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One Money, One Market: A Revised Benchmark

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Abstract

The introduction of the euro generated substantial interest in the impact of currency unions (CUs) on trade flows. Initial estimates suggested a tripling of trade, which gave rise to a literature in search of "more reasonable" CU effects. Theoretical derivations of the gravity model highlight, however that the CU literature neglects to control simultaneously for general equilibrium effects (multilateral resistance) and unobserved bilateral heterogeneity among trade partners. Once we introduce the appropriate controls, CU trade effects are shown to range around 50%. We also highlight that the practice of reporting average CU effects generates misleading results. The average effect is shown to be a composite of disparate individual CU effects ranging from 40% (euro) to about 100% (Central African franc).

1. Introduction

The advent of the euro generated keen interest in quantifying currency unions' (CUs') impact on trade. Rose (2000) and Frankel and Rose (2002) estimated that CUs triple bilateral trade with other currency union members. Such increases seemed questionably high and gave rise to an entire literature comprised of alternative approaches designed to "shrink the Rose effect." Rose and Stanley (2005) surveyed the literature and reported that subsequent papers with similar samples still report CU trade effects that exceed 100%. Baldwin's (2006) theoretical derivation of the gravity model highlights that the CU literature neglects to control simultaneously for general equilibrium effects (multilateral resistance) and unobserved bilateral heterogeneity among trade partners. Large panel studies of CU effects in the tradition of Rose (2000) that include these necessary controls simultaneously do not exist.

Previous approaches to controlling for multilateral resistance included geography-based remoteness (Rose, 2000) and country fixed effects (Rose and van Wincoop, 2001), which neglected changes in general equilibrium effects over time.[1] Below we outline theoretically and empirically how coefficients are affected by omitted variable bias when comprehensive multilateral resistance controls are absent from the analysis. As importantly, Baldwin (2006) highlighted the omitted variable bias introduced when the empirical strategy does not account for unobserved determinants of bilateral trading relationships.[2] Failure to include the adequate fixed effect controls leads to such severe bias that Baldwin recommends ignoring any other estimates for policy purposes.

Baldwin and Taglioni (2006) implemented the two crucial sets of simultaneous controls advocated by Baldwin (2006) in a small panel of EU countries and found either negative or zero trade effects of the euro.[3] Frankel (2010) revisited Rose

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(2000) in a large panel of countries, but controlled only for unobserved bilateral heterogeneity.[4]

Here we provide a revised benchmark for CU trade effects by simultaneously addressing all methodological issues raised by Baldwin (2006) and by updating and extending Rose's (2000) panel dataset to cover 50 years and 117 countries. We also address an issue that has not been appropriately emphasized in the CU literature: previous studies focus on one average CU effect, but trade effects may differ dramatically depending on the specific CU. With the exception of Nitsch (2002) and Frankel (2010), this heterogeneity has not been investigated. Not only does the omission of individual CU effects introduce potential omitted variable bias, but for policy purposes, average CU trade effects are uninformative. For instance, the trade effects of the euro should not be indicated by a coefficient that is a mix of euro, Central African franc (CFA), and East Caribbean Currency Union effects.

We posit that all three aspects—multilateral resistance, country-pair fixed effects, and individual CU effects—must be introduced simultaneously in order to eliminate bias to CU trade effects. First we show that Rose's (2000) average CU trade effect remains statistically and economically significant once we control for the proper fixed effects specification. The effect is, however, reduced to a more realistic 52%.[5] Then we highlight how dramatically the trade effects differ across the individual CUs, with a strong, positive effect equivalent to a 40% trade increase for the Eurozone countries, and over 100% trade increases for CFA countries. Other CUs show no effect, however, including the East Caribbean Currency Union and hub-and-spoke CUs, in which the spoke country adopts a major world currency. Finally we note that preferential trade agreements (PTAs) produce generally greater trade effects than CUs—with the exception of Europe, where euro trade effects challenge those produced by the common market.

2. Data

Our dataset is based on Subramanian and Wei (2007), who in turn derived their data from Rose (2004). Our 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) and we have 76,081 observations (1224 in CUs), 177 countries, and 16,941 bilateral trade relationships. The increased size of the sample is important because it covers more countries in CUs: 330 observations in Rose (2000) vs 1224 in our dataset.

Our dependent variable is bilateral imports at five-year intervals, deflated by the US consumer price index. 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 the previous literature. First, we augment the dataset to include a large list of major PTAs obtained from Ghosh and Yamarik (2004).[6] 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 ($CBeuro_{mxt}$). 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 gross domestic product (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 Interntional Monetary Fund (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 three years). Regressions including foreign exchange (FX) volatility reduce the dimension of the dataset to 66,619 observations with 15,833 pairs starting in 1960. The appendix tables in the working paper version of this study (Eicher and Henn, 2009) summarize membership in CUs, CBs, and PTAs, other explanatory variables, and country coverage of the dataset.

3. Empirical Implementation of the Gravity Model

Baldwin (2006) leveled two fundamental critiques against recent 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) started 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:

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

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}, however, 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).

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

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

Aside from transport costs (proxied by distance), currency arrangements, and PTAs, 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.[7] 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 time-invariant unobservables in bilateral trade relationships. The absence of country-pair fixed effects is not usually due to oversight on the part of the researcher. Especially in the CU literature, the paucity of observations that represent countries entering/exiting CUs may render the introduction of these effects too restrictive in small datasets. Our dataset instead proves sufficiently large to provide significant results.

The third methodological aspect that we address 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 drastically.[8] Hence, we allow not only for individual PTAs but also examine results for individual CUs in our analysis below.

4. Multilateral Resistance and the Trade Effects of Currency Unions

Our empirical strategy proceeds in stages. We first introduce controls for multilateral resistance; then we 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% to 136% in the process. The Rose and van Wincoop (2001) strategy sufficiently addressed multilateral resistance in a cross-section; however, the paper did not capture the time-varying nature of trade costs in panel data. Baldwin and Taglioni (2006) addressed 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 (1) and (2) according to

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

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.

Regressions (1)–(3) in Table 1 present our baseline results for CU trade effects with multilateral resistance controls. Regression (1) can be directly compared with Rose's (2000) benchmark regression except for the addition of multilateral resistance controls.[9] At 0.65, the CU coefficient estimate is roughly six standard deviations lower than Rose's original 1.21. This reduces the CU trade increase to 91% ($\approx e^{0.648} - 1$) as opposed to Rose's tripling estimate (the 235% increase). The estimate is also significantly smaller than Rose and van Wincoop's (2001), who did not consider the time-varying

Table 1. Trade Effects of Currency Unions

Fixed effects	*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 regression		# 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}	0.648***			0.424***		
(Catch-all for CUs)	(0.102)			(0.106)		
$CUcfa_{mxt}$		1.091***	1.174***		0.677*	0.682*
(African CFA franc)		(0.155)	(0.155)		(0.352)	(0.352)
$CUcarib_{mxt}$		0.428	0.486		–0.696	–0.707
(East Caribbean $)		(0.343)	(0.342)		(0.531)	(0.533)
$CUeuro_{mxt}$		0.381***	0.077		0.537***	0.339***
(Euro)		(0.118)	(0.116)		(0.094)	(0.097)
$CUgbp_{mxt}$		0.091	0.107		0.196	0.212
(British Pound)		(0.198)	(0.192)		(0.166)	(0.166)
$CUusd_{mxt}$		0.332	0.351		–0.145	–0.148
(US dollar)		(0.267)	(0.269)		(0.212)	(0.209)
$CUother_{mxt}$		0.994***	1.039***		0.551**	0.556**
(Other/Extinct CUs)		(0.303)	(0.302)		(0.247)	(0.247)
CB_{mxt}	0.232			0.483*		
(Catch-all for CBs)	(0.156)			(0.295)		
$CBusd_{mxt}$		0.305	0.373		0.174	0.180
(US dollar CB)		(0.249)	(0.251)		(0.235)	(0.234)

Table 1. Continued

Fixed effects	*Multilateral resistance controls only*			*Multilateral resistance and bilateral heterogeneity*		
Regression	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
$CBeuro_{mxt}$		0.094	0.122		0.653	0.651
(D-mark/Euro CB)		(0.206)	(0.207)		(0.433)	(0.433)
PTA_{mxt}	0.539***			0.414***		
(Catch-all for PTAs)	(0.096)			(0.055)		
$BilateralPTA_{mxt}$		0.408***	0.407***		0.049	0.060
		(0.073)	(0.073)		(0.086)	(0.086)
$NAFTA_{mxt}$		0.533**	0.256		0.419***	0.349**
		(0.263)	(0.268)		(0.148)	(0.160)
EU_{mxt}		–1.057***	–1.304***		0.477***	0.220***
		(0.098)	(0.101)		(0.067)	(0.073)
$CACM_{mxt}$		2.049***	2.144***		1.963***	1.942***
		(0.195)	(0.193)		(0.275)	(0.275)
$CARICOM_{mxt}$		2.607***	2.639***		0.740**	0.723**
		(0.205)	(0.205)		(0.353)	(0.354)
$MERCOSUR_{mxt}$		1.551***	0.988***		0.438**	0.425**
		(0.262)	(0.262)		(0.197)	(0.197)
$AFTA_{mxt}$		0.000	–0.181		–0.329	–0.372*
		(0.180)	(0.183)		(0.229)	(0.227)
$ANZCERTA_{mxt}$		2.353***	2.137***		0.858***	0.800***
		(0.320)	(0.318)		(0.150)	(0.149)
$SPARTECA_{mxt}$		2.006***	2.047***		0.810***	0.804***
		(0.289)	(0.289)		(0.205)	(0.205)

EEA_{mxt}			0.659***			0.449***
			(0.090)			(0.082)
$EFTA_{mxt}$			0.059			0.063
			(0.145)			(0.113)
AP_{mxt}			0.668***			0.921***
			(0.201)			(0.201)
$LAIA_{mxt}$			0.769***			1.385***
			(0.123)			(0.268)
$APEC_{mxt}$			0.497***			0.099
			(0.072)			(0.081)
Fxvolatility	–0.008	–0.012	–0.007	–0.011	–0.011	–0.009
(Ex. rate volatility)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
$CurColony_{mxt}$	0.632***	0.625***	0.606***	0.100	0.096	0.103
(Current colony)	(0.229)	(0.221)	(0.220)	(0.173)	(0.170)	(0.170)
$EverColony_{mx}$	1.395***	1.366***	1.399***			
(Ever colony)	(0.089)	(0.084)	(0.084)			
$ComColonizer_{mx}$	0.594***	0.509***	0.524***			
(Common colonizer)	(0.058)	(0.059)	(0.059)			
$ComLang_{mx}$	0.336***	0.289***	0.237***			
(Common language)	(0.038)	(0.038)	(0.039)			
$ComNat_{mx}$	1.956***	1.838***	1.838***			
(Same nation)	(0.429)	(0.442)	(0.442)			
$Border_{mx}$	0.148	0.206***	0.175**			
(Common border)	(0.092)	(0.087)	(0.087)			
$Dist_{mx}$	–1.286***	–1.276***	–1.246***			
(Log of distance)	(0.021)	(0.021)	(0.022)			

Notes: *,**,*** denote 10%, 5%, and 1% significance levels, respectively. Standard errors (clustered by country-pairs) in parentheses. PTAs are detailed in note 6.

nature of multilateral resistance. Their estimate of 0.86 (implying a 136% 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, 2010). 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 trade within the African CFA zone is estimated to be 197–224% higher than trade with outsiders. The hub-spoke arrangements of $CUother_{mxt}$ show a similar trade increase of 157–183%. CUs involving the British pound, US dollar, and East Caribbean dollar show no statistically significant effects.

The trade effect of the euro is the surprise in this set of results. In regression (2), our estimated euro trade increase ($46\% \approx e^{0.381} - 1$) 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. Regression (3) suggests that subsequent trade flows were mainly affected by PTA-based integration and hardly by the formation of the Eurozone. These results underscore 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 over-predict trade systematically for 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.

5. Benchmark Trade Effects of Currency Unions: Accounting for 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 4000 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 (3), we obtain our new estimation equation:

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

All time-invariant pair-specific variables are now absorbed into the country-pair fixed effects α_{mx}.[10]

In large trade datasets, the estimation of three-way fixed effect structures as in (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 provided strong economic and statistical arguments in favor of our proposed three-way error components model. They, however, motivate the time-varying importer and exporter dummies with business cycles and country-specific political and institutional conditions rather than omitted price terms. Eicher and Henn (2011) exploited 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 1 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% ($\approx e^{0.42} - 1$), the average CU effect has about halved and differs by more than two standard deviations from our previous estimate of 91% (regression (1)).[12] The 53% estimate is statistically significant but dramatically lower than the 123% ($\approx e^{0.80} - 1$) 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 US dollar remain insignificant. The African CFA and Other (extinct) hub-spoke CUs, however, stay significant but show reduced trade impact. In percentage terms, their effects halve to 97% and 73%, respectively. The CU coefficients in regressions (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% ($\approx e^{0.34} - 1$). 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. Interestingly, Persson (2001) produced results almost identical to ours (a 44% trade increase) using a matching estimator to control unobserved country-pair heterogeneity. However, his estimate is not statistically significant. Using a similar matching technique, Chintrakarn (2008) produced a 14% effect of the euro. Our estimate is also higher than those of Micco et al. (2003), Flam and Nordstrom (2003) and Bun and Klaassen (2007), who used dramatically shorter panels covering fewer countries. Except for Flam and Nordstrom (2003), none of the cited studies simultaneously controlled for pair heterogeneity and multilateral resistance. We will explain below that much of the CU effect accrues post-accession, hence the shorter time periods may produce smaller effects.

As expected, country-pair fixed effects also provide a remedy for the negative EU effect since they correct for "natural" trade levels in Europe. Therefore, the EU dummy can now reflect a 25% ($\approx e^{0.22} - 1$) increase in trade and the EEA trade effect is about 57% ($\approx e^{0.449} - 1$). The impact of the euro, at 40%, eclipses the EU effect. However, EEA, which extended the common market to some non-EU member countries from 1994 onwards, is also estimated to have had a 57% trade-enhancing effect. Thus, Europe managed to amalgamate CU and PTA-based integration during the 1990s to reap substantial trade benefits. Outside of Europe, however, PTA effects are generally larger and more precisely estimated than effects 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). To some degree the result is to be expected, since the theoretical literature indicates that FX volatility generates 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.

Accession Dynamics

The results above refer to average trade effects over the entire course of CU membership. However, CU effects may not be constant over time. The mere announcement of a future accession might already anticipate trade increases and benefits of CU membership might continue to accrue over time. The evolution of the CU effect over time can be further examined by separating the average CU dummy into time periods that identify the pre-accession, accession, and post accession periods. Since a number of countries exit CUs in our sample, we can also identify secession effects.

Table 2. Accession/Secession Dynamics of Currency Unions

Fixed effects	*Multilateral resistance controls only*	*Multilateral resistance and bilateral heterogeneity*
Regression	*1A*	*4A*
Adj R^2	0.743	0.867
F-statistic vs regression	1	4
Prob > F	0.96	0.00
CU_{mxt}	0.648***	0.424***
Average[a]	(0.102)	(0.106)
CU_{mxt}	−0.902***	0.107
Pre-accession (t–1)	(0.179)	(0.098)
CU_{mxt}	0.081	0.305***
Accession (t)	(0.102)	(0.114)
CU_{mxt}	0.981***	0.525***
Post-accession ($t+1, n$)	(0.145)	(0.140)
CU_{mxt}	0.548***	0.092
Post-secession ($n+1$)	(0.175)	(0.130)
CU_{mxt}	−0.007	−0.147
Post-secession ($n+2$)	(0.157)	(0.139)
CB_{mxt}	0.195	0.652***
(Catch-all for CBs)	(0.157)	(0.249)
PTA_{mxt}	0.609***	0.440***
(Catch-all for PTAs)	(0.097)	(0.053)
Fxvolatility	−0.007	−0.007
(Ex. rate volatility)	(0.008)	(0.007)
$CurColony_{mxt}$	0.690***	0.046
(Current colony)	(0.228)	(0.171)
$EverColony_{mx}$	1.375***	
(Ever colony)	(0.088)	
$ComColonizer_{mx}$	0.585***	
(Common colonizer)	(0.058)	
$ComLang_{mx}$	0.334***	
(Common language)	(0.038)	
$ComNat_{mx}$	2.093***	
(Same nation)	(0.433)	
$Border_{mx}$	0.150	
(Common border)	(0.091)	
$Dist_{mx}$	−1.285***	
(Log of distance)	(0.021)	

Notes: *,**,*** denote 10%, 5%, and 1% significance levels, respectively. Standard errors (clustered by country-pairs) in parentheses.
[a] The average effect (estimated in Table 1) is inserted into this table for comparison purposes.

Table 2 provides the results for our benchmark specifications. While the other covariates remain largely unchanged, the CU regressor shows clear evidence of accession dynamics. We focus on our preferred specification, regression (4A), since analog regression (1A), which omits unobserved heterogeneity controls, shows again some signs of misspecification.

The pre-accession effect, measured during the five years before joining the CU, is negligible and even negative, if unobserved bilateral heterogeneity is ignored.

However, immediately in the accession period, the trade effect of CUs turns positive. The largest trade effects are reserved, however, for the post-accession phase, indicating that the benefits of CUs accrue over time. The post-accession CU effect is estimated to be nearly 70% ($\approx e^{0.525} - 1$), and shows clearly that the average CU effect is a composite of the weaker CU accession effect and the stronger CU post-accession effect. When a country leaves a CU, its trade swiftly reverts to its expected level and no significant trade creation remains observable. Interestingly, currency boards become more significant, when we control for accession dynamics.

6. 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.[13] 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.[14] 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 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 3 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% for our preferred specification and 117–135% for the version without unobserved bilateral heterogeneity controls. Our preferred estimate of the average CU effect thus remains unambiguously on the order of 50%.

Table 4 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 4 are direct variants of regression (6). Like the aggregate CU impacts in Table 3, individual CU impacts are concentrated in narrow intervals. The CFA franc is estimated between 96% and 123%, slightly skewed around our 97% benchmark. Interestingly, the CFA coefficient rises in both magnitude and significance when we control for factor endowment differences (which, according to our results, increase bilateral trade). The euro trade effect also remains robust at 34–40%. Likewise, British pound and other/extinct CUs' effects hardly change. Our conclusion that dollarization does not improve trading relations with the USA also remains intact. The US dollar CU impacts remain negative and even turn statistically significant in some specifications.

7. Conclusion

The introduction of the euro raised interest in quantifying the trade effects associated with CUs. Early estimates suggested a tripling of trade because of CUs. While subsequent studies find smaller effects, most large panel studies still imply trade effects of

Table 3. Sensitivity Analysis: Average Currency Union Trade Effects

Fixed effects	*Multilateral resistance controls only*							*Multilateral resistance and bilateral heterogeneity*						
CU_{mxt}	1.078***	0.831***	0.798***	0.853***	0.829***	0.834***	0.774***	0.382*	0.374***	0.371***	0.383***	0.372***	0.380***	0.347***
(Catch-all for CUs)	(0.141)	(0.101)	(0.101)	(0.108)	(0.101)	(0.097)	(0.097)	(0.223)	(0.109)	(0.109)	(0.121)	(0.109)	(0.106)	(0.108)
Rose original dataset	yes							yes						
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: *,**,*** denote 10%, 5%, 1% significance levels, repectively. 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 4. Sensitivity Analysis: Trade Effects of Individual Currency Unions

	Multilateral resistance and bilateral heterogeneity				
$CUcfa_{mxt}$	0.717**	0.682*	0.673*	0.803**	0.683*
(CFA franc)	(0.345)	(0.352)	(0.352)	(0.375)	(0.352)
$CUcarib_{mxt}$	0.057	0.017	0.022	–0.312	0.025
(*East Caribbean dollar*)[a]	(0.675)	(0.785)	(0.794)	(0.894)	(0.783)
$CUeuro_{mxt}$	0.302***	0.327***	0.339***	0.292***	0.339***
(Euro)	(0.098)	(0.097)	(0.097)	(0.096)	(0.097)
$CUgbp_{mxt}$	0.224	0.212	0.211	0.255	0.336*
(British pound)	(0.177)	(0.166)	(0.166)	(0.181)	(0.192)
$CUusd_{mxt}$	0.080	–0.146	–0.150	–0.340**	–0.532**
(US dollar)	(0.250)	(0.208)	(0.206)	(0.163)	(0.269)
$CUother_{mxt}$	0.598***	0.556**	0.552**	0.415	0.564**
(Other CUs)	(0.239)	(0.247)	(0.246)	(0.285)	(0.246)
Currency board controls	yes		yes	yes	yes
Individual PTA controls	yes	yes	yes	yes	yes
Exchange rate volatility control		yes	yes	yes	yes
GATT/WTO, IMF, OEEC/OECD controls			yes		
Factor endowment controls				yes	
Broad currency union definition					yes

Notes: *,**,*** denote 10%, 5%, 1% significance levels, respectively. 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.

over 100%. Baldwin (2006) surveyed this literature and derived appropriate gravity model controls to assess the trade effect of CUs. He noted that previous studies neglect to control simultaneously for unobserved heterogeneity in bilateral trade relationships as well as for multilateral-resistance general-equilibrium effects.

We implement Baldwin's econometric specification to provide an updated benchmark of the original Rose (2000) and Glick and Rose (2002) results. Three important results emerge: first, the proper use of controls reduces omitted variable bias and lowers the magnitudes but not the significance of average CU trade effects. Second, individual CUs generate distinctly different trade effects, so coefficients that average over all CUs might produce misleading results. Finally, we find that trade effects of PTAs generally exceed those of CUs, although latter remain important drivers of trade and many cases. For the Eurozone, our results attribute a key share of recent trade increases to the introduction of the euro.

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Notes

1. The latter approach acknowledges that variations in the relative prices of trading partners matter for trade, but assumes that a country's price vector and trade costs with the rest of the world remain constant over time.
2. Hummels and Levinsohn (1995) first emphasized such unobserved bilateral heterogeneity by including country-pair fixed effects in the estimation. Recent CU papers that also include these fixed effects are Glick and Rose (2002), Pakko and Wall (2001), and Klein and Shambaugh (2006).
3. Their dataset holds only 4837 observations to serve as an example, compared to Rose's (2000), which featured 22,948 observations, or ours (76,081 observations). Baldwin and Taglioni speculate 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.
4. The dimension of the dataset matters for three important reasons. First, previous studies report an average CU trade effect, but coverage of different years implies the inclusion of different CUs. Second, datasets that cover only subsets of countries (e.g. OECD) exclude the effects of important CUs such as the Central African Franc (CFA) in the average effect. Third, the longer the time series, the more precise are the estimates of other trade flow determinants before and after CU accession, which allows for a more precise estimate of the CU effect.
5. Glick and Rose reduce the effect to 91% ($\approx e^{0.65} - 1$) by introducing country-pair fixed effects alone.
6. PTAs included are bilateral PTAs ($BilateralPTA_{mxt}$), North America FTA ($NAFTA_{mxt}$), European Union (EU_{mxt}), Central American Common Market ($CACM_{mxt}$), Caribbean Community ($CARICOM_{mxt}$), Southern Cone Common Market ($MERCOSUR_{mxt}$), Association of South East Asian Nations FTA ($AFTA_{mxt}$), Australia–New Zealand Closer Economic Relations Trade Agreement ($ANZCERTA_{mxt}$), South Pacific Regional Trade and Economic Cooperation Agree-

ment ($SPARTECA_{mxt}$), European Economic Area (EEA_{mxt}), European Free Trade Association ($EFTA_{mxt}$), Andean Community/Pact (AP_{mxt}), Latin America Integration Agreement ($LAIA_{mxt}$), and Asia Pacific Economic Community ($APEC_{mxt}$).

7. For example, political relationships, networks of business leaders, transport infrastructure, cultural affinities, and institutional similarities.

8. Soloaga and Winters (2001) examined nine PTAs (1982–1996) and reported effects ranging from an 8% increase (EU) to 17-fold increases (CACM). Following the convention of the literature, we refer to a "percent increase in trade," while the exact terminology would be the "average percent increase in trade for the years of CU membership" for member countries. Eicher et al. (2011) examined 12 PTAs and found effects ranging from 60% (EEA) to five-fold increases (CACM). Cipollina and Salvatici (2010) conducted a meta analysis for 75 studies that cover nineteen PTAs to find estimates to range from an astonishing 2000% increase (Baltic PTA) to a 29% decrease (Canada–USA FTA). Frankel (2010) surveyed the literature on CU trade impacts and reported EMU effects ranging from 14% to 40% (in Chintrakarn, 2008; and Bun and Klaassen, 2002, respectively), while CFA is reported to have increased trade by 8%.

9. 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 1970–1990 in Rose), and (3) we additionally insert a currency board dummy, which has, however, no impact on the results.

10. Hausman and Taylor (1981) and Hsiao (1986, p. 50) showed how coefficients of time-invariant regressors that have been absorbed into the fixed effects can be recaptured. The process involves a two-stage procedure where the first stage is given by (4) and the second stage uses the estimated country-pair fixed effect coefficients as dependent variables and time-invariant regressors as explanatory variables.

11. This is due to the number of fixed effects being large in all dimensions and to the panel being unbalanced. We use the "FEiLSDVj" estimation procedure of Andrews et al. (2006), which is based on partitioned regression techniques.

12. The magnitude of the CU coefficient produced by the specification in regression (4) is not significantly different from the coefficients produced when we use Rose's original sample period (see robustness analysis in Table 3).

13. All other variables that were included in our analysis above are now absorbed into the fixed effects.

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