WP/09/186

IMF Working Paper

One Money, One Market—A Revised Benchmark

Theo S. Eicher and Christian Henn

INTERNATIONAL MONETARY FUND

 WP/09/186

IMF Working Paper

Strategy, Policy, and Review Department

One Money, One Market—A Revised Benchmark

Prepared by Theo Eicher and Christian Henn[1]

Authorized for distribution by Thomas Dorsey

September 2009

Abstract

The introduction of the euro generated substantial interest in measuring the impact of currency unions (CUs) on trade flows. Rose's (2000) initial estimates suggested a tripling of trade and created a literature in search of "more reasonable" CU effects. A recent meta-analysis of this literature shows that subsequent papers quantify CU trade impacts at 30–90 percent. However, most recent studies use shorter time series and fewer countries than Rose in his original work. We revisit Rose's original benchmark, extend the dataset, and address Baldwin's (2006) critiques regarding the proper specification of gravity models in large panels by simultaneously accounting for multilateral resistance and unobserved bilateral heterogeneity. This produces a robust average CU trade effect of 45 percent. Yet, the trade impacts of individual CUs vary substantially and are generally lower than those of preferential trade agreements (PTAs). Our revised benchmark can be used as a yardstick for future studies to delineate how estimates differ due to new data or differences in econometric specifications.

JEL Classification Numbers: F10, F15, C23

Keywords: Currency Unions, Preferencial Trade Agreements, Gravity Equation, Multilateral Resistance

Author's E-Mail Address: te@u.washington.edu, chenn@imf.org

[1] Theo Eicher, Department of Economics, University of Washington; and Christian Henn, Strategy, Policy, and Review Department, IMF. We thank Subramanian and Wei (2007) for sharing their trade data. Christian Henn acknowledges support from the Henry Buechel Memorial Fellowship.

Contents Page

I. Introduction

The advent of the euro awoke keen interest in whether a currency union (CU) generates trade benefits over and above those of eliminating exchange rate fluctuations. If trade relationships are costly to establish, a more durable CU commitment should yield additional trade benefits compared to a conventional fixed rate peg. It is important to quantify these trade benefits for two reasons. First, countries outside of CUs need to know how much extra gains from trade their consumers can expect in deciding whether it is worthwhile to abandon independent monetary policy and thereby possibly incur greater volatility in output and inflation (Karam et al, 2008). Second, higher trade makes a CU more resilient through more integrated business cycles among member countries.

Rose (2000), in a seminal paper, was first to empirically test for a CU trade effect. He found that, on average over time, CUs double or even triple bilateral trade between members. And because the CU effects' magnitude typically increases over time (e.g. Flam and Nordstrom, 2003), presumably trade creation would be even larger after a CU is well established. This notion of the tripling estimate being unreasonably high is further reinforced by a look at the raw trade data. For instance, since euro introduction, German-Irish trade has increased by only 30 percentage points more than German-British trade.[2]

Thus, it is not surprising that Rose's estimate sparked a controversy out of which emerged an entire literature attempting to "shrink the Rose effect." This literature is meta-analyzed by Rose and Stanley (2005), who report that subsequent papers find much smaller changes in trade volumes, usually around 30–90 percent. However, these recent papers used much smaller datasets over shorter time series than Rose (2000). For large panel datasets, Rose and Stanley still report trade gains exceeding 100 percent (confirmed by the latest large panel study of Frankel, 2008). Thus, recent literature shows that the CU trade effect's magnitude has not been settled and that dataset dimensions and econometric approaches profoundly influence results.

Baldwin (2006) provides a comprehensive survey of econometric approaches used in the CU literature and suggests two crucial sets of controls necessary to obtain unbiased CU trade effects from the gravity equation. Baldwin and Taglioni (2006) implement these controls in a small panel to find either negative or zero trade effects of the euro.[3] Their results highlight

---

[2] The growth rates were calculated for 1998 to 2008 from the IMF's Direction of Trade Statistics.

[3] Baldwin and Taglioni (2006) focus solely on trade effects of the euro. Hence with 4,837 observations, their dataset is much smaller than Rose's (2000), who featured 22,948 observations, and ours (76,081 observations). Baldwin and Taglioni speculate that the implausible negative effect is the result of insufficient cross-sectional variation. However, when they add data (back to 1980) to address the high standard errors, their euro coefficient is small, positive and insignificant.

estimates' sensitivity with respect to the suggested sets of controls, but do not resolve what the implied CU or euro impact may be in large datasets. Frankel (2008) revisits Rose (2000) in a large panel but controls only for the second of two elements in Baldwin's critique. Here we provide a revised benchmark for CU trade effects by simultaneously addressing the key methodological issues raised by Baldwin (2006) in an updated and extended version of Rose's (2000) large dataset.

Baldwin's (2006) first fundamental insight was that multilateral resistance (Anderson and van Wincoop, 2003) must be comprehensively accounted for. Multilateral resistance captures the notion that trade decisions are based on relative, rather than absolute, prices. Two countries' decisions of how much to trade with each other is not only affected by the bilateral trade costs between them, but also the average (or multilateral) trade costs faced by each of these countries.[4] Because multilateral trade costs are an average of bilateral trade costs, they are affected by any factors that change the latter, such as geographical location, transit connections, tariff regime etc. As many of these determinants change from year to year, multilateral resistance thus varies not only by country but also over time. Therefore time-varying country fixed effects are required to comprehensively control for multilateral resistance in panel datasets. Previous approaches to controlling for multilateral resistance have focused on geography only with a remoteness measure (Rose, 2000) or used time-*in*variant country fixed effects (Rose and van Wincoop, 2001).[5] The latter approach acknowledges that various determinants matter for a country's average trade cost, but also assumes that a country's average trade costs with the rest of the world remain constant over time. Below we outline theoretically and empirically how coefficients are affected by omitted variable bias, if comprehensive multilateral resistance controls are absent from the analysis.

Baldwin's second issue is that further omitted variable bias may result when the empirical strategy does not account for unobserved determinants of bilateral trading relationships. Hummels and Levinsohn (1995) first emphasized this unobserved bilateral heterogeneity by including country-pair fixed effects in the estimation. Recent papers on currency regimes and trade that employ a similar approach include Glick and Rose (2002), Pakko and Wall (2001), Baldwin and Taglioni (2006), Klein and Shambaugh (2006) and Frankel (2008). Failure to include the adequate fixed effect controls can lead to such severe bias that Baldwin (2006) recommends ignoring any other estimates for policy purposes. While the above cited papers address either multilateral resistance or unobserved bilateral heterogeneity, only Baldwin and Taglioni address both—but, as mentioned, in a much smaller panel without overlap with

---

[4] For instance, the distances between Australia and New Zealand, on the one hand, and Spain and Poland, on the other hand, are roughly equal. However, Australia-New Zealand trade is substantially higher than Spain-Poland trade, because average trade costs (=multilateral resistance) faced by Australia and New Zealand are quite high owing to their remote geographical location.

[5] Time-varying fixed effects have since been introduced to the gravity literature, for example, in Subramanian and Wei (2007) in the context of WTO trade effects.

Rose (2000). In this paper, we implement both methodological approaches simultaneously in a long panel covering 10 cross sections over 50 years and 177 countries.

In addition, we address another crucial issue that is underemphasized in the CU literature: individual CUs and preferential trade agreements (PTAs) produce widely varying trade effects as different as their member country groupings. With exception of Nitsch (2002), this heterogeneity is not addressed in the CU literature. In the presence of such heterogeneity, we show that average CU and PTA effects captured by single "catch-all" dummies generate biased and uninformative results. Thus, for policy purposes, there exists no single CUs trade effect; and this must be addressed by the empirical strategy. For instance, CU trade effects for the euro and the African CFA Franc are unlikely to be equal, given different average development levels of their members. Similarly, it is crucial to allow for separate effects of multilateral and unilateral ("hub and spoke") CUs.

Our results show that it is crucial to account for all three outlined shortcomings simultaneously to eliminate bias to CU trade effects. Rose's (2000) average CU trade effect remains statistically and economically significant, although we find it reduced to a more realistic 45 percent. However, our results indeed confirm strong heterogeneity in PTA and CU trade effects. In contrast to Baldwin and Taglioni's (2006) result of no euro effect, we find a statistically significant 40 percent trade increase. Furthermore, our simultaneous account of multilateral resistance and unobserved bilateral heterogeneity conveys a 100 percent trade effect of the African CFA franc. On the contrary, the multilateral East Caribbean CU is never found to have a trade effect. Hub and spoke CUs featuring the British pound and the US dollar generally do not boost trade between spokes and the hub. Thus, and in contrast to Glick and Rose (2002) and Frankel and Rose (2002), we find dollarization to be insignificant for trade (as reported by Klein, 2005). Generally, trade effects of PTAs are greater than those of CUs. The reverse is true only in Europe: There we find the euro to boost trade by 40 percent, while the EU increases trade by only 25 percent.

The remainder of the paper is organized as follows. Our dataset is presented in Section 2. Section 3 reviews the Baldwin (2006) critique of gravity methodology. Sections 4 and 5, sequentially incorporate multilateral resistance and unobserved bilateral heterogeneity. Section 6 presents extensive robustness analysis. Section 7 concludes.

II. Data

Our dataset is an expanded version of Subramanian and Wei (2007). Subramanian and Wei (2007) in turn base their data on Rose (2004). The dataset ranges from 1950 to 2000 and represents a significant expansion of Rose's (2000) 1970–1990 data. Rose (2000) featured 22,948 observations (330 in CUs); we have 76,081 observations (1,224 in CUs) in 16,941 bilateral trade relationships across 177 countries (see Appendix Tables A1-A2).[6] The additional observations are crucial, because they enable us to introduce extensive fixed effects without compromising estimation precision.

Our dependent variable is bilateral imports at five-year intervals, deflated by the U.S. consumer price index.[7] A number of CU studies employ the *average* of imports and exports as the dependent variable, to reduce measurement error (e.g. Rose, 2000; Rose and van Wincoop, 2001; Glick and Rose, 2002). Recent approaches favor our unidirectional trade data, which is more closely aligned with theoretical implications and allows for proper multilateral resistance controls.

We expand the original Subramanian and Wei (2007) dataset to include a comprehensive set of explanatory variables suggested by previous literature. First, we augment the dataset to include a large list of major PTAs obtained from Ghosh and Yamarik (2004). Second, we add information on individual CUs as reported by Glick and Rose (2002). Third, we update the CU variable to include more recent CUs. Fourth, we include a currency board (CB) dummy and split it into arrangements that peg to the US Dollar ($CBusd_{mxt}$) and the D-Mark/Euro ($CBeuro_{mxt}$). Appendix Tables A2-A4 summarize the membership in CUs, CBs, and PTAs. Fifth is the addition of controls that are frequently encountered in the CU literature, which include current/historical colonial relationships as well as common languages/territories/borders. Sixth, we include regressors to control for differences in factor endowments (absolute log differences in per capita GDP and population density), based on the Penn World Tables, version 6.2. Finally, we add bilateral exchange rate volatility, which is computed from the IMF International Financial Statistics using Ghosh and Yamarik's (2004) methodology (the standard deviation of the first difference in the bilateral exchange rate in the previous 3 years). Regressions including FX volatility reduce the dimension of the dataset to 66,619 observations in 15,833 pairs starting in 1960.

III. Empirical Implementation of the Gravity Model

Baldwin (2006) leveled two fundamental critiques against popular empirical implementations of the gravity equation. His arguments are best understood by following a theory-based derivation of the gravity equation based on Anderson (1979) and Anderson and van Wincoop (2003). Baldwin (2006) starts with the trade expenditure share identity to derive a version of the gravity equation that relates bilateral imports, V_{mxt}, at time t to expenditures, E, of importers, m, and exporters, x:

[6] The reason for our larger number of observations rests on the longer panel employing unidirectional trade data.

[7] Deflating the trade data by the U.S. consumer price index is common in the literature, given that trade price indices are unavailable for many countries and years. Any possible bias induced is picked up by the time-varying importer and exporter fixed effects (Baldwin and Taglioni, 2006).

$$V_{mxt} = \frac{\tau_{mxt}^{1-\sigma} E_{mt} E_{xt}}{\Delta_{mt} \Omega_{xt}}. \tag{1}$$

The numerator illustrates that "size" of trading partners (proxied by E_m *or* E_x) "attracts" more bilateral trade, akin to Newton's Law of Gravity. Greater bilateral trade costs, τ_{mxt}, on the other hand, reduce bilateral imports (as $\sigma>1$ for substitutes). The denominator contains multilateral resistance terms for exporters and importers that represent these countries' openness to the rest of the world. Formally, $\Delta_{mt} \equiv \sum_k n_{kt} \tau_{mkt}^{1-\sigma}$ is the importer's trade costs with k global trading partners for n varieties, while the global cost/demand index for the exporter nation is $\Omega_{xt} = \tau_{xt}^{1-\sigma} E_{kt} / \Delta_{kt}$.

Equation 1 clearly shows that both changes in bilateral trade costs (for example, countries m and x join a CU) and changes in multilateral trade costs (e.g. country k changes tariffs across the board) affect the bilateral trade relationship, V_{mxt}, in general equilibrium. *Time-varying* multilateral resistance controls are thus necessary to avoid bias. Otherwise changes in multilateral trade costs may be falsely attributed to changes in bilateral relationships (e.g., formation of a CU). Feenstra (2002) argues that time-varying fixed effects are the method of choice to control for multilateral resistance in large panels for which the relevant cost indices are unavailable. Baldwin (2006) makes the same point in a currency-union-specific context.

Bilateral trade cost can be disaggregated to highlight its individual determinants:

$$\tau_{mxt}^{1-\sigma} = F\left[Distance_{mx}, CU_{mxt}, CB_{mxt}, PTA_{mxt}, Z_{mxt}\right]. \tag{2}$$

Aside from transport costs (proxied by distance), currency arrangements, and preferential trade agreements, trade costs are determined by a vector of regressors, Z_{mxt}, that controls for countries' "natural" inclinations to trade with each other. Variables commonly included in Z_{mxt} are bilateral exchange rate volatility, $FXvola_{mxt}$; current and historical colonial relationships, $CurColony_{mxt}$ and $EverColony_{mx}$, respectively; common colonizer post-1945, $ComColonizer_{mx}$; shared official languages, $ComLang_{mx}$; as well as territorial dependencies and contingencies, $ComNat_{mx}$ and $Border_{mx}$, respectively.

It is difficult to specify an exhaustive Z_{mxt} vector, since some bilateral characteristics may be unobservable.[8] This is the origin of Baldwin's (2006) second criticism: whenever Z_{mxt} is not comprehensively specified, the gravity equation is immediately subject to omitted variable bias. Therefore, the gravity equation must contain not only time-varying importer and exporter fixed effects but also country-pair fixed effects, which control for all unobservables in bilateral trade relationships. The absence of pair fixed effects is not usually due to oversight on the part of the researcher. Especially in the CU literature, the paucity of

[8] For example, personal relationships between business leaders, transport infrastructure, political relationships, cultural affinities, and institutional similarities.

observations entering/exiting CUs may render the introduction of these effects too restrictive in small datasets. Our dataset proves sufficiently large to provide significant results.

The third methodological aspect addressed by us relates to the distinct trade effects of individual CUs and PTAs. If PTAs and CUs do not generate identical trade benefits, estimating an average coefficient using a catch-all CU or PTA dummy introduces bias not only to bilateral trade costs (Equation 2) but also to the multilateral resistance terms (Equation 1). A large literature has documented that trade effects of individual PTAs and CUs differ substantially.[9] Hence, we allow not only for individual PTAs but also examine results for individual CUs.

IV. Multilateral Resistance and the Trade Effects of Currency Unions

Our empirical strategy proceeds in stages. We first introduce controls for multilateral resistance; later we then include the additional fixed effects to address unobserved bilateral heterogeneity. This sequential approach allows us to examine the marginal impact of each set of controls on the CU coefficients.

Multilateral resistance controls have long been part of the CU literature. Rose (2000) included a time-invariant "remoteness" term to proxy for multilateral resistance. Rose and van Wincoop (2001) included country-specific fixed effects and reduced Rose's (2000) CU trade effect from 235 percent to 136 percent in the process. The Rose and van Wincoop (2001) strategy sufficiently addresses multilateral resistance in a cross-section; however, it does not capture the time-varying nature of trade costs in panel data. Baldwin and Taglioni (2006) address this issue by including time-varying fixed effects but find either zero or negative trade effects of the Euro in a small dataset. Here we establish a new revised benchmark for a large panel by estimating equations (1) and (2) according to

$$\begin{aligned}\log(Imports_{mxt}) = \alpha + \delta_{mt} + \lambda_{xt} + \beta_1 CU_{mxt} + \beta_2 CB_{mxt} + \beta_3 PTA_{mxt} + \beta_4 FXvola_{mxt} \\ + \beta_5 CurColony_{mxt} + \beta_6 EverColony_{mx} + \beta_7 ComColonizer_{mx} \\ + \beta_8 ComLang_{mx} + \beta_9 ComNat_{mx} + \beta_{10} Border_{mx} + \beta_{11} Distance_{mx} + \varepsilon_{mxt}\end{aligned} \quad . \qquad (3)$$

Equation (3) includes time-varying fixed effects for importers, δ_{mt}, and exporters, λ_{xt}, to address multilateral resistance. Note that these fixed effects absorb country-year specific regressors, such as importer and exporter expenditures, E_{mt} and E_{xt}, which are proxied by GDP in canonical gravity equations. Equation (3) is easily extended to account for individual CUs, CBs, and PTAs by converting β_1, β_2, and β_3 to coefficient vectors $\tilde{\beta}_1$, $\tilde{\beta}_2$, and $\tilde{\beta}_3$ representing membership in individual arrangements.

[9] See Frankel (1997), Soloaga and Winters (2001), Carrere (2006), Eicher, Henn and Papageorgiou (2007), Rose (2004 and 2005), Subramanian and Wei (2007), Nitsch (2002) and Eicher and Henn (2008).

Regressions 1–3 in Table 1 present our baseline results for CU trade effects with multilateral resistance controls. Regression 1 can be directly compared to Rose's (2000) benchmark regression except for the addition of multilateral resistance controls.[10] At 0.65, the CU coefficient estimate is roughly 6 standard deviations lower than Rose's original 1.21. This reduces the CU trade increase to 91 percent $\left(\approx e^{0.648}-1\right)$ as opposed to Rose's tripling estimate (the 235 percent increase). The estimate is also significantly smaller than Rose and van Wincoop's (2001), who did not consider the time-varying nature of multilateral resistance. Their estimate of 0.86 (implying a 136 percent increase) settles right between ours and Rose's (2000).

Regressions 2 and 3 allow for individual CU and PTA effects. Regression 2 first introduces all PTAs included in Rose's (2000) PTA dummy; then Regression 3 expands the set of PTAs to those considered by Ghosh and Yamarik (2004). One reason put forth to exclude individual PTAs from CU studies is that CU and PTA membership may overlap, particularly in Europe (see e.g., Frankel, 2008). This overlap, however, does not justify their exclusion. Rather, by the very same reasoning, the exclusion of individual PTAs introduces omitted variable bias to CU estimates. Even if CU and PTA membership generated multicorrelation, and therefore the standard errors of PTAs and CUs were inflated, coefficients resulting from their simultaneous inclusion are nevertheless the best linear unbiased estimates. In our dataset, we find that potentially inflated standard errors are not a serious problem for statistical significance. Most of the individual CUs and PTAs are estimated with sufficient precision to infer statistical significance even when included in tandem.

Regressions 2 and 3 show the importance of splitting the catch-all CU dummy into the individual CU arrangements. Individual CU trade effects differ substantially from each other and from the average trade effect estimated in Regression 1. Consequently, individual CUs improve fit considerably throughout: Convincing evidence is provided by the relevant F-Statistics, and by CU and other estimates' robustness and significance across specifications.

Large and significant effects for individual CUs exist for the African CFA and for (mostly extinct) hub-spoke arrangements represented by $CUother_{mxt}$. Regressions 2 and 3 show that African CFA franc internal trade is estimated to be 197–224 percent higher than trade with outsiders. The hub-spoke arrangements of $CUother_{mxt}$ show a similar trade increase of 157–183 percent. CUs involving the British Pound, US dollar, and East Caribbean dollar show no statistically significant effects.

---

[10] As outlined in the data section, further differences lie in (1) the specification of the dependent variable (unidirectional trade flow data, vs. Rose's bidirectional), (2) time frame (1950–2000 vs. 1970–1995 in Rose), and (3) one additional regressor (we insert a currency board dummy, which has, however, no impact on the results).

The trade effect of the euro is the surprise in this set of results. In Regression 2, our estimated euro trade increase $\left(46\% \approx e^{0.381} - 1\right)$ is substantially smaller than effects of other CUs and the CFA in particular. Moreover, the euro effect even turns insignificant when the European Economic Area (EEA) is included (Regression 3). The formation of the EEA in 1994 extended the EU's Common Market to most members of the European Free Trade Agreement (EFTA) and deepened European trade integration. Regressions 3 suggests that subsequent trade flows were mainly affected by PTA-based integration and hardly by the formation of the eurozone. These results underline the importance of including a comprehensive set of individual PTA dummies when estimating CU effects.

A counterintuitive result in Regression 3 is *negative* trade creation of the main European PTA—the EU. The EU instituted far-reaching integration by removing border controls and harmonizing the entire spectrum of public policy; the resulting reduction in transaction costs should have augmented trade volumes.

This predicted negative EU effect, however, is well understood in the literature (see e.g., Linnemann, 1966; Aitken, 1973; Pollak, 1996; Rose, 2004; Baldwin 2006). Dating back to Linnemann (1966), the gravity equation has been known to systematically over-predict trade among large, geographically proximate country pairs. Europe-specific variables thus tend to pick up the negative residuals resulting from proximate European countries' under-trading relative to gravity model predictions. Since the EU variable most closely resembles a Europe dummy, its coefficient turns negative in Regressions 2 and 3. This negative coefficient indicates the omission of crucial variables that would help the gravity equation predict intra-European trade correctly. This omission is not surprising: because the flaw in the gravity specification relates to unobserved effects specific to country *pairs*, multilateral resistance controls cannot remedy the issue. That is, the negative EU effect alerts us that the empirical approach is missing crucial unobserved bilateral heterogeneity controls. We add these controls in Section 5.

V. Benchmark CU Trade Effects addressing Multilateral Resistance and Unobserved Bilateral Heterogeneity

In this section, we add country-pair fixed effects to control for any relevant unobservables in bilateral trade relationships. The estimates presented in this section thus account for the most comprehensive set of controls for omitted variable bias and are the most policy relevant. As outlined in the introduction, *either* multilateral *or* unobserved heterogeneity among trading partners has been addressed by previous CU papers. Here we account for both effects simultaneously to provide a revised benchmark of Rose's (2000) results. In a CU context, only Baldwin and Taglioni (2006) have undertaken such a simultaneous approach before—on a small dataset of roughly 4,000 recent observations (that does not overlap with Rose, 2000). The size of the dataset matters because the inclusion of comprehensive fixed effects reduces the number of degrees of freedom substantially. By adding country-pair fixed effects to equation (3), we obtain our new estimation equation:

$$\log(Imports_{mxt}) = \alpha_{mx} + \delta_{mt} + \lambda_{xt} + \beta_1 CU_{mxt} + \beta_2 CB_{mxt} + \beta_3 PTA_{mxt} + \beta_4 FXvola_{mxt} \\ + \beta_5 CurColony_{mxt} + \varepsilon_{mxt} \quad . \quad (4)$$

All time-invariant pair specific variables are now absorbed into the pair fixed effects, α_{mx}.

In large trade datasets, the estimation of three-way fixed effect structures as in equation (4) is computationally demanding.[11] Despite the growing interest of labor economists in analyzing three-way error component models since Abowd et al. (1999), only three papers exploit this setup in a gravity context aside from Baldwin and Taglioni (2006). Baltagi et al. (2003) also provide strong economic and statistical arguments in favor of our proposed three-way error components model. They do not motivate the time-varying importer and exporter dummies with omitted price terms but with country-specific political and institutional conditions, and business cycles. Eicher and Henn (2007) exploit the methodology in a large dataset to test for the trade implications of regionalism and multilateralism. Baier and Bergstrand (2007) chose the three-way structure as their preferred technique to address possible endogeneity problems.

Regressions 4–6 in Table 2 present the estimates based on equation (4). The F-Statistics overwhelmingly confirm the importance of country-pair fixed effects. Moreover, Regression 4 already reveals that we previously attributed much of "naturally" occurring trade to CUs. At 53 percent $\left(\approx e^{0.42} - 1\right)$, the average CU effect has about halved and differs by more than two standard deviations from our previous estimate of 91 percent (Regression 1). The 53 percent estimate is statistically significant but dramatically lower than the 120 percent reported by Glick and Rose (2002, Table 5). Their paper features country-pair fixed effects but no time-varying multilateral resistance controls.

By disaggregating CUs and PTAs in Regressions 5 and 6, we find that individual CU estimates are significantly reduced compared to Regressions 2 and 3. The exception is again the trade effect of the euro. It turns positive now after accounting for unobserved bilateral heterogeneity and will be discussed further below. Again we show that catch-all dummies masked highly heterogeneous individual CU and PTA effects. The estimates for hub-spoke CUs involving the British Pound or U.S. Dollar remain insignificant. The African CFA and Other (extinct) hub-spoke CUs, on the other hand, stay significant but show reduced trade impact. In percentage terms, their effects halve to 97 and 73 percent, respectively. Overall, the country-pair fixed effects cause a slight reduction in estimates' precision, because Regressions 4–6 exploit only the time dimension. That is, the CU coefficients in Regressions

[11] This is due to the number of fixed effects being large in all dimensions and that the panel is unbalanced. We use the "FEiLSDVj" estimation procedure of Andrews et al. (2006), which is based on partitioned regression techniques. We are thus forced to create and store 2000+ time-varying importer and exporter dummies (with 76,089 observations each) before algebraically stripping out the country-pair fixed effects.

4–6 reveal exclusively the time-series impact of CU accessions and exits and thus constitute the policy relevant measure we seek.

The euro is the only CU for which trade effects become both larger and more significant when we add unobserved bilateral heterogeneity controls. This supports Baldwin's (2006) hypothesis that non-euro CUs carry essentially zero informational content for euro trade effects, because these CUs' members differ dramatically from eurozone countries. Our preferred regression 6 shows that the euro increased trade by about 40 percent $\left(\approx e^{0.34}-1\right)$. This result contrasts with Balwin and Taglioni (2006), who only find negative or zero trade effects of the eurozone. The magnitude of our preferred euro estimate is comparable to those of Barr et al. (2003) and Bun and Klaassen's (2002) long-run estimates. However, our estimate is higher than those of Micco et al. (2003), Flam and Nordstrom (2003) and Bun and Klaassen (2007) who use drastically shorter panels covering fewer countries. Except for Flam and Nordstrom (2003), none of the cited studies control for pair heterogeneity and multilateral resistance.

As expected, country-pair fixed effects also provide a remedy for the negative EU effect, because they allow to correctly predict "natural" trade levels in Europe. Therefore, the EU dummy can now reflect a 25 percent $\left(\approx e^{0.22}-1\right)$ increase in trade. Furthermore, the EEA trade effect is about 57 percent $\left(\approx e^{0.449}-1\right)$. While, at 40 percent, the CU effect is smaller than the PTA trade effect for the eurozone, the combined effects of European integration (CU and PTA) caused a substantial trade increase during the 1990s. Outside of Europe, however, PTA effects are generally larger and more precisely estimated than those of CUs covering similar countries.

It is notable that FX volatility shows no significant impact on trade throughout. Currency boards are significant when aggregated (Regression 4) but insignificant when disaggregated (Regressions 5 and 6). This may be due to an insufficient number of observations in the presence of multiple fixed effects. These fragile FX volatility and currency board effects are in line with the recent empirical literature on the subject (see, e.g. Clark et al., 2004). Furthermore, theoretical literature also indicates that FX volatility may generate ambiguous trade effects in general equilibrium (Bacchetta and van Wincoop, 2000). Remaining control variables for geography, culture, and colonial history are stable, significant and of the expected magnitudes.

VI. Sensitivity Analysis

It is common in the CU literature to provide extensive sensitivity analysis to explore a range of alternative specifications. Through five perturbations to our preferred regressions, our sensitivity analysis covers virtually all remaining variables proposed by earlier literature.[12]

[12] All other previously suggested variables are already included in our analysis (absorbed into the fixed effects).

Our first perturbation follows Rose (2005) and adds regressors for membership in the three international organizations intended to promote trade: GATT/WTO, IMF and OEEC/OECD.[13] Our second perturbation adds two measures of factor endowment differences from Frankel et al. (1995) to proxy for Heckscher-Ohlin trade. These two measures are the absolute log differences in per capita GDP and population density. In the third and fourth perturbations, we drop FX volatility and the CB variables. The omission of FX volatility extends our dataset to back to 1950 and increases the number of observations by roughly ten thousand. Finally, our fifth perturbation adopts a broader CU definition (as in Glick and Rose, 2002), which defines trade flows between spokes in hub-spoke arrangements also as CU-internal.

Table 2 presents the robustness results for the aggregate CU effect with and without additional unobserved bilateral heterogeneity controls. All regressions expand on the baseline Regressions 1 and 4 but include the entire disaggregated set of individual PTAs. The implied trade increases are 42–47 percent for our preferred specification and 117–135 percent for the version without unobserved bilateral heterogeneity controls. Our preferred estimate of the average CU effect thus remains unambiguously on the order of 45 percent.

Table 3 presents robustness for the individual CU effects. To conserve space, it focuses exclusively on our preferred specification with simultaneous multilateral resistance and unobserved bilateral heterogeneity controls. That is, all results in Table 3 are direct variants of Regression 6. Like their aggregate CU counterpart in Table 2, individual CU impacts are concentrated in narrow intervals. The CFA franc is estimated between 96 and 123 percent, slightly skewed around our 97 percent benchmark. Interestingly, the CFA coefficient rises in both magnitude and significance when we control for factor endowment differences (which our results find to increase bilateral trade). The euro trade effect also remains robust at 34–40 percent. Likewise, British Pound and other/extinct CUs' effects hardly change. Our conclusion that dollarization does not improve trading relations with the United States also remains intact. The US Dollar CU impacts remain negative and even turn statistically significant in some specifications.

VII. Conclusion

Rose (2000) provided provocative estimates of the trade effects of currency unions, suggesting a tripling of trade. The subsequent literature finds smaller effects but differs from Rose's original study either in methodology or in the size of the panel. Smaller panels that cover recent CU trade effects produce significantly smaller estimates, while larger panel

---

[13] GATT = General Agreement on Tariffs and Trade, WTO = World Trade Organization, IMF = International Monetary Fund, OEEC = Organization for European Economic Co-operation, OECD = Organization for Economic Co-operation and Development. Data on GATT/WTO membership is taken from Subramanian and Wei (2007). Data on IMF and OEEC/OECD membership is taken from these institutions' websites at www.imf.org and www.oecd.org, respectively.

studies still find large trade effects. These large panel studies are, however, subject to Baldwin's 2006 critique that global trade and general equilibrium considerations ("multilateral resistance") as well as country pair specific characteristics ("unobserved bilateral heterogeneity") should be accounted for comprehensively and simultaneously to prevent omitted variable bias.

We provide an updated benchmark of the original Rose (2000) and Glick and Rose (2002) results, using an expanded dataset and simultaneous controls for multilateral resistance and unobserved bilateral heterogeneity. Three main results emerge: first, these simultaneous controls reduce the magnitudes but not the significance of CU trade effects. Yet, *individual* CUs may still generate trade effects exceeding 100 percent. The euro trade effect is, however, significantly smaller than the estimates for developing country CUs. Second, we show that a comprehensive set of PTA dummies should be included in any CU estimation, because individual PTAs exert strong and heterogeneous impacts on trade. Omission of their individual effects would thus introduce substantial omitted variable bias. Third, trade effects of PTAs seem to generally outpace those of currency unions. This, however, may result from the member country composition of particular CUs and PTAs.

Table 1: Trade Effects of Currency Unions

	Multilateral Resistance Controls only			Multilateral Resistance and Bilateral Heterogeneity		
Regression #	**1**	**2**	**3**	**4**	**5**	**6**
Adj R^2	0.734	0.738	0.739	0.866	0.867	0.867
F Statistic vs. Regr.#		# 1	# 2	# 1	# 2 # 4	# 3 # 5
Prob>F:		0.00	0.00	0.00	0.00 0.00	0.00 0.00
CU_{mxt} (Catch-all for CUs)	0.648*** (0.102)			0.424*** (0.106)		
$CUcfa_{mxt}$ (African CFA franc)		1.091*** (0.155)	1.174*** (0.155)		0.677* (0.352)	0.682* (0.352)
$CUcarib_{mxt}$ (East Caribbean $)		0.428 (0.343)	0.486 (0.342)		-0.696 (0.531)	-0.707 (0.533)
$CUeuro_{mxt}$ (Euro)		0.381*** (0.118)	0.077 (0.116)		0.537*** (0.094)	0.339*** (0.097)
$CUgbp_{mxt}$ (British Pound)		0.091 (0.198)	0.107 (0.192)		0.196 (0.166)	0.212 (0.166)
$CUusd_{mxt}$ (US Dollar)		0.332 (0.267)	0.351 (0.269)		-0.145 (0.212)	-0.148 (0.209)
$CUother_{mxt}$ (Other/Extinct CUs)		0.994*** (0.303)	1.039*** (0.302)		0.551** (0.247)	0.556** (0.247)
CB_{mxt} (Catch-all for CBs)	0.232 (0.156)			0.483* (0.295)		
$CBeuro_{mxt}$ (D-Mark/Euro CB)		0.094 (0.206)	0.122 (0.207)		0.653 (0.433)	0.651 (0.433)
$CBusd_{mxt}$ (US Dollar CB)		0.305 (0.249)	0.373 (0.251)		0.174 (0.235)	0.180 (0.234)
FXvolatility (Ex. rate volatility)	-0.008 (0.008)	-0.012 (0.008)	-0.007 (0.008)	-0.011 (0.008)	-0.011 (0.008)	-0.009 (0.008)
PTA_{mxt} (Catch-all for PTAs)	0.539*** (0.096)			0.414*** (0.055)		
$BilateralPTA_{mxt}$		0.408*** (0.073)	0.407*** (0.073)		0.049 (0.086)	0.060 (0.086)
$NAFTA_{mxt}$		0.533** (0.263)	0.256 (0.268)		0.419*** (0.148)	0.349** (0.160)
EU_{mxt}		-1.057*** (0.098)	-1.304*** (0.101)		0.477*** (0.067)	0.220*** (0.073)
$CACM_{mxt}$		2.049*** (0.195)	2.144*** (0.193)		1.963*** (0.275)	1.942*** (0.275)
$CARICOM_{mxt}$		2.607*** (0.205)	2.639*** (0.205)		0.740** (0.353)	0.723** (0.354)
$MERCOSUR_{mxt}$		1.551*** (0.262)	0.988*** (0.262)		0.438** (0.197)	0.425** (0.197)
$AFTA_{mxt}$		0.000 (0.180)	-0.181 (0.183)		-0.329 (0.229)	-0.372* (0.227)
$ANZCERTA_{mxt}$		2.353*** (0.320)	2.137*** (0.318)		0.858*** (0.150)	0.800*** (0.149)
$SPARTECA_{mxt}$		2.006*** (0.289)	2.047*** (0.289)		0.810*** (0.205)	0.804*** (0.205)
EEA_{mxt}			0.659*** (0.090)			0.449*** (0.082)
$EFTA_{mxt}$			0.059 (0.145)			0.063 (0.113)
AP_{mxt}			0.668*** (0.201)			0.921*** (0.201)
$LAIA_{mxt}$			0.769*** (0.123)			1.385*** (0.268)
$APEC_{mxt}$			0.497*** (0.072)			0.099 (0.081)
$CurColony_{mxt}$ (Current colony)	0.632*** (0.229)	0.625*** (0.221)	0.606*** (0.220)	0.100 (0.173)	0.096 (0.170)	0.103 (0.170)
$EverColony_{mx}$ (Ever colony)	1.395*** (0.089)	1.366*** (0.084)	1.399*** (0.084)			
$ComColonizer_{mx}$ (Common colonizer)	0.594*** (0.058)	0.509*** (0.059)	0.524*** (0.059)			
$ComLang_{mx}$ (Common language)	0.336*** (0.038)	0.289*** (0.038)	0.237*** (0.039)			
$ComNat_{mx}$ (Same nation)	1.956*** (0.429)	1.838*** (0.442)	1.838*** (0.442)			
$Border_{mx}$ (Common border)	0.148 (0.092)	0.206*** (0.087)	0.175** (0.087)			
$Dist_{mx}$ (Log of distance)	-1.286*** (0.021)	-1.276*** (0.021)	-1.246*** (0.022)			

Notes: *, **, *** are 10, 5, 1% significance levels. Standard errors (clustered by country-pairs) in parentheses. Coefficients of Fixed Effect controls are suppressed.

Table 2. Sensitivity Analysis: Average Currency Union Effects on Trade

	Multilateral Resistance Controls only						Multilateral Resistance and Bilateral Heterogeneity					
CU_{mxt} **(Catch-all for CUs)**	0.831*** (0.101)	0.798*** (0.101)	0.853*** (0.108)	0.829*** (0.101)	0.834*** (0.097)	0.774*** (0.097)	0.374*** (0.109)	0.371*** (0.109)	0.383*** (0.121)	0.372*** (0.109)	0.380*** (0.106)	0.347*** (0.108
Individual PTA controls	Yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
GATT/WTO, IMF, OEEC/OECD controls		yes						yes				
Factor Endowment controls			yes						yes			
Currency Board controls				no						no		
Exchange Rate volatility control					no						no	
Broad CU definition						yes						yes

Notes: *, **, *** are 10, 5, 1% significance levels. Standard errors (clustered by country-pairs) in parentheses. Coefficients of Fixed Effect and remaining controls are suppressed. The remaining controls are as in Table 1, Regression 1 (for the left half of the table) and as in Table 1, Regression 4 (for the right half of the table). The estimates in the left half of the table above are obtained by including time-varying importer and exporter fixed effects only. In the right half of the table, country-pair fixed effects are additionally included.

Table 3. Sensitivity Analysis: Trade Effects of Individual Currency Unions

	Multilateral Resistance and Bilateral Heterogeneity				
$CUcfa_{mxt}$ **(CFA franc)**	0.717** (0.345)	0.682* (0.352)	0.673* (0.352)	0.803** (0.375)	0.683* (0.352)
$CUcarib_{mxt}$ **(East Caribbean $)[a]**	0.057 (0.675)	0.017 (0.785)	0.022 (0.794)	-0.312 (0.894)	0.025 (0.783)
$CUeuro_{mxt}$ **(Euro)**	0.302*** (0.098)	0.327*** (0.097)	0.339*** (0.097)	0.292*** (0.096)	0.339*** (0.097)
$CUgbp_{mxt}$ **(BritishPound)**	0.224 (0.177)	0.212 (0.166)	0.211 (0.166)	0.255 (0.181)	0.336* (0.192)
$CUusd_{mxt}$ **(US Dollar)**	0.080 (0.250)	-0.146 (0.208)	-0.150 (0.206)	-0.340** (0.163)	-0.532** (0.269)
$CUother_{mxt}$ **(Other CUs)**	0.598*** (0.239)	0.556** (0.247)	0.552** (0.246)	0.415 (0.285)	0.564** (0.246)
Individual PTA controls	yes	yes	yes	yes	yes
Exchange Rate volatility control	no				
Currency Board controls		no			
GATT/WTO, IMF, OEEC/OECD controls			yes		
Factor Endowment controls				yes	
Broad Currency Union definition					yes

Notes: *, **, *** are 10, 5, 1% significance levels. Standard errors (clustered by country-pairs) in parentheses. Coefficients of Fixed Effect and remaining controls are suppressed. The remaining controls are as in Table 1, Regression 6.

Appendix

Table A1. Countries in Sample

Albania	Dominican Republic	Lithuania	Slovak Republic
Algeria	Ecuador	Luxembourg	Slovenia
Angola	Egypt	Macedonia, for. Yug. Rep. of	Solomon Islands
Antigua and Barbuda	El Salvador	Madagascar	Somalia
Argentina	Equatorial Guinea	Malawi	South Africa
Armenia	Estonia	Malaysia	Spain
Australia	Ethiopia	Maldives	Sri Lanka
Austria	Fiji	Mali	St. Kitts and Nevis
Azerbaijan	Finland	Malta	St. Lucia
Bahamas, The	France	Mauritania	St. Vincent & The Grenadines
Bahrain, Kingdom of	Gabon	Mauritius	Sudan
Bangladesh	Gambia, The	Myanmar	Suriname
Barbados	Georgia	Mexico	Swaziland
Belarus	Germany	Moldova	Sweden
Belgium	Ghana	Mongolia	Switzerland
Belize	Greece	Morocco	Syrian Arab Republic
Benin	Grenada	Mozambique	Tajikistan
Bermuda	Guatemala	Namibia	Tanzania
Bhutan	Guinea	Nepal	Thailand
Bolivia	Guinea-Bissau	Netherlands	Togo
Botswana	Guyana	New Zealand	Tonga
Brazil	Haiti	Nicaragua	Trinidad and Tobago
Bulgaria	Honduras	Niger	Tunisia
Burkina Faso	Hungary	Nigeria	Turkey
Burundi	Iceland	Norway	Turkmenistan
Cambodia	India	Oman	Uganda
Cameroon	Indonesia	Pakistan	Ukraine
Canada	Iran, Islamic Republic of	Panama	United Arab Emirates
Cape Verde	Iraq	Papua New Guinea	United Kingdom
Central African Rep.	Ireland	Paraguay	United States
Chad	Israel	Peru	Uruguay
Chile	Italy	Philippines	Uzbekistan
China	Jamaica	Poland	Vanuatu
China, Hong Kong SAR	Japan	Portugal	Venezuela, Rep. Bol.
Colombia	Jordan	Qatar	Vietnam
Comoros	Kazakhstan	Reunion	Yemen, Republic of
Congo, Dem. Rep. of (Zaire)	Kenya	Romania	Yugoslavia, Soc. Fed. R. of
Congo, Republic of	Kiribati	Russia	Zambia
Costa Rica	Korea	Rwanda	Zimbabwe
Côte d'Ivoire (Ivory Coast)	Kuwait	Samoa	
Croatia	Kyrgyz Republic	Sao Tome & Principe	
Cyprus	Lao People's Dem.Rep	Saudi Arabia	
Czech Republic	Latvia	Senegal	
Denmark	Lesotho	Seychelles	
Djibouti	Liberia	Sierra Leone	
Dominica	Libyan Arab Jamahiriya	Singapore	

Table A2. Membership and Observations for Currency Unions and Boards

Currency Union or Board	Number of CU Observations: Strict Definition			Broad Definition			Membership
	Total	Entry	Exit	Total	Entry	Exit	
CUother$_{mxt}$ **(Total)**	1224	146	328	1371	151	438	
CUcfa$_{mxt}$**:** (African CFA Franc)	671	53	48	671	53	48	Equatorial Guinea (since 1984), Gabon, Guinea (until 1969), Guinea-Bissau (since 1996), Madagascar (until 1982), Mali (until 1962 and since 1984), Mauritania (until 1974), Niger, Reunion (until 1976), Senegal, Togo, Benin, Burkina Faso, Cameroon, Central African Rep., Chad, Comoros (until 1994), Republic of the Congo, Cote d'Ivoire
CUcarib$_{mxt}$ (East Caribbean Dollar)	101	0	16	101	0	16	Antigua and Barbuda (since 1965), Dominica (since 1965), Grenada (since 1965), St. Vincent and the Grenadines (since 1965), St. Kitts and Nevis (since 1965), St. Lucia (since 1965), Barbados (1965-1975), Guyana (1971-1975)
CUeuro$_{mxt}$ (Euro)	110	72	0	110	72	0	Austria (since 1999), Belgium (since 1999), France (sine 1999), Germany (since 1999), Italy (since 1999), Netherlands (since 1999), Finland (since 1999), Ireland (since 1999), Portugal (since 1999), Spain (since 1999), Luxembourg (since 1999)
CUgbp$_{mxt}$ (British Pound)	122	0	122	177	0	176	United Kingdom, Ireland (until 1979), Malta (until 1971), New Zealand (until 1967), South Africa (until 1961), Bahamas (until 1966), Bermuda (until 1970), Jamaica (until 1969), Cyprus (until 1972), Iraq (until 1967), Israel (until 1954), Jordan (until 1967), Kuwait (until 1967), Gambia (until 1971), The, Ghana (until 1965), Kenya (until 1967), Libya (until 1971), Malawi (until 1971), Nigeria (until 1967), Zimbabwe (until 1967), Sierra Leone (until 1965), Somalia (until 1967), Uganda (until 1967), Zambia (until 1967)
CUusd$_{mxt}$ (U.S. Dollar)	84	18	28	148	23	60	United States, Dominican Republic (until 1985), Guatemala (until 1986), Panama, Bahamas (since 1967), Bermuda (since 1969), Liberia
CUother$_{mxt}$ **(Total) (Other & Extinct)**	**136**	**3**	**114**	**146**	**3**	**138**	
French Franc	13	2	11	17	2	15	France, Algeria (until 1969), Morocco (until 1959), Reunion (1977-1998)
Austrialian Dollar	16	0	8	18	0	8	Australia, Kiribati, Tonga (until 1991), Solomon Islands (until 1979)
East African Schilling	13	1	13	13	1	13	Kenya (1966-1978), Tanzania (1966-1978), Uganda (1966-1978), Somalia (1966-1971)
Dirham/Riyal	10	0	0	10	0	0	United Arab Emirates (since 1973), Qatar (since 1973)
Portuguese Escudo	25	0	25	47	0	45	Portugal, Angola (until 1976), Cape Verde (until 1977), Guinea-Bissau (until 1977), Mozambique (until 1977)
Malayasian Dollar	2	0	2	2	0	2	Malaysia, Singapore (1966-1973)
Indian Rupee	57	0	55	57	0	55	India, Bangladesh (until 1974), Oman (until 1970), Bhutan, Myanmar (until 1971), Sri Lanka (until 1966), Pakistan (until 1967), Mauritius (until 1967), Seychelles (until 1966)
CB$_{mxt}$ **(Total)**	**89**	**61**	**0**				
CBeuro$_{mxt}$ (Mark/Euro peg)	56	44	0				Bosnia-Herzegovina (since 1997), Bulgaria (since 1999), Estonia (since 1992), Lithuania (since 1994)
CBusd$_{mxt}$ (U.S. Dollar peg)	33	17	0				East Caribbean CU members (since 1976), Hong Kong (since 1983), Argentina (1991-2002)

Notes: Table includes only countries in our dataset. Broad CU definition includes trade between spokes in hub-spoke arrangements as intra-CU trade (see Glick and Rose, 2002). Entries (exits) recorded only for country-pairs with observations prior (posterior) to entry (exit).

Table A3. Membership in Preferential Trade Agreements

Abbreviation	Name of PTA	Start	Member countries
ANZCERTA	Australia – New Zealand Closer Economic Relations Trade Agreement	1983	Australia, New Zealand
APEC	Asia Pacific Economic Community	1989	Australia, Brunei, Canada, China (1991), Chile (1994), Taiwan Province of China (1991), Hong Kong (1991), Indonesia, Japan, South Korea, Malaysia, Mexico (1993), New Zealand, Papua New Guinea (1993), Peru (1998), Philippines, Singapore, Thailand, United States, Vietnam (1998).
AP	Andean Community / Andean Pact	1969	Bolivia, Colombia, Ecuador, Peru, Venezuela (1973), Former: Chile (1969-76)
AFTA	Association of South East Asian Nations (ASEAN) Free Trade Area	1967	Brunei (1984), Cambodia (1998), Indonesia, Laos (1997), Malaysia, Myanmar (1997), the Philippines, Singapore, Thailand, Vietnam (1995).
CACM	Central American Common Market	1960	Costa Rica (1963), El Salvador, Guatemala, Honduras, Nicaragua.
CARICOM	Caribbean Community/ Carifta	1968	Antigua and Barbuda, Bahamas (1983), Barbados, Belize (1995), Dominica (1974), Guyana (1995), Grenada (1974), Jamaica, Montserrat (1974), St. Kitts and Nevis, St. Lucia (1974), St. Vincent and the Grenadines, Suriname (1995), Trinidad and Tobago.
EEA	European Economic Area	1994	Austria, Belgium, Denmark, Finland, France, Germany, Greece, Luxembourg, Iceland, Italy, Ireland, Liechtenstein, Netherlands, Norway, Portugal, Spain, Sweden, United Kingdom.
EFTA	European Free Trade Association	1960	Iceland, Liechtenstein (1991), Norway (1986), Switzerland Former: Denmark (1960-72), United Kingdom (1960-72), Portugal (1960-85), Austria (1960-94), Sweden (1960-94), Finland (1986-94).
EU	European Union	1958	Austria (1995), Belgium, Denmark (1973), Finland (1995), France, Germany, Greece (1981), Luxembourg, Ireland (1973), Italy, Netherlands, Portugal (1986), Spain (1986), Sweden (1995), United Kingdom (1973).
LAIA/LAFTA	Latin America Integration Agreement	1960	Argentina, Bolivia (1967), Brazil, Chile, Colombia (1961) Ecuador (1961), Mexico, Paraguay, Peru, Uruguay, Venezuela (1966).
MERCOSUR	Southern Cone Common Market	1991	Argentina, Brazil, Paraguay, Uruguay
NAFTA	Canada-US Free Trade Arrangement / North America Free Trade Agreement	1988	Canada, United States, Mexico (1994).
SPARTECA	South Pacific Regional Trade and Economic Cooperation Agreement	1981	Covers trade relations between the Cook Islands, Fiji, Kiribati, Marshall Islands, Micronesia, Nauru, Niue, Palau, Papua-New Guinea, Salomon Islands, Samoa, Tonga, Tuvalu, Vanuatu, on the one hand, and Australia and New Zealand on the other
BilateralPTA	Bilateral Preferential Trade Agreements		All bilateral agreements considered are listed in Table A2.

Table A4. Bilateral Preferential Trade Agreements

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

References

Abowd, J., Kramarz, F., Margolis, D., 1999, "High wage workers and high wage firms", *Econometrica*, 67, p. 251-333.

Andrews, M., Schank, T., Upward, R., 2006, "Practical fixed effects estimation methods for the three-way error components model", *The Stata Journal*, 6(4), pp. 461-481.

Aitken, N.D., 1973, "The Effect of the EEC and EFTA on European Trade: A Temporal Cross-Section Analysis", *American Economic Review*, 63 (5), pp. 881-892.

Anderson, J., 1979, "A theoretical foundation for the gravity equation", *American Economic Review*, 69 (1), pp. 106-116.

Anderson, J., van Wincoop, E., 2003, "Gravity with Gravitas: a solution to the border puzzle", *American Economic Review*, 93 (1), pp. 170-192.

Bacchetta, P., van Wincoop, E., 2000, "Does Exchange-Rate Stability Increase Trade and Welfare?", *The American Economic Review*, 90 (5), pp. 1093-1109.

Baier, S.L., Bergstrand, J.H., 2001, "The growth of world trade: tariffs, transport costs, and income similarity", *Journal of International Economics*, 53 (1), pp. 1-27.

Baier, S.L., Bergstrand, J.H., 2007, "Do free trade agreements actually increase members' international trade?", *Journal of International Economics*, 71, pp. 72-95.

Baldwin, R., 2005. "The euro's trade effects", Working Paper prepared for ECB Workshop "What effects is EMU having on the euro area and its member countries?", Frankfurt 16 June 2005.

Baldwin, R., Taglioni, D., 2004, "Positive OCA criteria: Microfoundations for the Rose effect", Working Paper, Graduate Institute of International Studies, Geneva.

Baldwin, R., Taglioni, D., 2006, "Gravity for Dummies and Dummies for Gravity Equations", NBER Working Paper No. 12516.

Baltagi, B.H., Egger, P., Pfaffermayr, M., 2003. "A generalized design for bilateral trade flow models", *Economics Letters*, 80, pp. 391-397.

Barr, D., Breedon, F., Miles, D., 2003, "Life on the outside: economic conditions and prospects outside euroland", *Economic Policy*, 18 (37), pp. 573-613.

Bergstrand, J.H., 1985, "The gravity equation in international trade: some microeconomic foundations and empirical evidence", *Review of Economics and Statistics*, 67 (3), pp. 474-481.

Bernard, A.B., Jensen, J.B., 2004, "Entry, Expansion, and Intensity in the U.S. Export Boom", *Review of International Economics*, 12 (4), pp. 662-675.

Bun, M., Klaassen, F., 2002. "Has the Euro increased trade?", Working Paper, University of Amsterdam, Netherlands.

Bun, M., Klaassen, F., 2007, "The Euro Effect on Trade is not as Large as Commonly Thought", Oxford Bulletin of Economics and Statistics, *forthcoming*.

Carrere, C., 2006, "Revisiting the effects of regional trade agreements on trade flows with proper specification of the gravity model", *European Economic Review*, 50, pp. 223-247.

Cheng, I., Wall, H.J., 2005, "Controlling for Heterogeneity in Gravity Models of Trade and Integration", *Federal Reserve Bank of St. Louis Review*, 87 (1), pp. 49-63.

Clark, P., Tamirisa, N., Wei, S.J., Sadikov, A., Zeng, L., 2004. "Exchange Rate Volatility and Trade Flows - Some New Evidence", Working Paper, International Monetary Fund.

Deardorff, A.V., 1998, "Determinants of bilateral trade: does gravity work in a classical world", in: Frankel, J. (Ed.), *The Preferentialization of the World Economy*. University of Chicago Press, Chicago, pp. 7 –22.

De Grauwe, P., 1992, "The Benefits of a Common Currency", in De Grauwe, P. (ed.): *The Economics of Monetary Integration*, Oxford University Press, New York.

Eaton, J., Kortum, S., 2002. "Technology, geography, and trade", *Econometrica*, 70 (5), pp. 1741-1779.

Evenett, S. J., Keller, W., 2002, "On Theories Explaining the Success of the Gravity Equation", *Journal of Political Economy*, 110, pp. 281-316.

Eicher, T.S., Henn, C., 2007, "Preferential Trade Agreements Promote Trade Strongly But Unevenly: An Exhaustive Account for Omitted Variable Biases and WTO Effects", Working Paper, University of Washington.

Eicher, T.S., Henn, C., Papageorgiou, C., 2007, "Trade Creation and Diversion Revisited: Model Uncertainty, Natural Trading Partners, and Robust PTA Effects", Working Paper, University of Washington.

European Commission, 1990, "One market, one money", *European Economy*, 44.

Flam, H., Nordstrom, H., 2003, "Trade Volume Effects of the Euro: Aggregate and Sector Estimates", Institute for International Economic Studies, Stockholm University, Seminar Paper No. 746.

Frankel, J., 1997, *Preferential Trading Blocs in the World Trading System*, Institute for International Economics, Washington, D.C.

Frankel, J., Rose, A.K., 2002, An estimate of the effect of common currencies on trade and income. *Quarterly Journal of Economics* 117 (2), 437e466.

Frankel, J., Stein, E., Wei, S.J., 1995, "Trading blocs and the Americas: the natural, the unnatural and the supernatural?" *Journal of Development Economics*, 47, pp. 61– 95.

Ghosh, S., Yamarik, S., 2004, "Are preferential trade agreements trade creating? An application of extreme bounds analysis", *Journal of International Economics*, 63, pp. 369-395.

Glick, R., Rose, A.K., 2002, "Does a currency union affect trade? The time-series evidence", *European Economic Review*, 46, pp. 1125-1151.

Hooper, P., Kohlhagen, S. 1978, "The Effect of Exchange Rate Uncertainty on the Prices and Volume of International Trade," *Journal of International Economics*, 8, pp. 483-511.

Hummels, D., Levinsohn, J., 1995, "Monopolistic Competition and International Trade: Reconsidering the Evidence", *Quarterly Journal of Economics*, 110 (3), pp. 799-836.

Karam, P., Laxton, D., Rose, D., Tamarisa, N., "The macroeconomic costs and benefits of adopting the euro", IMF Staff Papers, 55 (2), pp. 339-355.

Klein, M.W., 2005, "Dollarization and Trade", *Journal of International Money and Finance*, 24, pp. 935-943.

Klein, M.W., Shambaugh, J.C., 2006, "Fixed exchange rates and trade", *Journal of International Economics*, 70, pp. 359-383.

Linnemann, H., 1966, *An Econometric Study of International Trade Flows*, North-Holland Publishing Company, Amsterdam.

Mancini-Griffoli, T., 2006, "Explaining the Euro's Effect on Trade? Interest Rates in an Augmented Gravity Equation", HEI Working Paper No. 10/2006, Graduate Institute of International Studies, Geneva.

Mancini-Griffoli, T., Pauwels, L., 2006, "Is There a Euro Effect on Trade? An Application of End-of-Sample Structural Break Tests for Panel Data", HEI Working Paper No. 04/2006, Graduate Institude of International Studies, Geneva.
McCallum, J., 1995, "National Borders Matter: Canada-U.S. Regional Trade Patterns", *American Economic Review*, 85 (3), pp. 615-623.
Melitz, M., 2003, "The impact of trade on intra-industry reallocations and aggregate industry productivity", *Econometrica*, 71 (6), pp. 1695-1725.
Micco, A., Stein, E., Ordonez, G., 2003, "The currency union effect on trade: early evidence on EMU", *Economic Policy*, 18 (37), pp. 315-336.
Nitsch, V., 2002, "Honey, I Shrunk the Currency Union Effect on Trade", The World Economy, 25 (4), pp. 457-474.
Nitsch, V., 2004, "Comparing Apples and Oranges: The Effect of Multilateral Currency Unions on Trade", in: Alexander,V., Melitz, J., Furstenberg, G.M., eds., *Monetary Unions and Hard Pegs: Effects on Trade, Financial Development and Stability*, Oxford University Press.
Oguledo, V.I., MacPhee, C.R., 1994, "Gravity models: A reformulation and an application to discriminatory trade arrangements", *Applied Economics*, 26 (2), pp. 107-120.
Pakko, M.R., Wall, H.J., 2001. "Reconsidering the Trade-Creating Effects of a Currency Union", *Federal Reserve Board of St. Louis Review*, 83 (5), pp. 37-45.
Persson, T., 2001, "Currency unions and trade: how large is the treatment effect?", *Economic Policy*, 33, pp. 435-448.
Pollak, J.J., 1996, "Is APEC a Natural Regional Trading Bloc? A Critique of the 'Gravity Model' of International Trade", *World Economy*, pp. 533-543.
Rose, A.K., 2000, "One money, one market? The Effects of Common Currencies on International Trade", *Economic Policy*, 15, pp. 7-46.
Rose, A.K., van Wincoop, E., 2001, "National Money as a Barrier to International Trade: The Real Case for a Currency Union", *American Economic Review*, 91 (2), pp. 386-390.
Rose, A.K., 2004, "Do We Really Know That the WTO Increases Trade?", *American Economic Review*, 13 (4), pp. 682-698.
Rose, A.K., 2005, "Which International Institutions Promote Trade?", *Review of International Economics*, 13 (4), pp. 682-698.
Rose, A.K., 2005, Stanley, T., 2005, "A Meta-Analysis of the Effect of Common Currencies on International Trade", *Journal of Economic Surveys*, 19, pp. 347-365.
Soloaga, I., Winters, L.A., 2001, "Preferentialism in the nineties: what effect on trade?" *North American Journal of Economics and Finance*, 12, pp. 1-29.
Subramanian, A., Wei, S., 2007, "The WTO Promotes Trade, Strongly But Unevenly", *Journal of International Economics*, forthcoming.
Yi, K.M., 2003, "Can vertical specialization explain the growth of world trade?", *Journal of Political Economy*, 111, pp. 52-102.