (1967), Seychelles (1967), Sri Lanka (1967).
IX.
Mauritius - Seychelles (1976)
X.
Qatar - United Arab Emirates (from 1981)
XI.
Denmark, Faeroe Islands, Greenland.
XII.
France, Algeria (1969), French Guiana, Guadeloupe, Martinique, Morocco (1959), Tunisia (1958), Reunion (from 1976), St. Pierre & Miquelon (from 1976).
XIII.
United Kingdom, Bahamas (1966), Bermuda (1970), Cyprus (1972), Falkland Islands, Ghana (1965), Gibraltar, Iraq (1967), Ireland (1979), Israel (1954), Jamaica (1969), Jordan (1967), Kenya (1967), Kuwait (from 1961-1967), Libya (1967), Malawi (1971), Malta (1971), New Zealand (1967), Nigeria (1967), Samoa (1967), Somalia (1967), Tanzania (1967), Uganda (1967), Zambia (1967), Zimbabwe (1967), Gambia (1971), Sierra Leone (1965), South Africa (1961), People's Democratic Republic of Yemen (from 1953 - 1972), St. Helena, Oman (only 1970).
XIV.
United States, Bahamas (from 1966), Bermuda (from 1970), Dominican Republic (1985), Guatemala (1986), Liberia, Panama, Guam, American Samoa, East Timor (only 2000), Ecuador (only 2000).
XV.
Singapore, Brunei, Malaysia (1971).
XVI.
Portugal, Angola (1976), Cape Verde (1977), Guinea Bissau (1977), Mozambique (1977), Sao Tome and Principe (1977).
XVII.
New Caledonia, French Polynesia, Vanuatu (1971), Wallis and Futuna.
XVIII.
Austria (from 1999), Belgium (from 1999), Luxembourg (from 1999), Finland (from 1999), France (from 1999), Germany (from 1999), Ireland (from 1999), Italy (from 1999), Netherlands (from 1999), Portugal (from 1999), Spain (from 1999), and St. Pierre & Miquelon (from 1999).

Notes: The parentheses contain a currency union's year of exit, except where noted otherwise. While Glick and Rose (2002) is the main source consulted, additional sources include Bogetić (2000), Friberg and Matha (2004), Schuler (2005), Rose (2006), Gómez-Oliver et al. (1999), Boughton (1991), Rose and Engel (2002), and Moheeput (2008).