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Disaggregated trade and disaggregated currency unions: a ranking of common currency effects

Gregory William WHITTEN, *Lingnan University, Hong Kong*

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Abstract

Andrew Rose has long argued that a common currency has a large effect on increasing trade. Recently, Rose has called into question the reliability of this conclusion, as new techniques have emerged for estimating gravity equations. This paper uses the sector-specific gravity model developed by Anderson and Yotov (2010a) to investigate if disaggregated trade can provide a reliable estimate of a common currency's effect. Disaggregating trade alone is insufficient to obtain a reliable estimate of a currency union, regardless of econometric technique, when the effect of a common currency on trade is uniform across all unions. Disaggregating the universe of currency unions with individual effects provides a reliable ranking of currency unions, independent of estimation method, by the effect that each union's currency has on increasing trade. These rankings differ across sectors.

Keywords: international trade, gravity equation, agriculture, manufacturing, currency union

JEL: F14, F15, O19, Q17

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[*] gregorywhitten@ln.edu.hk, Lingnan University Department of Economics, 8 Castle Peak Road, Tuen Mun, Hong Kong SAR. I am deeply grateful to Steven Husted and James Cassing. I also thank the participants of the Pittsburgh International Trade & Development Seminar series of 2013 as well as the 2014 GTAP Annual Conference on Global Economic Analysis in Dakar for their input. All remaining errors are, regretfully, mine.

1 Introduction

The effect of a common currency on trade has long been a question of interest for economists. The presence of a common currency over one or more areas of economic activity suggests a high level of economic integration within those areas (See Mundell (1961)). Andrew Rose and his coauthors began in the late 1990s to provide quantitative measures of the extent of this integration (See Frankel and Rose (1998), Rose (2000), Rose and van Wincoop (2001), Frankel and Rose (2002), Rose and Engel (2002), and Glick and Rose (2002)). These articles generally estimate a gravity equation of bilateral trade in order to quantify the extent of integration, the measurement coming from the coefficient for a variable indicating that the two countries in the bilateral transaction share a common currency. These results from these regressions have been surprisingly large, ranging from 90% to over 200%. However, owing to recent developments regarding the estimation of gravity equations, Rose has called into question the reliability of his results, stating "it is currently beyond our ability to estimate the effect of currency unions on *aggregate* trade with much confidence," (Glick and Rose (2015) p. 19, emphasis mine).

While analyzing aggregate trade may not yield reliable estimates, Rose's work leaves open the possibility that analyzing disaggregate trade may be more promising. This supposition is not a mere and literal exploitation of Rose's comments. Sharing a common currency alters the barriers to trade. Anderson and Yotov (2010a) and Anderson and Yotov (2010b) emphasize that barriers to trade, and their removal, likely have differential effects across the range of goods traded. Furthermore, Rose's recent reflection comes at a time when the gravity equation has emerged as a reliable, theoretically-sound component of international economics. Recognition of gravity's importance has come as researchers have directed greater attention to theoretical models generating a gravity equation as well as to the econometric methods used to conduct regression analysis.[1]

This paper builds on these recent developments to revisit Rose's work in order to see if *disaggregate* trade may provide a reliable estimate of a common currency's effect on trade, independent of estimation technique, when *aggregate* trade has not. To do so, this paper first uses the sector-specific gravity equation of Anderson and Yotov (2010a) in order to confirm that determinants of agricultural trade differ noticeably from the determinants of manufacturing trade, confirming the supposition of Anderson and Yotov (2010a) that trade costs differ noticeably with the types of goods traded.

Given different results for the determinants of trade between agricultural and manufacturing trade, the paper makes its second contribution. Disaggregating trade does not lead to an estimated coefficient for a single, currency union indicator variable that is similar across the estimation techniques used (OLS and Pseudo-Poisson Maximum Likelihood or PPML). This result matches the result of Glick and Rose (2015). However, if one considers that each currency union has a different effect on trade (just as each type of good traded responds to a common currency and other determinants of trade differently), then it is possible to estimate a reliable *ranking* of currency union effects across estimation techniques. I.e., individual currency unions with large, estimated OLS coefficients will likely have large, estimated PPML coefficients, even though the level effects implied by the two estimation techniques for each currency union may differ. The effects are substantially different for a given type of good traded. Hence, disaggregating trade and disaggregating the definition of currency union provides a precise and richer understanding of a common currency's effect on trade. Further disaggregation within manufacturing provides a more nuanced examination of the currency union effect.

This paper draws together several topics, each of which has developed an extensive literature: the theoretical derivation of gravity equations, the estimation of

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[1] See Head and Mayer (2014)

gravity equations, the study of currency unions, and the analysis of disaggregated trade flows. However, few papers have analyzed these topics jointly. Anderson and Yotov (2010b) and Anderson and Yotov (2012) use the PPML method to estimate gravity equations on panels of data for individual manufacturing industries in order to understand how trade barriers affect producers and consumers differently. The authors do not explicitly compare and contrast the determinants to trade between agricultural trade and manufacturing trade overall. Huchet-Bourdon and Cheptea (2011) estimate the effect of a common currency on disaggregated trade, but limit themselves to the EMU and agricultural trade. This paper continues in their spirit by providing estimates for multiple currency unions and multiple, disaggregated categories of trade. Herwartz and Weber (2013) also examine the eurozone to demonstrate cross-sectional and dynamic differences in trade costs across different goods belonging to manufacturing trade.

2 Literature Review

This article draws on two related but distinct strands of literature: first, microeconomic or microeconometric theory-based literature focused on the derivation and estimation of the gravity equation; second, an empirical-based literature that uses gravity equations to characterize international trade. Within this second strand exist two largely, but not exclusively distinct, further strands: a literature that uses gravity equations on aggregate trade to characterize currency unions and a literature that adapts gravity equations to describe disaggregated trade. This section will demonstrate how this paper arises from and responds to work within each of these strands.

Though use of the gravity equation or gravity-like equations dates back at least as early as Tinbergen (1962), Anderson and van Wincoop (2003) point out that estimates of the gravity equation made up to the late 1990s generally omitted important control variables required by the theoretical foundations of the gravity equation. In particular, Anderson and van Wincoop (2003) demonstrate that the theory predicts the presence in the estimating equation of "multilateral resistance terms," defined as the average prices of goods in a particular economy inclusive of all trade barriers from all countries in the economy. Papers published at the time of Anderson and van Wincoop (2003) generally neglected any controls for these terms. Consequently, Anderson and van Wincoop (2003) conclude that the results in these papers suffer potentially from an omitted variable bias. To correct for this bias, Anderson and van Wincoop (2003) obtain an estimating equation that, with further refinements by Feenstra (2003) and Redding and Venables (2004), provides a straight-forward approach to estimating gravity equations through the use of country fixed effects.[2]

Largely contemporaneous with Anderson and van Wincoop (2003), the work of Andrew Rose uses gravity equations in order to explore the effect of a common currency on trade. In these papers, Rose and his coauthors present robust evidence to argue that a common currency has a large effect on increasing trade among the countries over which the currency is defined. Subsequent attempts to confirm or explain Rose's findings have met with mixed results (see Persson (2001) and Barro and Tenreyro (2007)).

These attempts, along with the increasing popularity of gravity equations to explain trade-related phenomena, of which the work of Rose is an example, brought greater scrutiny to the estimation of gravity equations. The publication of Santos Silva and Tenreyro (2006) has proven to be a seminal paper in this area. Section 3.2.3 will provide a more detailed explanation of the contribution of this paper. Suffice it to say at this moment that Santos Silva and Tenreyro (2006) prompted researchers to apply regression and estimation models other than OLS to gravity equations.

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[2] See Head and Mayer (2014) section 3 for a more thorough discussion.

Using these alternative models, Glick and Rose (2015) revisit the question of a common currency's effect on trade using historical data, the question analyzed in Glick and Rose (2002). The subtitle of Glick and Rose (2015), "A post-EMU mea culpa," gives away the conclusion. The authors state in the abstract that "Estimates of the currency union effect on trade are sensitive to the exact econometric methodology; the lack of consistent and robust evidence undermines confidence in our ability to reliably estimate the effect of currency union on trade."

This paper will offer a partial vindication for Rose's efforts. To do so, it calls on previous work exploring differences across the set of currency unions examined by Rose and his coauthors to show that even if a consistent statement robust to econometric technique about the magnitude of a currency union effect may not be obtainable, a consistent *ranking* of currency unions by the effect of each union's ability to increase trade internally *is* obtainable. Consider, for example, dollarized countries which have generally adopted the US dollar on an *ad hoc* basis with the two zones of the *Communauté financière africaine* (CFA), a set of long-standing, treaty-based monetary unions that emerged from French colonial rule. Klein (2005) shows that the experience of dollarized countries differs tremendously from what Rose and his coauthors found for currency unions overall (inclusive of dollarized countries). The CFA zones in Africa consist of the CEMAC and UEMOA, each with its own central bank and distinct fiat currency. Though these currencies are fixed at a 1:1 exchange rate, Gulde and Tsangarides (2008) argue that the two zones differ tremendously in most economic characteristics, including international trade. For example, CEMAC members export more petroleum products outside their zone than do UEMOA members export from their zone. Whitten (2013) examines different effects of currency unions on trade and links these effects to TFP and tariff differences across unions.

Overall, empirical studies deriving from the Anderson and van Wincoop (2003) tradition, such as Rose's, use aggregate trade as the dependent variable, an empirical application consistent with the theory used to derive the estimating equation. As the derivation in Anderson and van Wincoop (2003) assumes that "all goods are differentiated by place of origin," that "each region is specialized in the production of only one good," and requires summing purchases and sales over all destinations, Anderson and van Wincoop (2003) effectively "[suppresses] finer classifications of goods," in favor of aggregate trade (p. 174). That is to say, the theoretical justification supporting the use of the gravity equation for purposes of viewing the estimating equation as "structural gravity" and conducting comparative static questions in the context of "theory consistent estimation," (both terms used by Head and Mayer (2014)) takes as given that the trade examined is aggregate trade.

Nonetheless, applications of Anderson and van Wincoop (2003) work in the second, empirical strand often replace aggregate trade as the dependent variable with a "finer classification" of trade. This "finer classification" has often been agricultural trade.[3] Rauch (1999) examines a wider array of traded goods by disaggregating internationally-traded goods according to market characteristics of the goods: "[goods] traded on organizes exchanges, [goods] not traded on organized exchanges but nevertheless possessing ... 'reference prices', and all other commodities," (p. 8).

As in this paper, Grant and Lambert (2008) use a gravity model to compare and contrast the determinants of trade differ between agricultural and non-agricultural trade. However, this paper differs from Grant and Lambert (2008) in three important dimensions. First, Grant and Lambert (2008) use pooled OLS and linear panel models for their estimating equation, methods that may present inconsistent estimates as noted in Santos Silva and Tenreyro (2006). Second, the authors rely on GDP instead of the value of production and expenditure on agricultural and non-agricultural products (a point to be discussed below). Third, Grant and Lam-

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[3] E.g., see Sarker and Jayasinghe (2007), Lambert and McKoy (2009), Vollrath and Hallahan (2011), Villoria (2012), Philippidis et al. (2013), and Grant (2013)

bert (2008) depart from a common practice in the gravity literature by excluding any variables controlling for colonial and post-colonial status between countries. Though the effects of colonial status have weakened over time, as demonstrated by Head et al. (2010), this paper will demonstrate that colonial and post-colonial status remain important nonetheless.[4]

The above-mentioned studies all adapt the traditional gravity model to the cases of disaggregate trade by replacing aggregate trade flows with trade flows in the particular class of goods of interest to the authors. The authors of these papers do not adjust the GDPs of the countries concerned on the right-hand side of the estimating equation. Nor do the authors generally provide an alternative theory from which their estimating equation for disaggregated trade is derived. Head and Mayer (2014) state that for a gravity equation at an industry or disaggregated level to be structural, "one should [include as a regressor] gross production of traded goods ... and apparent consumption of goods. However, in practice GDP is often used as a proxy," (p. 138). For panel data or repeated cross-sections, this proxy is valid under the assumption of a constant expenditure share on or production share of a particular class of goods over the duration of the panel. Under this assumption, the share of GDP allocated to the expenditure on or production of the goods in question will be absorbed into the coefficient for GDP. To the extent that this assumption may not hold, the regression may produce inconsistent estimates.

This problem may be absent in purely cross-sectional studies such as Rauch (1999) and Villoria (2012). Rauch (1999) provides some theoretical justification for using GDP as a proxy for production of and expenditure on a particular class of goods by adapting the model used by Helpman (1987) to express output of expenditure on commodity k in countries i and j as a share of GNP in the respective

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[4] As an example, to understand the importance of colonial status in explaining agricultural trade between European countries and their former colonies, see Cosgrove (1994) for an explanation of the Lomé accords.

country (w_{ik} and w_{jk}). This assumption permits consistent estimates under the assumption that commodity k accounts for the same share of output across all countries. To the extent that this assumption is not true, Rauch states "I will assume that $w_{ik} + w_{jk}$ is absorbed into a multiplicative error term," (p. 11). As Rauch applies a logarithm to the gravity equation before estimating the equation by OLS, the results will be consistent provided that the expectation of the log of this multiplicative error term, inclusive of the country-varying sums of shares, is 0. However, as Santos Silva and Tenreyro (2006) note, justifying the consistency of OLS estimates from the gravity equation, particularly in the presence of heteroskedasticity, requires strong assumptions. These assumptions become even stronger if one must assume that the regression error, combined with $w_{ik} + w_{jk}$, come from a distribution the logarithm of whose mean is 0.

Anderson and Yotov (2010a) provide a theory and estimation strategy to overcome the above-mentioned problems. They derive a theory-consistent, estimable gravity equation for a particular class of goods. The dependent variables in this equation include the values of expenditure on and production of the particular class of goods concerned, not GDP. Anderson and Yotov (2010a) use the PPML method, as advocated by Santos Silva and Tenreyro (2006), to estimate the model, avoiding any concern for inconsistent estimates that may arise owing to the changing shares in GDP of the particular class of goods being analyzed.

Recent papers have incorporated the contribution of Anderson and Yotov (2010a) by using actual or approximate production data for agricultural and related goods. Sorgho and Larue (2014) use production data from the databases maintained by the *Food and Agriculture Organization* and *UNIDO*, following Anderson and Yotov (2010a) and Santos Silva and Tenreyro (2006). Bergstrand et al. (2015) examine manufacturing trade but rely on time-varying exporter and importer fixed effects to control for production, expenditure, and the MRTs. The authors' estimating equation differs noticeably from the estimating equation in this paper as it contains leads and lags of entry into economic integration agreements. Such a specification is not feasible in this paper as currency unions, often long-standing arrangements from decolonization in the 1960s (except for the EMU), pre-date the availability of sector-specific data (available only roughly from the 1970s forward). Consequently, the results in this paper and directly comparable to Bergstrand et al. (2015).

A few papers have managed to combine the various topics that this paper addresses (estimation techniques, currency unions, and specific sectors). Rose and Engel (2002) compare and contrast countries inside and outside of currency union countries on the basis of industrial concentration, using Herfindahl indices. Huchet-Bourdon and Cheptea (2011), instead of relying on aggregate GDP to proxy for agricultural production and expenditure, follow Baldwin et al. (2005) and use value-added figures from the World Bank's *World Development Indicators.* The authors use the PPML method to estimate their regression and find that introducing the euro increased agricultural trade within the EU by approximately 40%. Herwartz and Weber (2013) model a sector-specific gravity equation as a state-space model and estimate the model with the Kalman filter on data for select EMU countries and trading partners outside EMU for 94 different trade sectors.

3 Gravity equation, Estimation Methodologies, & Data

Although use of the gravity or gravity-like estimating equations dates back as far as Tinbergen (1962), the last decade has born witness to substantial improvements to the theoretical derivation and empirical estimation of the gravity equation, as summarized in Head and Mayer (2014). This section will describe how this paper incorporates, when possible, the advances made to gravity equation estimation. This section will also explain why the nature of the investigation undertaken here (a panel study of bilateral trade) makes the adoption of some procedures infeasible.

3.1 Structural Gravity Equation

Structural gravity, in the words of Head and Mayer (2014), refers to models sharing the following characteristics:

$$X_{ij} = \frac{Y_i}{\Pi_i} \frac{E_j}{P_j} b_{ij} \tag{1}$$

where $P_j = \sum_l \frac{Y_l b_{lj}}{\Pi_l}$ and $\Pi_i = \sum_l \frac{E_l b_{il}}{P_l}$. X_{ij} represents exports from i to j. Y_i denotes the value of goods produced in i. E_j^k represents expenditures on goods in class k. Π_i and P_j are the "multilateral resistance terms" (MRTs) that reflect the country-i and country-j prices of goods produced in all countries in the model. Controlling for these terms properly has become a key concern of the literature.[5] b_{ij} represents bilateral trading costs and is often modeled as a product of terms: $b_{ij} = \prod_l z_{ijl}^{\alpha_l}$. The bilateral resistance terms (BRTs) embody barriers to trade specific to a particular country pair. These barriers may observable or unobservable to the econometrician. Controls this resistance generally include geographic variables (distance, contiguity); socio-political variables (colonial and postcolonial relationships, common language); and international economic accords (free trade agreements and common currency). A following subsection addresses how this article controls for the BRTs.

This paper examines a particular element in the above-mentioned class, the sector-specific gravity equation developed by Anderson and Yotov (2010a):

$$X_{ij}^k = \frac{Y_i^k E_j^k}{Y^k} \left(\frac{b_{ij}^k}{\Pi_i^k P_j^k} \right)^{1-\sigma_k} \qquad \sigma_k > 1 \tag{2}$$

where Y^k is the sum of Y_i^k over all i: $Y^k \equiv \sum_i Y_i^k$.

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[5] See Head and Mayer (2014), section 3.

Estimating equation 2 on a panel of data, the methodology in this paper, requires confronting the following difficulties: 1) how to control for the multilateral resistances; 2) how to control for the bilateral resistances and levels of expenditure; 3) how to specify the conditional expectation; 4) how to prevent serial correlation from contaminating the coefficient and standard error estimates. The following subsections address these questions.

3.2 Estimation Methodologies

3.2.1 Controlling for the MRTs

To obtain consistent estimates of the regressors of interest, the estimating equation should contain controls for the multilateral resistance terms. In general, these controls are country-specific fixed effects where each country has potentially 2 different indicator variables: one for the country's role as an exporter and one for the country's role as an importer. Taking seriously the role that these variables play in controlling for changes in multilateral resistance requires that the fixed effects vary dynamically for a panel of data. However, this requirement can complicate the estimation, especially non-linear estimation, for datasets tracking trade over numerous years. As Head and Mayer (2014) say, "the estimation might run into computational feasibility issues due to the very large number of resulting dummies to be estimated," (p. 152). In some instances, the number of dummies is so large that it impedes estimation.

Is there a way to reduce the number of dummy variables necessary? This section demonstrates that an alternative approach for controlling the multilateral resistance terms leads to nearly-identical estimates for most variables of interest. In particular, this section shows that if importer and exporter fixed effects vary not yearly but less frequently, the coefficients and standard errors for most variables of interest are largely unaffected. Hence, researchers can obtain consistent estimates for coefficients from the gravity equation but with a fraction of the dummy variables required to control for the multilateral resistance terms. As a result, the computation time required to estimate parameters falls.

To demonstrate this claim, this section uses OLS to estimate the gravity equation on subsets of the data. This section uses OLS rather than non-linear estimation methods, owing to the speed and convenience of the former. As the controlling or not controlling for the multilateral resistance terms is a problem distinct from the objective function used to obtain parameter estimates (e.g., OLS vs. PPML), there is no reason to believe that the results would differ for PPML. For each type of trade (agriculture versus manufacturing), this section divides the available data into blocks of no more than 10 years.[6] For each block, this section reports estimation results from the gravity equation where the exporter and importer fixed effects are interacted not with the year dummy but with a `year_block` dummy. This `year_block` dummy groups years into a block of decreasing length. For example, the first length is 10 years: i.e., one fixed effect each for the country as an importer and as an exporter. The next length is 9 years: one fixed effect for each country's role (exporter and importer) for the first 9 years, followed by a different fixed effect for the remaining year. The next length is 8 years: one fixed effect for each country's role (exporter and importer) for the first 8 years, followed by a different fixed effect for the remaining 2 years. Once the length falls to 4 years, there are 3 fixed effects for each of the country's roles (exporter and importer), dividing the time span into two groups of four years and one group of two years. The most refined specification is interacting the importer effect and the exporter effect with a dummy for year.

`Put tables` 9 `and` 10 `approximately here.`

The results are in tables 9 and 10. For issues of space, only the broadest (1

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[6] The choice of 10 years is dictated by Stata's limit of no more than 11,000 independent variables for a regression.

fixed effect per decade per role) and the narrowest (1 fixed effect per year per role) results are included. Other results are available upon request. Overall, most coefficients and standard errors change little as the number of fixed effects changes, as evidenced from the correlations in table 1, summarizing the results in tables 9 and 10.[7] The correlations in table 1 do not carry any statistical meaning but serve merely to show whether an array of numbers is "similar" to another array of numbers.

Table 1: Correlation between effects of trade when year_block = 1 and when year_block = 10 (based on tables 9 and 10)

	1980s	1990s	2000s
Agriculture (including Y_{it}^k and $Y_{world,t}^k$)	0.65	0.51	0.84
Manufacturing (including Y_{it}^k and $Y_{world,t}^k$)	0.03	0.51	0.75
Agriculture (excluding Y_{it}^k and $Y_{world,t}^k$)	0.99	0.99	0.99
Manufacturing (excluding Y_{it}^k and $Y_{world,t}^k$)	0.99	0.99	0.99

The only exceptions are the coefficients for the value of production at the country and world level. It is not immediately clear why estimates for these variables should respond differently than do other estimates. Output certainly fluctuates over time in a way distinct from the fluctuations of the bilateral resistance terms (e.g., changes in membership for multilateral regional trade accords). Changing the frequency with which the dynamic multilateral resistance changes may have some unexpected influence on the coefficients of output. Additionally, this paper follows the practice of Anderson and Yotov (2010a) by using the importer's fixed effect to control simultaneously for the importer's expenditure as well as the multilateral resistance term, a practice that may contribute to the changing coefficients on Y_{it}. However, the coefficient for regional trade agreements is largely constant as the

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[7] Note that when comparing regression results across decades but within trade type and the value of `year_block` implies, in some cases, noticeably different effects (e.g., the estimate coefficient for contiguity). Though explaining why the effects may change over time is an interesting and important question, it is beyond the scope of this paper.

`year_block` dummy changes, even though membership in trade agreements does change over time. Nonetheless, if one excludes the coefficients for output, one can see that whether one follows a strict practice of changing the fixed effects every year or a looser practice of changing the fixed effects less frequently, the estimated coefficients for the variables of interest change little. Hence, this paper will "split the difference" and estimate a model based on blocks of 5 to 7 years, depending on the time required for the non-linear estimation to converge.

3.2.2 Controlling for the BRTs

The bilateral resistance terms embody barriers to trade specific to a particular country pair. A complete list and definitions of the variables used to control for the BRTs is given in appendix 2 of this paper. However, there still remains the possibility that other country-specific characteristics are responsible for explaining key determinants of trade. As Rose and van Wincoop note in their 2001 paper, despite a larget set of controls for BRTs in their paper, "there may still be some omitted factor that drives countries both to participate in currency unions and to trade more," (p. 387).

Given the "known unknown," of an assumed and unobserved pair-specific effect, panel methods may provide a way to obtain reliable estimates of the regressors. However, these methods are not without limitations. Random Effect models require that the unobserved, pair-specific effect be uncorrelated with the observed regressors, a strong assumption.[8] Fixed Effect models do not require this assumption, but eliminate the ability to estimate the effect of time-invariant determinants of trade (e.g., distance). The inability to estimate time-invariant regressors means that estimating the effect of currency unions, especially individual currency unions such as the ECCU or the UEMOA whose memberships have been largely stable over time, becomes infeasible or sensitive to the small number of entries and exits pro-

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[8] See Wooldridge (2002) p. 257

viding identification. Identifying the effect of unions other than the EMU (formed in 1999, providing a sharp pre- and post-effect) in a Fixed Effects framework relies on a small number of changes to membership as to question the reliability of estimates arising from such an identification strategy.[9] Hence, neither random Effects nor Fixed Effects can function as an appropriate estimation strategy.

A pooled estimation strategy with a time-invariant dummy variable for each country pair can absorb any time-invariant determinant of trade specific to a country pair but distinct from any of the other control variables for bilateral resistance. However, technical limitations pose a problem. Stata can handle only 11,000 independent variables for a regression.[10] Given that the number of country pairs greatly exceeds 11,000, this strategy is also infeasible. Hence, though the results in this paper use a rich set of regressors to control for bilateral resistances to trade, there remains the possibility that some unobserved, time-invariant, country-pair specific factors may be responsible for explaining the results found in this article.

3.2.3 Regression Model

Traditional estimation of a gravity equation of the form of equation 2 would first apply a functional transformation (the logarithm) to the equation in order to generate a linear form, upon which the researcher would perform OLS. Santos Silva and Tenreyro (2006) note that using such a procedure causes problems when the goal is to obtain consistent estimates of the marginal effect of the level of the independent variables on the level of the dependent variable. Consistency for the coefficients from an OLS regression requires that the conditional expectation of the error term from a regression of the form $\ln y_i = \sum_j \beta_j \ln x_{ij} + \epsilon_i$ be mean 0. To use the estimates from such a regression to make a consistent inference about the marginal effects of the level x on the level of y requires that the conditional expectation of the error

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[9] See appendix 1 for the entries into and exits from currency unions.

[10] See `http://www.stata.com/products/which-stata-is-right-for-me/`

term from a regression of the form $y_i = \exp(x_i'\beta) + \upsilon_i$ also be mean 0. Satisfying both conditional expectations requires either 1) a violation of Jensen's inequality or 2) a very particular form of heteroskedasticity in the variance-covariance matrix. As neither one of these conditions is likely, Santos Silva and Tenreyro (2006) note that OLS applied to a log-transformed gravity equation will result in inconsistent estimates. To remedy this problem, the authors propose using a pseudo-Poisson Maximum Likelihood (PPML) estimation procedure which estimates equation 2 in levels.

Head and Mayer (2014) compare and contrast the use of OLS, PPML, and Gamma PML.[11] As PPML has gained wide acceptance in the gravity literature (see Anderson and Yotov (2010a), Anderson and Yotov (2010b), Egger et al. (2011), Anderson and Yotov (2012), Glick and Rose (2015)), this paper will present results from both OLS and PPML estimation procedures in order to conclude whether a currency union effect can be estimated reliably regardless of the estimation procedure. Interpretations of the other variables will be based on the PPML results, owing to the econometric argument presented in Santos Silva and Tenreyro (2006) and to some questionable results from OLS. The estimating equations for each regression model are:

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[11] Head and Mayer (2014) also discuss a ratio-type estimation as used in Head et al. (2010) in order to reduce the number of parameters to estimate. In this procedure, the conditional expectation is neither a sum of logarithmic values or a product of levels but a ratio of variables where each country's variables are expressed relative to a reference importer and exporter. Although easy to implement, it is not clear if this method is also susceptible to the concern raised by Santos Silva and Tenreyro (2006). I.e., under what conditions for the error term and its variance-covariance matrix will the ratio estimation yield consistent estimates for the effect of the levels of independent variables on the level of the dependent variable?

$$\begin{aligned}\ln X_{ijt} &= \alpha_1 \ln Y_{it}^k + \alpha_2 \ln Y_t^k + \sum_{\tau\in T}\sum_{u=1}^{C} \gamma_{\tau u} d_{\tau u} + \sum_{\tau\in T}\sum_{v=1}^{C} \gamma_{\tau v} d_{\tau v} + \gamma CU_{ijt} \\ &\quad + \sum_{k=1}^{K} \beta_k b_{ijkt} + \epsilon_{ijt} \end{aligned} \tag{3}$$

$$\begin{aligned}X_{ijt} &= \exp\Bigg(\alpha_1 \ln Y_{it}^k + \alpha_2 \ln Y_t^k + \sum_{\tau\in T}\sum_{u=1}^{C} \gamma_{\tau u} d_{\tau u} + \sum_{\tau\in T}\sum_{v=1}^{C} \gamma_{\tau v} d_{\tau v} + \gamma CU_{ijt} \\ &\quad + \sum_{k=1}^{K} \beta_k b_{ijkt}\Bigg) + \upsilon_{ijt} \end{aligned} \tag{4}$$

where C is the number of countries. Let T denote the set of years and let $\tau \in T$ refer to the fact that a single country fixed-effect, indexed by τ, covers more years than just one. The $d_{\tau u}$ and $d_{\tau v}$ variables are indicator variables that control for the MRTs and equal 1 if $u = i$ and $v = j$, respectively, and if t corresponds to the group of years subsumed by τ. For the models where a common currency has a common effect in all currency unions (homogeneous integration), CU_{ijt} should be interpreted as a single indicator variable and γ as the scalar parameter of interest. For models where a common currency is allowed to have a different effect depending on the currency union in question (heterogeneous integration), CU_{ijt} should be interpreted as a vector of indicator variables and γ as the parameter vector of interest.

3.2.4 Serial Correlation

A concern with panels, particularly those that span decades, is the possibility for serial correlation of errors to bias coefficient and standard error estimates. Taking first differences of the data and estimating a model on those differences is a straightforward way to remove the serial correlation, provided that the model is linear (e.g., pooled OLS). First differencing is not an available option for a non-linear model such as PPML. To understand why, note that the estimating equation for PPML is of the form $y_{it} = \exp(x'_{it}\beta) + \upsilon_{it}$ and suppose that $\upsilon_{it} = \upsilon_{it-1} + \epsilon_{it}$ where ϵ_{it} is a mean zero, identically and independently distributed disturbance term. First differencing the estimating equation for such a series of data would imply estimating a vector of coefficients from the following estimating equation in order to eliminate the serial correlation in the error term: $\Delta y_{it} = \exp(x'_{it}\beta) - \exp\left(x'_{it-1}\beta\right) + \epsilon_{it}$. That is to say, this estimation procedure requires estimating β from a difference of exponentials, not an exponential of differences. Estimating β from such a regression would eliminate any contamination from the serial correlation of errors, but is a far more difficult task than doing so in a pooled OLS framework.

To address concerns about serial correlation in the data, this paper will appeal to the fixed-T, large-N asymptotics identified by Wooldridge (2002). If the cross-sectional size is large relative to the number of years covered by the data, then a pooled OLS estimator "is fully robust to arbitrary heteroskedasticity - conditional or unconditional - and arbitrary serial correlation across time (again, conditional or unconditional)," (pp 175-176). If the cross-sectional size is large relative to the number of years covered by the data, then a pooled Poisson estimation with a robust, sandwich-form variance covariance matrix will yield consistent coefficient estimates and standard errors "robust to the presence of serial correlation in the score and arbitrary conditional variances," without assuming that the data truly follow a Poisson distribution. (p. 670). In this paper, the number of years or T ranges from 32 to 36. The size of the cross-section, N, is the number of country pairs and ranges from 12,820 to 24,954. The values of N and T arguably satisfy the condition of fixed-T and large-N. All standard errors come from a fully-robust, sandwich form, variance-covariance matrix.

3.3 Data

The data are a panel of country pairs at yearly frequency from 1976 to 2011 (for agricultural data) and from 1980 to 2011 (for manufacturing data). Though the coverage of the two types of trade does not match perfectly, the additional four years of observations combined with time-varying MRTs as controls are unlikely to render the results uncomparable.[12] Export data come from UN COMTRADE, part of World Integrated Trade Solutions (WITS). When possible, agricultural and manufacturing are defined according to ISIC Revision 3, first introduced in 1988. For data from earlier years and for countries that maintained ISIC Revision 2 as the nomenclature for classifying industrial activity, the ISIC Revision 2 classification is used (1976-1995). Following the practice in Baldwin (2006) and in Huchet-Bourdon and Cheptea (2011), data on world agricultural and manufacturing output are proxied with data on value-added, recorded in the World Bank's *World Development Indicators.*

The control variables that compose the b_{ijt} include the currency union relationships, distance, other economic and geographic features, and colonial heritage. I use the CEPII database for Great Circle distances, augmented with information obtained from the CIA Factbook and from `http:\\www.timeanddate.com`.[13] As the CEPII data political relationships is time-invariant and as my period of interest spans the European decolonization of Africa and Southeast Asia, I use the independence dates provided by the Factbook in order to construct time-varying measures of political relationships. Information on regional trade agreements comes from the WTO's Regional Trade Agreement (RTA) database, augmented with information provided by the various secretariats of the RTAs on changes in RTA membership. Currency union membership comes from Glick and Rose (2002) and Glick and Rose

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[12] I thank a reviewer for mentioning this point.

[13] The CEPII database has been since replaced with a dataset constructed by Keith Head, Thierry Mayer and John Ries.

(2015), augmented by IMF staff reports and other publications. The term currency union in this literature refers not only to formal unions such as the EMU or CFA but also to countries that fix their own currency to or use the currency of another country, such as the use of the US Dollar in El Salvador and Liberia. A list of currency union members and a full list of definitions of the other proxy variables for bilateral resistance can be found in appendices 1 and 2, respectively. I refer to the Eurozone as countries in the EMU and those that have adopted the euro unilaterally (such as Macedonia). Appendix 1 lists the composition of the currency unions.

4 Results

4.1 Agriculture versus manufacturing

Given the multiple dimensions along which agricultural and manufacturing trade can be compared (OLS estimates versus PPML estimates, currency union coefficients, other covariates), this results section will first explain why the PPML results appear more plausible. Next, the section will compare the effect of a common currency under a homogeneous level of integration with the effect under a heterogeneous level of integration. Finally, the section will address differences between agriculture and manufacturing trade in other covariates.

`Tables 11 and 12 approximately here.`

4.1.1 PPML versus OLS

First, consider the differences between regression models between each type of trade. Comparing the OLS and PPML results overall for each type of good traded show that the results are generally similar. The correlation between OLS and PPML coefficients for agricultural trade is 0.831 while the same correlation for manufacturing trade the correlation is 0.703 (see table 2). These correlations do not carry any statistical meaning but serve merely to show whether an array of numbers is "similar" to another array of numbers.[14] However, the theoretical argument proposed by Santos Silva and Tenreyro (2006) as well as some individual estimates suggest that the inconsistency introduced by the logarithmic transformation of the gravity equation for OLS has some serious effects. First, note that OLS yields a significantly negative effect for manufacturing trade in the EMU. PPML returns a negative but small (in absolute value) and insignificant effect for the same variable. While the true magnitude of the European single currency's effect on trade is debatable, to believe that introducing the euro has *reduced* trade within Europe seems incredible in the literal sense of the word.[15] Though the PPML results imply that membership in the dollarized zone predicts a lower level of manufacturing trade, this result is more plausible. A country adopting the US dollar ad hoc often does so for political or historical reasons (e.g., Panama) or to combat severe inflation (e.g., Ecuador). It is not surprising that the consequences of dollarization under these idiosyncratic factors would predict a lower level of trade.

Second, note that OLS implies that distance has a larger (in absolute value) effect on manufacturing trade than does it have on agricultural trade. In contrast, PPML implies a larger effect for distance on agricultural trade than on manufacturing trade, a finding consistent with existing literature regarding the importance of distance for perishable goods within agricultural goods.

To explain why distance likely matters more for agricultural goods than for manufactured goods, note that the transportation costs needed to maintain quality in transit for perishable agricultural goods (e.g. cut flowers, fresh produce, dairy products) may manifest themselves in a gravity equation through a larger coefficient

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[14] This table is based on columns 3 and 4 from tables 11 and 12.

[15] Glick and Rose (2015) also find a negative coefficient for the EMU from OLS, but for aggregate trade, and argue in favor of the PPML results in their paper.

for distance than for other coefficients, unlike the results of manufactured goods, which tend to be more durable over time (see Wang et al. (2000) and Djankov et al. (2010)).

Given the unlikely results for the eurozone and distance, the PPML results appear to be more reliable than are the OLS estimates. Subsequent of discussion of individual components of the BRT (other than currency unions) will be based on the PPML results.

Table 2: Correlation between all coefficients

	Ag PPML	Manuf OLS	Manuf PPML
Ag OLS	0.831	0.497	
Ag PPML			0.263
Manuf OLS			0.703

4.1.2 Currency union coefficients

For both agricultural and manufacturing trade, the effect of a common currency differs noticeably across estimation methods and across types of trade under a homogeneous representation of currency unions (see columns 1 and 2 of tables 11 and 12). OLS implies no significant effect for agricultural trade but a significant (if relatively small when compared to other regressors) effect for agricultural trade. OLS implies a large and significant effect of a common currency on manufacturing trade while the PPML procedure yields and insignificant effect with a negative point estimate.

Do these results support the conclusion of Glick and Rose (2015) that the currency union effect cannot be estimated consistently across estimation procedures? While disaggregating across trade types does not lead to consistent estimation, disaggregating across the individual currency unions *does* lead to a consistent ranking of the effectiveness of each currency union at promoting trade (see columns 3 and 4 of tables 11 and 12). Although each estimation method yields different, sometimes noticeably different, coefficient estimates for the individual currency unions (e.g., the EMU estimates for manufacturing trade), these estimates follow a pattern; if OLS yields a high parameter estimate for one type of trade, PPML will also yield a high parameter estimate for the same type of trade. The correlation between the OLS currency union parameter estimates and the PPML currency union parameter estimates for agricultural trade is 0.92 (see table 3). The corresponding correlation for manufacturing trade is 0.88. Hence, though we may be unable to state definitively what is the quantitative significance of a common currency on trade, we may be able to state which currency union arrangements promote trade to a greater or lesser degree. Note also that for a fixed estimation technique, the correlation between the coefficients between agriculture and manufacturing are rather weak. These relatively small correlations justify disaggregation in trade as necessary to understand how each currency union benefits from its own common currency.

Table 3: Correlation between currency union coefficients

	Ag PPML	Manuf OLS	Manuf PPML
Ag OLS	0.918	0.638	
Ag PPML			0.538
Manuf OLS			0.88

Tables 11 and 12 demonstrate that the individual effects of some unions are quite large (e.g., the East Caribbean Currency Union (ECCU) for manufacturing), consistent with Rose's earlier findings. Others are more modest (e.g., the West African Economic and Monetary Union (UEMOA) for agriculture). For some unions, such as the dollarized zone and agriculture, the quantitative significance differs noticeably between estimation methods. Using the formula $\left(\exp\left(\hat{\beta}\right) - 1\right) \times 100$ to obtain the percentage increase in trade associated with a common currency, OLS implies that the US increases agricultural trade among dollarized countries trade by approximately 60%. The PPML result implies a larger increase, 80%. Hence, different estimation techniques imply different effects on levels. The rankings, as indicated in table 3, are more robust to estimation technique.

Differences exist among unions and between goods. The ECCU, the UEMOA, the Australia zone, the rand zone, the Krone zone, and India-Bhutan all have effects on manufacturing that are larger than are the effects on agricultural trade for both OLS and PPML. The dollarized zone has insignificant effects under OLS but significant effects under PPML. However, the dollarized zone has a larger effect in absolute value on manufacturing trade than on agricultural trade, though the manufacturing effect is negative while the agricultural effect is positive. Pick and Carter (1994) note that the US dollar commonly serves as the currency to price agricultural products, making the result in table 11 not too surprising.

The Eurozone is an exception to the stylized fact that effects of a common currency on manufacturing trade are larger than are the effects on agricultural trade. The result for the Eurozone suggests that agricultural trade within the union increased by 27%. This result is in line with the finding of 40% in Huchet-Bourdon and Cheptea (2011). Superficially, the Eurozone result suggests that introducing the euro has led to greater trade within the EMU of agricultural products. However, the Common Agricultural Policy (CAP) of the EU may work as a confounding factor in the analysis. The EU implemented a reform package entitled "Agenda 2000: for a stronger and wider Union" beginning in 1997 and finishing in 1999.[16] The euro was introduced formally in 1999 when exchange rates for the initial member countries were fixed. Notes and coins were introduced in 2002.[17] The CAP is a subsidy that overlaps almost entirely with the EMU, the currency union in question. This poses a problem unlike that encountered with agricultural subsidies implemented by individual countries. Country-specific agricultural subsidies that may distort trade patterns are controlled for implicitly through the time-varying country effects which

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[16] See `http://europa.eu/legislation_summaries/enlargement/2004_and_2007_enlargement/l60001_en.htm`, accessed 3 January 2015

[17] See `http://ec.europa.eu/economy_finance/euro/index_en.htm`, accessed 3 January 2015

also control for the price indices. These controls are separate from the controls for the currency union to which a country may belong.

4.1.3 Other bilateral resistance terms

Focusing on the PPML estimates over the OLS estimates, the estimated coefficients of bilateral resistance are larger in magnitude for manufacturing trade than for agricultural trade. The exceptions to this pattern are distance (as mentioned earlier); the colonizer-colonized variable (for a pair of countries, one country was a former colonizer of the other country); the transitional colony variable (a country pair where one country has gained independence from a colonial empire while the other country has not); and the generic common currency variable for homogeneous integration. All of these variables are significant for agricultural trade while all of them, except the transitional colonialism variable, are insignificant for manufacturing trade. Overall, manufacturing trade and agricultural trade appear to have very different determinants. Table 2 shows that for a fixed regression model, the correlation between coefficients across trade types suggests a weakly positive association (agricultural trade: 0.497) or a nearly non-existent association (manufacturing trade: 0.263). Hence, the results bear out the message of Anderson and Yotov (2010a), Anderson and Yotov (2010b), and Anderson and Yotov (2012). Trade costs and country-pair features that mitigate those trade costs differ substantially across the types of goods traded.

Larger estimates of bilateral resistances for manufacturing trade than for agricultural trade may not be expected for three reasons. First, agricultural products tend to be homogeneous across producers. Hence, the gravity equation, often motivated through the supposition of a CES objective function conveying a “love-of-variety”-like motivation, may be less appropriate for agricultural goods than for manufacturing goods. A consequence of the homogeneity in product characteristics is a greater emphasis on price for customers when selecting a producer rather than other considerations (e.g., a common language).

Second, a large share of agricultural products may rely more heavily on natural endowments which tend to be country-specific in ways distinct from a comparable share of manufactured goods.[18] As endowments of natural resources tend to be country-specific, the flow of trade within a given country pair may depend more heavily on particular aspects of one of those countries (through natural endowments) than on a characteristic of the pair itself.[19] Consequently, the greater importance of country-specific factors means controls for the multilateral resistance terms will absorb these effects in the gravity equation, leaving less explanatory power (and smaller bilateral resistance coefficients) for agricultural trade.

A third reason, that also explains the larger effect of a common currency for agricultural trade than for manufacturing trade, arises from the greater emphasis on price for agricultural than for manufacturing trade. A worldwide currency such as the U.S. dollar is generally the currency used for pricing and then purchasing agricultural goods (see Pick and Carter (1994)). If neither importer nor exporter use the currency in which the agricultural goods are priced, exchange rate uncertainty can affect both buyer and seller (in contrast to affecting the seller for producer-currency pricing or the buyer for local-currency pricing). As the trade-enhancing effects of a common currency are traditionally believed to arise from the elimination of exchange rate uncertainty, the removal of uncertainty from both ends of the

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[18] I thank a referee for pointing out this distinction.

[19] Heerman et al. (2015) propose a model that recognizes the ability of changes in a single exporter's trade costs to alter the ratio of the market shares of two other exporters in some other market, a phenomenon that they label as a violation of the "independence of irrelevant exporters," (IIE) an international trade counterpart to the independence of irrelevant alternatives property commonly used in the discrete choice literature. The violation of IIE owing to the importance of natural resources in a single country is a point that Heerman et al. (2015) justify by citing Eaton and Kortum (2002) and their caution against using their Ricardian model in a context where natural resources are critical for determining trade.

transaction should increase trade to a larger extent than does removing uncertainty from just one end of the transaction (see Anderson and van Wincoop (2004)).

Table 4: Correlation between non-currency union coefficients

	Ag PPML	Manuf OLS	Manuf PPML
Ag OLS	0.902	0.504	
Ag PPML			-0.483
Manuf OLS			0.307

4.2 Individual manufacturing industries

The results in the previous section indicate distinct differences between the determinants of agricultural trade and the determinants of manufacturing trade. The results also show that there exists a common ranking of currency unions, independent of regression model, by the magnitude of their effect on trade.

`Put tables 13 through 16 approximately here.`

Hence, this section will explore the robustness of this ranking within manufacturing trade by examining four industries: textiles, machinery, chemicals, and food, tobacco, and beverages. These are industries for which the value of production can be computed over time and across countries using data from the World Bank's *World Development Indicators*. The years covered range from 1991 to 2006. Although these industries all belong to the manufacturing sector, textiles and food, beverage, and tobacco rely on inputs more closely linked with agricultural production than do the machinery and chemicals industries. This subsection and its results closely resemble those in Anderson and Yotov (2010b), but pays particular attention to the importance of currency union effects. Owing to difficulties in obtaining convergence for the PPML estimates, the MRTs are time-invariant fixed effects where each country has up to two fixed effects for the two roles it might play (importer and exporter). There is a separate, year fixed effect.

Table 5: Correlation between currency union coefficients across estimation procedures (OLS versus PPML)

Chemicals	Food, beverage, & tobacco	Machinery	Textiles
0.773	0.884	0.871	0.912

Food, beverage, and tobacco and textiles have currency union effects that tend to be the most robust regardless of estimation technique (see table 5). The correlation between the OLS and PPML results for food et al is 0.884 while the correlation for textiles is 0.912. The correlation for machinery, however, is also quite high: 0.871. Chemicals has a weaker correlation, 0.773. Note that except for chemicals, these correlations are roughly equal to or larger the correlation between the PPML and OLS currency union coefficients for agricultural and manufacturing trade (compare table 5 with table 3). Hence, disaggregating trade can lead to a reliable ranking, independent of estimating method, of the effects of individual currency unions.

Table 6: Correlation for all coefficients between Agriculture, Manufacturing, and individual industries

	Chemicals	Food, beverage, & tobacco	Machinery	Textiles
Agriculture	0.42	0.492	0.498	0.563
Manufacturing	0.956	0.849	0.96	0.632

How do the results of individual industries compare with the estimates for agricultural trade and manufacturing trade overall? Consider the PPML results as the previous section established the greater plausibility of PPML over OLS. Overall and unsurprisingly, the coefficients of individual manufacturing industries are more closely correlated with the coefficients from manufacturing trade (see table 6). Textiles is an exception as it appears to be nearly equally and weakly related to both agricultural trade overall and manufacturing trade overall.

The closeness between manufacturing and individual manufacturing industries continues when examining just the coefficients of the currency unions (see table

Table 7: Correlation for currency union coefficients between Agriculture, Manufacturing, and individual industries

	Chemicals	Food, beverage, & tobacco	Machinery	Textiles
Agriculture	0.291	0.308	0.373	0.448
Manufacturing	0.94	0.762	0.952	0.473

7). Looking at individual unions and industries, the UEMOA, CEMAC, Kroner zone, and India-Bhutan generally have larger and more significant coefficients in chemicals and manufacturing than in the agricultural-related industries. The coefficients for manufacturing are negative or insignificant for the dollar and euro zones, respectively, but are positive and significant for food et al.

Table 8: Correlation for non-currency union coefficients between Agriculture, Manufacturing, and individual industries

	Chemicals	Food, beverage, & tobacco	Machinery	Textiles
Agriculture	0.585	0.926	0.731	0.876
Manufacturing	0.832	0.901	0.911	0.958

The correlation between other covariates for individual industries and the larger categories of trade offers a slightly different story (see table 8). The correlation between manufacturing trade and each of the individual industries is high. However, the correlations between agricultural trade and each of food et al and textiles are also high. As both of these manufacturing industries rely on agricultural inputs to a large extent, the high correlation may not be too surprising. However, the correlation of 0.731 between machinery and agriculture is less easily explained. Food et al appears to be the industry best-explained by the gravity model in the sense that the coefficients of the regressors are generally larger and more significant than are the regressors for the other industries. Colonial-era variables tend to be the most important for food et al than for the other industries. The regional trade agreement indicator has a large coefficient for machinery and textiles, though the estimate for food et al coefficient is also sizable.

5 Conclusion

A series of papers by Andrew Rose showed that a common currency has a large effect on increasing trade within the area covered by the currency. Recently, Rose has called into question the reliability of this conclusion, as new techniques have emerged for estimating gravity equations of bilateral trade.

A common currency is, from the perspective of a gravity equation, a bilateral concept that can enhance (or possibly even hinder) trade between the countries sharing it. As Anderson and Yotov (2010a) note, the effect of these bilateral resistances likely differ across types of goods. Consequently, examining disaggregate trade instead of aggregate trade may provide the reliable estimate of a common currency on trade sought by Rose and his coauthors.

Disaggregating trade alone is not sufficient to obtain a reliable estimate of a currency union when the effect of a common currency on trade is thought to be homogeneous (common across all currency unions), regardless of econometric technique. Disaggregating the universe of currency unions into unions with individual effects provides a reliable ranking of currency unions, independent of estimation method, by the effect that each union's currency has on increasing trade within the union. These rankings differ across the type of goods traded. The eurozone, dollar zone, and the rand zone have large positive effects for agricultural trade but very different effects for manufacturing trade. The East Carribean Currency Union has a large effect for manufacturing trade but an insignificant effect for agricultural trade. The reliability of the ranking independent of estimation method carries over to specific industries within manufacturing.

Although this paper establishes a clear and consistent ranking of the effects of a currency union on trade, it does not provide a meaningful explanation as to why currency unions are ranked in this particular way. I.e., this paper does not answer why dollarized countries should have a strong effect for agriculture but not for manufacturing, or why the reverse is true for the East Caribbean Currency Union. As these results come from a regression that controls for individual, time-varying country effects, any explanation should arise from institutional factors or multi-country sources. Future research will examine such factors in order to provide a more thorough understanding of currency unions and their effects on trade.

Appendix 1: Currency unions and their composition

East Caribbean Currency Union (ECCU)

Antigua and Barbuda
Barbados (1965-1972)
Dominica
St. Kitts and Nevis
St. Lucia
St. Vincent and the Grenadines

West African Economic and Monetary Union (UEMOA)

Benin
Burkina Faso
Côte d'Ivoire
Guinea-Bissau (1997-)
Mali
Mauritania (1960-1973)
Niger
Sénégal
Togo

Communauté Économique et Monétaire de l'Afrique Centrale (CEMAC)

Cameroon
Central African Republic
Chad
Congo, Rep.
Equatorial Guinea (1985-)
Gabon
Madagascar (1960-1972)

European Monetary Union (EMU) / Euroized

Austria (1999-)
Belgium (1999-)
Cyprus (2004-)
Estonia (2004-)
France (1999-)
Finland (1999-)
Germany (1999-)
Greece (2001-)
Ireland (1999-)

Italy (1999-)
Latvia (2005-)
Luxembourg (1999-)
Macedonia (2002-)
Malta (2005-)
Netherlands (1999-)
Portugal (1999-)
Slovak Republic (2006-)
Slovenia (2007-)
Spain (1999-)

Dollarized countries

American Samoa
The Bahamas (1973-)
Bermuda
Ecuador (2000-)
El Salvador (2001-)
Guam
Liberia
Marshall Islands
Federated States of Micronesia
Northern Mariana Islands
Palau
Panama
Puerto Rico
Virgin Islands (U.S.)

India

Bhutan
India

Denmark

Denmark
Færoe Islands
Greenland

Australia

Australia
Kiribati
Tonga (until 1990)

Appendix 2: Control variables and definitions

- CU_{ijt} is 1 if countries i and j belong to the same currency union in time t.
- $ECCU_{ijt}$ is 1 if countries i and j belong to the East Caribbean Currency Union in time t.
- $UEMOA_{ijt}$ is 1 if countries i and j belong to the *Union Economique et Monétaire Ouest Africaine* (West African Economic and Monetary Union) in time t.
- $CEMAC_{ijt}$ is 1 if countries i and j belong to the *Communauté Economique et Monétaire de l'Afrique Centrale* (Central African Economic and Monetary Coummnity) in time t.
- $AUSTRALIA_{ijt}$ is 1 if countries i and j both use the Australian dollar at time t.
- EMU_{ijt} is 1 if countries i and j both use the Euro at time t.
- $BHUTAN_{ijt}$ is 1 if $i =$ India and $j =$ Bhutan (or vice versa) at time t.
- $BRUNEI_{ijt}$ is 1 if if $i =$ Singapore and $j =$ Brunei (or vice versa) at time t.
- ln_dist_{ijt} is the log of Great Circle distance between countries i and j.
- $contig_{ijt}$ is 1 if countries i and j share a border.
- $comlang\_off_{ijt}$ is 1 if countries i and j share a common or official language.
- rta_{ijt} if 1 if countries i and j adhere to a trade agreement in time t.
- $colonizer\_variant_{ijt}$ is 1 if i maintains or has maintained some level of sovereignty over j up to time t, 0 otherwise. Sovereignty could be of an administrative nature, such as the US relationship with Guam, or complete sovereignty, such as France's control over Algeria before 1962.
- $comcol_{ijt}$ is 1 if countries i and j are both under the same, third-country colonizer in time t.
- $postcol_{ijt}$ is 1 if countries i and j were both under the same, third-country colonizer before time t but are now independent.
- $transcol_{ijt}$ is 1 if countries i and j were both under the same, third-country colonizer before time t but only 1 country has left the colonial relationship.

Appendix 3: Results

Table 9: Estimated coefficients under differing frequencies of MRT for Agriculture trade (estimated with OLS)

	1980s				1990s				2000s			
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
Countries share a common currency	0.314	(0.177)	0.316	(0.18)	-0.102	(0.549)	-0.046	(0.79)	-0.183	(0.165)	-0.198	(0.14)
ln Y_{it}^k	$0.128^{\dagger}$	(0.078)	1.564	(0.597)	0.458^{**}	(0.0)	-9.506^{**}	(0.003)	0.351^{**}	(0.0)	-0.102	
ln $Y_{world,t}^k$	0.388^{**}	(0.0)	-1.053	(0.564)	-0.620^{**}	(0.0)	-7.421	(0.193)	0.772^{**}	(0.0)	2.518	
ln Distance	-1.193^{**}	(0.0)	-1.206^{**}	(0.0)	-1.181^{**}	(0.0)	-1.209^{**}	(0.0)	-1.301^{**}	(0.0)	-1.306^{**}	(0.0)
Pair belongs to a Regional trade Accord	0.397^{**}	(0.002)	0.378^{**}	(0.004)	0.252^{**}	(0.0)	0.246^{**}	(0.0)	0.390^{**}	(0.0)	0.391^{**}	(0.0)
Colonizer-colonized relationship	1.415^{**}	(0.0)	1.407^{**}	(0.0)	1.457^{**}	(0.0)	1.459^{**}	(0.0)	1.475^{**}	(0.0)	1.485^{**}	(0.0)
Country pair transitioning from colonialism	0.072	(0.761)	0.089	(0.738)	0.326	(0.264)	0.292	(0.344)	$0.803^{\dagger}$	(0.086)	$0.785^{\dagger}$	(0.095)
Countries were colonies of same country	0.297^{**}	(0.002)	0.294^{**}	(0.002)	0.322^{**}	(0.0)	0.320^{**}	(0.0)	0.450^{**}	(0.0)	0.455^{**}	(0.0)
Countries are contiguous	0.529^{**}	(0.0)	0.513^{**}	(0.001)	1.052^{**}	(0.0)	1.051^{**}	(0.0)	1.257^{**}	(0.0)	1.257^{**}	(0.0)
Shared common or official language	0.456^{**}	(0.0)	0.465^{**}	(0.0)	0.332^{**}	(0.0)	0.346^{**}	(0.0)	0.404^{**}	(0.0)	0.410^{**}	(0.0)
Number of dummy variables per decade	1		10		1		10		1		10	
Number of observations	36513				71289				58894			
Number of country pairs	6921				13944				14392			

Table 10: Estimated coefficients under differing frequencies of MRT for Manufacturing trade (estimated with OLS)

	1980s				1990s				2000s			
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
Countries share a common currency	1.955^{**}	(0.0)	1.937^{**}	(0.0)	1.176^{**}	(0.0)	1.204^{**}	(0.0)	0.160	(0.151)	0.159	(0.159)
ln Y_{it}^{k}	0.499^{**}	(0.0)	16.837	(0.676)	0.491^{**}	(0.0)	0.286	(0.763)	0.333^{**}	(0.0)	0.016	(0.0)
ln $Y_{world,t}^{k}$	0.775^{**}	(0.0)	-18.527	(0.695)	1.902^{**}	(0.0)	-1.032	(0.684)	1.173^{**}	(0.0)	-0.719	(0.0)
ln Distance	-1.700^{**}	(0.0)	-1.709^{**}	(0.0)	-1.625^{**}	(0.0)	-1.639^{**}	(0.0)	-1.719^{**}	(0.0)	-1.721^{**}	(0.0)
Pair belongs to a Regional trade Accord	-0.129	(0.353)	-0.160	(0.264)	0.326^{**}	(0.0)	0.291^{**}	(0.0)	0.540^{**}	(0.0)	0.539^{**}	(0.0)
Colonizer-colonized relationship	1.118^{**}	(0.0)	1.104^{**}	(0.0)	1.343^{**}	(0.0)	1.344^{**}	(0.0)	1.261^{**}	(0.0)	1.258^{**}	(0.0)
Country pair transitioning from colonialism	-0.044	(0.86)	-0.148	(0.579)	0.315	(0.113)	0.249	(0.226)	$0.486^{\dagger}$	(0.051)	0.497^{*}	(0.047)
Countries were colonies of same country	0.436^{**}	(0.0)	0.440^{**}	(0.0)	0.676^{**}	(0.0)	0.678^{**}	(0.0)	0.728^{**}	(0.0)	0.728^{**}	(0.0)
Countries are contiguous	0.195	(0.191)	0.178	(0.242)	0.696^{**}	(0.0)	0.692^{**}	(0.0)	0.784^{**}	(0.0)	0.782^{**}	(0.0)
Shared common or official language	0.567^{**}	(0.0)	0.583^{**}	(0.0)	0.552^{**}	(0.0)	0.554^{**}	(0.0)	0.571^{**}	(0.0)	0.571^{**}	(0.0)
Number of dummy variables per decade	1		10		1		10		1		10	
Number of observations	44815				107917				98961			
Number of country pairs	8094				19363				21394			

Table 11: Estimated coefficients for agricultural trade

	(1) OLS		(2) PPML		(3) OLS		(4) PPML	
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
East Caribbean Currency Union					-0.516	(0.245)	-0.308	(0.76)
West African Economic & Monetary Union					0.097	(0.771)	1.096*	(0.022)
Central African Economic & Monetary Union					-1.341**	(0.009)	-0.814	(0.248)
Rand zone					0.137	(0.87)	1.967*	(0.043)
Australia zone					-0.385	(0.607)	-0.460	(0.283)
Dollarized zone					0.466	(0.161)	0.588**	(0.008)
Euro zone					0.347**	(0.008)	0.238**	(0.002)
Krone zone (Denmark)					2.986**	(0.0)	3.447**	(0.0)
India-Bhutan					0.858	(0.595)	2.153**	(0.007)
Countries share a common currency	-0.035	(0.78)	0.275**	(0.002)				
ln Y_{it}^{k}	0.406**	(0.0)	0.402**	(0.0)	0.407**	(0.0)	0.506**	(0.0)
ln $Y_{world,t}^{k}$	0.311**	(0.0)	0.624**	(0.0)	0.312**	(0.0)	0.398**	(0.0)
ln Distance	-1.225**	(0.0)	-0.921**	(0.0)	-1.225**	(0.0)	-0.944**	(0.0)
Pair belongs to a Regional trade Accord	0.363**	(0.0)	0.472**	(0.0)	0.356**	(0.0)	0.439**	(0.0)
Colonizer-colonized relationship	1.480**	(0.0)	0.566**	(0.0)	1.472**	(0.0)	0.803**	(0.0)
Countries are colonies of the same country	3.187**	(0.0)	2.093**	(0.001)	3.253**	(0.0)	2.141**	(0.0)
Country pair transitioning from colonialism	0.400†	(0.077)	-0.764*	(0.043)	0.425†	(0.063)	-0.738†	(0.051)
Countries were colonies of same country	0.357**	(0.0)	0.279**	(0.009)	0.367**	(0.0)	0.280*	(0.01)
Countries are contiguous	1.012**	(0.0)	0.498**	(0.0)	1.031**	(0.0)	0.511**	(0.0)
Shared common or official language	0.388**	(0.0)	-0.019	(0.841)	0.380**	(0.0)	-0.025	(0.783)
Number of observations	175365		288366		175365		288366	
Number of country pairs	17461		24954		17461		24954	

Table 12: Estimated coefficients for manufacturing trade

	(1) OLS		(2) PPML		(3) OLS		(4) PPML	
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
East Caribbean Currency Union					2.321**	(0.0)	3.227**	(0.0)
West African Economic & Monetary Union					1.905**	(0.0)	2.016**	(0.0)
Central African Economic & Monetary Union					1.479**	(0.002)	1.829**	(0.0)
Rand zone					0.397	(0.317)	1.285**	(0.005)
Australia zone					1.898*	(0.032)	2.340**	(0.0)
Dollarized zone					0.055	(0.772)	-0.893**	(0.002)
Euro zone					-0.841**	(0.0)	-0.023	(0.703)
Krone zone (Denmark)					3.468**	(0.0)	4.373**	(0.0)
India-Bhutan					2.780†	(0.06)	3.156**	(0.0)
Countries share a common currency	0.753**	(0.0)	-0.055	(0.356)				
ln Y^k_{it}	0.376**	(0.0)	0.613**	(0.0)	0.436**	(0.0)	0.736**	(0.0)
ln $Y^k_{world,t}$	1.079**	(0.0)	1.098**	(0.0)	1.172**	(0.0)	0.736**	(0.0)
ln Distance	-1.671**	(0.0)	-0.738**	(0.0)	-1.677**	(0.0)	-0.751**	(0.0)
Pair belongs to a Regional trade Accord	0.414**	(0.0)	0.619**	(0.0)	0.427**	(0.0)	0.584**	(0.0)
Colonizer-colonized relationship	1.280**	(0.0)	0.151	(0.255)	1.291**	(0.0)	0.234	(0.121)
Coutnries are colonies of the same country	0.326	(0.438)	-3.633**	(0.0)	0.326	(0.412)	-3.619**	(0.0)
Country pair transitioning from colonialism	0.277	(0.112)	-0.431*	(0.04)	0.287	(0.1)	-0.397†	(0.062)
Countries were colonies of same country	0.653**	(0.0)	0.337**	(0.0)	0.645**	(0.0)	0.335**	(0.0)
Countries are contiguous	0.629**	(0.0)	0.450**	(0.0)	0.607**	(0.0)	0.454**	(0.0)
Shared common or official language	0.560**	(0.0)	0.275**	(0.0)	0.555**	(0.0)	0.270**	(0.0)
Number of observations	254871		268571		254871		268571	
Number of country pairs	23579		24954		23579		24954	

Table 13: Estimated coefficients for chemical trade

	(1) OLS		(2) PPML		(3) OLS		(4) PPML	
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
East Caribbean Currency Union					1.428*	(0.043)	4.845**	(0.0)
West African Economic & Monetary Union					1.802**	(0.0)	2.174*	(0.018)
Central African Economic & Monetary Union					3.333**	(0.0)	3.233**	(0.0)
Australia zone					2.000**	(0.0)	2.493**	(0.0)
Dollarized zone					-0.111	(0.669)	-0.841**	(0.0)
Euro zone					0.178	(0.107)	-0.045	(0.628)
Krone zone (Denmark)					4.328**	(0.0)	5.574**	(0.0)
India-Bhutan					-0.068	(0.807)	2.522**	(0.0)
Singapore-Brunei					2.611**	(0.0)	1.802**	(0.002)
Countries share a common currency	0.517**	(0.0)	-0.118	(0.211)				
ln Y_{it}^{k}	0.318**	(0.0)	0.199**	(0.008)	0.315**	(0.0)	0.534**	(0.0)
ln $Y_{world,t}^{k}$	1.732**	(0.0)	1.207	(0.998)	1.747**	(0.0)	1.394**	(0.0)
ln Distance	-1.744**	(0.0)	-0.968**	(0.0)	-1.757**	(0.0)	-0.958**	(0.0)
Pair belongs to a Regional trade Accord	0.353**	(0.0)	0.240**	(0.003)	0.338**	(0.0)	0.238**	(0.004)
Colonizer-colonized relationship	1.077**	(0.0)	-0.344*	(0.044)	1.079**	(0.0)	-0.049	(0.794)
Country pair transitioning from colonialism	0.605*	(0.034)	-0.061	(0.787)	0.632*	(0.03)	-0.035	(0.871)
Countries were colonies of same country	0.592**	(0.0)	0.279*	(0.036)	0.591**	(0.0)	0.259*	(0.045)
Countries are contiguous	0.743**	(0.0)	0.147*	(0.045)	0.751**	(0.0)	0.158*	(0.045)
Shared common or official language	0.579**	(0.0)	0.311**	(0.008)	0.565**	(0.0)	0.297*	(0.011)
Number of observations	107355		149502		107355		149502	
Number of country pairs	12820		16984		12820		16984	

Table 14: Estimated coefficients for food, beverage, & tobacco trade

	(1) OLS		(2) PPML		(3) OLS		(4) PPML	
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
East Caribbean Currency Union					1.942*	(0.025)	2.818**	(0.001)
West African Economic & Monetary Union					0.534†	(0.057)	0.031	(0.94)
Central African Economic & Monetary Union					2.182**	(0.0)	1.387**	(0.008)
Australia zone					2.269**	(0.0)	1.426**	(0.001)
Dollarized zone					0.491†	(0.064)	0.508**	(0.004)
Euro zone					0.289*	(0.012)	0.182*	(0.042)
Krone zone (Denmark)					3.520**	(0.002)	2.352**	(0.003)
India-Bhutan					2.205**	(0.0)	3.071**	(0.0)
Singapore-Brunei					2.747**	(0.0)	2.737**	(0.0)
Countries share a common currency	0.488**	(0.0)	0.169†	(0.053)				
ln Y_{it}^{k}	0.199**	(0.0)	0.340**	(0.0)	0.197**	(0.0)	0.328**	(0.0)
ln $Y_{world,t}^{k}$	1.474**	(0.0)	0.831**	(0.0)	1.480**	(0.0)	0.796	(0.755)
ln Distance	-1.558**	(0.0)	-0.863**	(0.0)	-1.562**	(0.0)	-0.833**	(0.0)
Pair belongs to a Regional trade Accord	0.193**	(0.0)	0.466**	(0.0)	0.191**	(0.0)	0.508**	(0.0)
Colonizer-colonized relationship	1.336**	(0.0)	0.326*	(0.025)	1.328**	(0.0)	0.671**	(0.0)
Country pair transitioning from colonialism	0.229	(0.37)	-0.542*	(0.013)	0.248	(0.34)	-0.406†	(0.061)
Countries were colonies of same country	0.610**	(0.0)	0.468**	(0.0)	0.617**	(0.0)	0.594**	(0.0)
Countries are contiguous	0.933**	(0.0)	0.510**	(0.0)	0.934**	(0.0)	0.537**	(0.0)
Shared common or official language	0.562**	(0.0)	0.404**	(0.0)	0.557**	(0.0)	0.399**	(0.0)
Number of observations	107786		155409		107786		155409	
Number of country pairs	13729		17817		13729		17817	

Table 15: Estimated coefficients for machinery trade

	(1) OLS		(2) PPML		(3) OLS		(4) PPML	
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
West African Economic & Monetary Union					2.005**	(0.0)	2.662**	(0.0)
Central African Economic & Monetary Union					3.528**	(0.0)	3.759**	(0.0)
Australia zone					2.546**	(0.0)	2.933**	(0.0)
Dollarized zone					-0.844**	(0.003)	-1.971**	(0.0)
Euro zone					-0.344**	(0.002)	-0.108	(0.143)
Krone zone (Denmark)					3.556**	(0.0)	5.133**	(0.0)
India-Bhutan					1.672**	(0.0)	3.357**	(0.0)
Singapore-Brunei					1.312**	(0.0)	1.239**	(0.0)
Countries share a common currency	0.165	(0.205)	-0.155*	(0.038)				
$\ln Y_{it}^{k}$	0.381**	(0.0)	0.766**	(0.0)	0.381**	(0.0)	0.771**	(0.0)
$\ln Y_{world,t}^{k}$	1.302**	(0.0)	0.466**	(0.002)	1.312**	(0.0)	0.464**	(0.002)
ln Distance	-1.563**	(0.0)	-0.639**	(0.0)	-1.578**	(0.0)	-0.651**	(0.0)
Pair belongs to a Regional trade Accord	0.349**	(0.0)	0.730**	(0.0)	0.340**	(0.0)	0.716**	(0.0)
Colonizer-colonized relationship	1.073**	(0.0)	0.053	(0.714)	1.079**	(0.0)	0.091	(0.575)
Country pair transitioning from colonialism	-0.132	(0.704)	0.115	(0.786)	-0.120	(0.733)	0.078	(0.857)
Countries were colonies of same country	0.652**	(0.0)	0.220	(0.115)	0.641**	(0.0)	0.221	(0.118)
Countries are contiguous	0.732**	(0.0)	0.478**	(0.0)	0.718**	(0.0)	0.462**	(0.0)
Shared common or official language	0.683**	(0.0)	0.227**	(0.001)	0.670**	(0.0)	0.226**	(0.001)
Number of observations	119268		151572		119268		151572	
Number of country pairs	14050		17045		14050		17045	

Table 16: Estimated coefficients for textile trade

	(1) OLS		(2) PPML		(3) OLS		(4) PPML	
	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value	$\hat{\beta}_x$	p-value
East Caribbean Currency Union					0.209	(0.727)	-0.588	(0.476)
West African Economic & Monetary Union					1.171^{**}	(0.0)	1.453^{**}	(0.007)
Central African Economic & Monetary Union					2.534^{**}	(0.002)	2.216^{**}	(0.002)
Australia zone					2.084^{**}	(0.0)	2.122^{**}	(0.001)
Dollarized zone					0.308	(0.363)	0.364	(0.652)
Euro zone					-0.481^{**}	(0.0)	$-0.147^{\dagger}$	(0.059)
Krone zone (Denmark)					4.199^{**}	(0.0)	4.773^{**}	(0.0)
India-Bhutan					-2.125^{**}	(0.0)	-0.019	(0.975)
Singapore-Brunei					3.100^{**}	(0.0)	3.150^{**}	(0.0)
Countries share a common currency	0.002	(0.988)	$-0.140^{\dagger}$	(0.076)				
ln Y_{it}^{k}	0.403^{**}	(0.0)	0.715^{**}	(0.0)	0.400^{**}	(0.0)	0.735^{**}	(0.0)
ln $Y_{world,t}^{k}$	1.913^{**}	(0.0)	0.809	(0.0)	1.928^{**}	(0.0)	0.762	(0.0)
ln Distance	-1.524^{**}	(0.0)	-0.891^{**}	(0.0)	-1.533^{**}	(0.0)	-0.954^{**}	(0.0)
Pair belongs to a Regional trade Accord	0.582^{**}	(0.0)	0.627^{**}	(0.0)	0.581^{**}	(0.0)	0.542^{**}	(0.0)
Colonizer-colonized relationship	1.263^{**}	(0.0)	0.283	(0.247)	1.247^{**}	(0.0)	0.358	(0.192)
Country pair transitioning from colonialism	$0.513^{\dagger}$	(0.094)	-0.820^{**}	(0.005)	$0.534^{\dagger}$	(0.085)	-0.803^{**}	(0.006)
Countries were colonies of same country	0.540^{**}	(0.0)	-0.053	(0.711)	0.535^{**}	(0.0)	-0.068	(0.635)
Countries are contiguous	0.721^{**}	(0.0)	0.357^{**}	(0.001)	0.748^{**}	(0.0)	0.326^{**}	(0.002)
Shared common or official language	0.693^{**}	(0.0)	0.431^{**}	(0.0)	0.680^{**}	(0.0)	0.429^{**}	(0.0)
Number of observations	112134		155076		112134		155076	
Number of country pairs	14261		17717		14261		17717	

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