The Empirics of Exchange Rate Regimes and Trade[†]

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Abstract

This paper examines the impact of fixed exchange rate regimes on bilateral trade while differentiating the effects of “deeds” versus “words”. Our empirical findings, based on an extended database for de jure and de facto exchange rate classifications suggest that fixed exchange rate regimes in the form of currency unions and conventional pegged arrangements promote trade, through channels in addition to reduced exchange rate volatility. The trade creating effect is broadly similar for both de jure and de facto direct pegs, and currency unions, suggesting that exchange rate stability through policy actions as well as commitments matters for trade. In addition, we find evidence that (i) direct pegs have a larger impact when de jure and de facto regimes are aligned; (ii) the currency union effect is dependent on the geographical proximity of trading partners, with member countries closer to each other achieving larger trade gains; (iii) exchange rate stability reduces trade instability; and (iv) the trade generating effects of more stable regimes are heterogeneous across country groups.

JEL Classification Numbers: F1, F3, F4

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Contents Page

Tables

Figures

I. INTRODUCTION

The choice of exchange rate regime and its macroeconomic implications—a well-debated subject since the collapse of the Bretton-Woods system in the early 1970s—gained renewed interest with a series of financial crisis in the late 1990s. The exchange rate regimes adopted by the hardest-hit countries were widely believed to have played a role in triggering the crisis, which led to a greater scrutiny of exchange rate policy choices and their impact on the macroeconomy by both researchers and policy makers.

Most of the ensuing research focused on the influence of exchange rate regimes on economic growth, inflation, and macroeconomic stability. A notable exception to this is the seminal work of Rose (2000), which investigates the effect of currency unions on trade, and finds that two countries having a common currency tend to trade about three times as much as they would otherwise. Frankel and Rose (2002) further show that the growth enhancing benefits of currency unions occur through increased trade only and not through other channels (such as reduced inflation). These findings have generated immense interest and controversy, and numerous studies have followed which, in general, find a smaller magnitude of the currency union effect on trade than estimated by Rose and his co-authors.[1]

In two recent studies, Klein and Shambaugh (2006), and Adam and Cobham (2007) go beyond currency unions, and investigate the impact of other possible exchange rate regimes on bilateral trade flows. They find that exchange rate regimes with lower uncertainty and transaction costs—namely, pegs and currency unions—are significantly more pro-trade than flexible exchange rate regimes. Both studies use de facto exchange rate classifications that categorize regimes according to the actual behavior of exchange rates. Klein and Shambaugh use the de facto exchange rate regime classification developed by Shambaugh (2004), while Adam and Cobham use the regime classification of Reinhart and Rogoff (2004) for their analysis.

The use of de facto classification in these studies is based on the premise that it offers an improved characterization of the exchange rate regime in place, as deeds presumably matter more than words. In fact, since pervasive differences were highlighted by earlier research between the regimes officially announced by countries, and those followed in practice, the use of the former in empirical exchange rate analysis has been significantly reduced. However, to the extent that central banks' commitments affect expectations, the de jure exchange rate arrangement—which captures these commitments—may also affect economic outcomes. In this respect, both actual actions as well as policy announcements could matter, with each reflecting a different aspect of exchange rate stability. Their joint analysis could therefore provide insights into the importance of "deeds versus words" for cross-border trade activity.

In this paper, we therefore empirically revisit the relationship between trade and exchange rate regimes using both the de jure and de facto exchange rate regime classifications of the IMF. A notable advantage of using the IMF's de facto classification is that it is the only available

[1] See, for example, Nitsch (2002), Levy-Yeyati (2003), and Tsangarides et al. (2008).

classification that assesses central bank behavior in addition to supplementary indicators such as exchange rate movements. To this end, we employ: (i) a novel dataset on exchange rate regime classifications compiled by Anderson (2008), which harmonizes the chronological coverage of IMF's de jure and de facto classifications by extending the former up to 2006, and the latter backwards up to 1972, providing comprehensive data coverage for both classifications; and (ii) a recent bilateral trade data set comprising 159 countries over the period 1972-2006, which importantly includes the formation of the European Monetary Union (EMU).[2]

We begin by investigating the impact of currency unions, conventional pegs, and the possible indirect pegs created as a result of pegged arrangements with anchor countries, on bilateral trade. We then explore whether the impact of these regimes is conditional on the geographical distance between trading partners.[3] This is important since if the lower transaction cost benefits associated with less flexible arrangements, particularly, currency unions, diminish as other trading costs increase, then modeling a linear relationship may not be sufficient to extrapolate the effects of these regimes on bilateral trade. Further, unlike previous studies, which focus only on trade volumes, we also examine if fixed exchange rate regimes affect trade stability. If—through lower uncertainty or output-price volatility—fixed exchange rate regimes promote predictability and consistency in cross-border trade, then bilateral trade may be boosted further in the long-run. Importantly, we investigate these issues while addressing some of the concerns raised in previous literature about the proper econometric specification of gravity models, potential endogeneity of the explanatory variables, and heterogeneity of the trade-generating effect across sub-samples.

Our findings suggest that, on average, both de jure and de facto fixed exchange rate regimes promote bilateral trade, through channels in addition to reduced exchange rate volatility. Specifically, we find that: (i) countries belonging to a currency union trade substantially more with each other than with comparable countries that do not share a currency. This is true for both industrialized and non-industrialized dyads, although the effect is larger for the latter; (ii) both de jure and de facto pegs create trade—particularly, between industrialized and non-industrialized pairs—in the same order of magnitude as currency unions; (iii) the effect of direct pegs is magnified when deeds are aligned with words; (iv) indirect pegs lower trade, suggesting possible trade diversion effects, largely driven by the non-industrialized dyads; (v) the effect of currency unions is dependent on the geographical proximity of trading partners, with member countries closer to each other benefiting more; and (iv) exchange rate stability promotes trade stability.

In what follows, Section II provides a brief background to this research. Section III outlines the empirical strategy, and discusses relevant estimation issues. Section IV describes the data. Section V presents the estimation results and the sensitivity analysis. Section VI concludes.

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[2] In the context of currency unions, Baldwin (2006) notes that including the post-Euro period may have important implications for the estimated results.

[3] The choice of exchange rate regime may itself depend on the geographical distance between two countries. For example, Tenreyro (2001) shows that distance is an important determinant of currency union participation. However, to the extent that certain exchange rate regimes, particularly, those in developing countries, are motivated by political or economic stabilization concerns, it is important to examine if their benefits depend on other types of trading costs, notably, transportation costs.

II. BACKGROUND

An extensive body of literature in the late 1970s through the 1980s investigates the effect of more stable regimes exchange rate regimes on trade by analyzing the link between exchange rate volatility and trade.[4] These studies offer mixed evidence, not providing strong support to the argument that a reduction in exchange rate volatility increases trade. For example, Clark and Haulk (1972) and Hooper and Kohlhagen (1978) use different measures of nominal exchange rate volatility, but find no evidence that exchange rate risks affect trade volumes. However, Cushman (1983), Kenen and Rodrik (1986), and Thursby and Thursby (1987) show that real exchange rate variability affects trade flows negatively.

Studies in the following decade are more consistent in their findings and report negative, albeit small, effects of exchange rate volatility on trade. For example, Frankel and Wei (1993) and Frankel (1997) find some evidence that bilateral exchange rate variability has a small, negative effect on trade. Similarly, applying a gravity model to panel data from western Europe, Dell'Ariccia (1999) finds that exchange rate uncertainty has a significant negative effect on international trade, which is robust to different measures of exchange rate uncertainty, and remains even after controlling for the potential simultaneity bias.

Using a somewhat different approach, Rose (2000) models the effect of fixed exchange rate regimes on trade by explicitly including a binary variable for currency unions in the gravity model. His results show a large trade generating effect of currency unions: two countries sharing a currency tend to trade roughly three times as much as they would otherwise. Rose and van Wincoop (2001), Frankel and Rose (2002), and Glick and Rose (2002) confirm this result, and show that it is robust to various specifications and estimation techniques. Other studies have challenged the magnitude of the currency union effect reported in these studies largely on methodological grounds, but in general agree on the existence of a strong positive effect.[5]

In recent years, some studies go beyond currency unions, and investigate the impact of other possible exchange rate regimes on bilateral trade using the gravity model framework.[6] Notably, Klein and Shambaugh (KS, 2006) estimate the impact of currency unions, direct pegs, and indirect pegs on bilateral trade flows, while controlling for exchange rate volatility. They report large and significant bilateral trade gains from de facto direct pegs and currency unions, but find no significant impact of indirect pegs. Adam and Cobham (2007) use a fine classification, encompassing 27 exchange rate regimes, and show that regimes with lower exchange rate

[4] See IMF (1984) for a detailed survey of early literature on exchange rate volatility and trade.

[5] Rose and Stanley (2005) conduct a meta-analysis from 34 studies, and conclude that currency unions increase trade in the range of 30-90 percent.

[6] Earlier attempts include Brada and Mendez (1988) and Pozo (1992). The former analyze the impact of exchange rate regimes on bilateral trade flows for 30 countries using cross-sectional data from the mid-1970s. They find that trade flows between countries with floating exchange rates are higher than those with a fixed exchange rate regime. The latter examines the effect of exchange rate regimes and volatility on the exports from United Kingdom to the United States in 1900-40. Her findings suggest that higher real exchange rate volatility reduces trade volumes, but there is no significant effect of different regimes on trade.

uncertainty and transactions costs are significantly more pro-trade than the floating regime. They argue that the direct and indirect trade-creating effects of more stable regimes tend to outweigh any trade-diverting effects such that the net effect is positive, and of the same magnitude as of currency unions or regional free trade agreements.

III. EMPIRICAL STRATEGY

A. Analytical Framework

In line with recent literature, we employ the gravity model of bilateral trade flows to investigate the effect of exchange rate regimes on trade. The gravity model, a popular tool in empirical trade analysis, represents trade between two countries as a function of their respective economic sizes and obstacles to trade such as the distance between them. The initial criticism that these models lack a proper theoretical foundation has been addressed by several studies that use different approaches to establish a theoretical justification for these models (see, for example, Anderson, 1979; Bergstrand, 1985; and Deardorff, 1998). In its simple form, the new trade theory model of Helpman and Krugman (1985) can be used to derive the gravity equation as follows:[7]

$$X_{ij} = Y_i Y_j \left(\frac{p_i T_{ij}}{P_j} \right)^{1-\sigma}, \tag{1}$$

where X_{ij} represents the exports from country i to j, Y is total domestic output, p_i are marked up prices of domestic producers (net of any transport costs) in country i, P_j is the overall price index in country j, T_{ij} are the iceberg trading costs (such that $X'_T < 0$), and σ is the elasticity of substitution between products ($\sigma > 1$).[8]

Traditionally, T includes transportation costs that are proxied by geographical attributes (such as bilateral distance, access to sea, and contiguity). In recent years, several other factors that may affect trade, for example, common language, historical ties, free trade agreements, tariffs, and non-tariff barriers have also been included. However, to the extent that the introduction of a fixed exchange rate regime reduces conversion costs between countries, and lowers exchange rate volatility and uncertainty, trading costs could also depend on the exchange rate regime in place, which may have important consequences for bilateral trade flows.

Benchmark specification

To examine the trade effects of exchange rate regimes which lower transaction costs, we augment the traditional gravity equation, and include variables for currency unions, direct pegs, and indirect pegs, and estimate the following benchmark specification:

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[7] See Feenstra (2004) for a detailed derivation.

[8] Iceberg trading costs imply $p_j = T_{ij}\, p_i$ (where $T_{ij} \geq 1$), indicating that T_{ij} units of a product must be shipped to country j for one unit to arrive.

$$\log(X_{ijt}) = \beta_0 + \sum_{k=1}^{N} \beta_k Z_{ijt} + \gamma CU_{ijt} + \delta DirPeg_{ijt} + \varepsilon IndPeg_{ijt} + \zeta Vol_{ijt} + \lambda_t + u_{ijt}\,, \qquad (2)$$

where X_{ijt} denotes bilateral trade between countries i and j in year t; Z is a vector consisting of traditional time-varying and in-variant trade determinants;[9] *CU* is binary variable that is unity if i and j share the same currency; *DirPeg* is also a binary variable that is unity if i's exchange rate is pegged to j, or vice versa (but i and j are not members of the same currency union); *IndPeg*—defined in a somewhat similar manner to KS (2006)—is a binary variable that takes the value of 1 if i is indirectly related to j through its peg with an anchor country;[10] *Vol* refers to real exchange rate volatility defined over a specific horizon; λ_t are the year-specific effects indicating common shocks across countries; and u_{ij} is the error term, assumed to be independently and normally distributed ($u_{ij} \sim N(0,\sigma)$).

To construct direct and indirect pegs, we need information on anchor countries. Our list of anchors includes major currencies as well as regionally important currencies (see Table B4). We focus on strict (or explicit) anchors, whereby countries serving as anchors of monetary policy or multiple anchors (basket pegs) are not included. Further, since the depth or level of the indirect peg relation between a trading pair may imply a different impact on trade, we use two alternative coding schemes for indirect pegs. In the first scheme, we include the shortest indirect linkage where a dyad pegged to the same base is considered as having an indirect peg. In the second, we include longer indirect linkages, such as those between an anchor country and another country pegged to the anchor country; or, between two countries which are pegged to different base countries, but their base countries are pegged to the same anchor country.[11]

Overall, the three exchange rate regime categories included in our estimation—currency unions, direct and indirect pegs—are mutually exclusive such that at a point in time, each country pair is coded as one of the three. The reason for including exchange rate volatility in the benchmark specification is to examine if more stable exchange rate regimes improve trade through multiple channels (such as a reduction in transaction costs, increased transparency, and competition). To construct the n-horizon real exchange rate volatility, *Vol*, we follow Ghosh, Gulde and Wolf (GGW, 2003), and proceed in two steps. First, for each month in a given year, we take the absolute value of the percentage change in the exchange rate over the previous n months. In the

[9] The time-variant variables include: the (logs of) product of real gross domestic product, and the real GDP per capita of the trading partners; and a dummy variable equal to unity if the two countries share a free trade agreement. The time invariant variables include: the (log of) product of land areas of the trading partners, the (log of) distance between them, and whether the countries are landlocked or island, and binary variables which equal unity if the countries share colonial ties, language, and border.

[10] By this definition, two countries (B and C) that are pegged to the same anchor (A) would be classified as having an indirect peg with each other. Similarly, if a country, D, is pegged to B, then it would also have indirect pegs with A and C, and so forth. See Figure B1 for a diagrammatic illustration of possible indirect peg relations.

[11] Specifically, the first definition of indirect peg includes links between pairs defined by relation=2 (which is equivalent to the "sibling" relationship of KS) in Figure B1. The second definition includes indirect relations=2, 3, 4 and 5.

next step, we take the average of the absolute values over *n* months to obtain a measure corresponding to that particular year, given by:

$$Vol_t = \sum_{p=1}^{n} \frac{\left| R_{t+p-1} - R_{n+p-1} \right|}{n},$$

where *R* is the bilateral real exchange rate between countries *i* and *j*. We define *Vol* over two horizons—12 and 36 months—to represent short- and long-run volatility, respectively. In addition, we construct two other measures of volatility to verify the robustness of our results to different definitions of the variable. These are:

$$Vol2_t = SD[r_{t+p-1} - r_{t+p-2}], \text{ and}$$

$$Vol3_t = \frac{\overline{R}^2}{\left[\frac{1}{n} \sum_{p=1}^{n} (R_{t+p-1} - \overline{R})^2 \right] + \overline{R}^2},$$

where *r* is the natural log of bilateral real exchange rate between countries *i* and *j*; and $\overline{R}$ is the average bilateral real exchange rate over the given period. *Vol*2 defines volatility as the standard deviation of the first difference of (logs of) the real exchange rate. The first difference is computed over one month (with end-of-month data), while the standard deviation is calculated over a one- and three-year period to measure short- and long-run volatility, respectively. *Vol*3 represents a linear transformation of the coefficient of variation of real exchange rates, and is also computed for the short- and long-runs.

Does distance matter?

The trading costs, *T*, that enter equation (1) include different types of transportation and transaction costs. If *T* is a multiplicative function of these trading costs, then the trade generating effect of lower currency conversion costs would depend on other trading costs as well as on the elasticity of substitution between products. Specifically, if

$$T_{ij} = \prod_{m=1}^{M} (\tau_{ij}^{m}), \qquad \tau_{ij}^{m} = 1 \text{ if } i{=}j, \quad \tau_{ij}^{m} \geq 1 \text{ otherwise,} \tag{3}$$

then the trade gains from lower transaction costs are proportional to:

$$\frac{\partial X_{ij}}{\partial \tau_{ij}^{M}} \sim \left(\prod_{m=1}^{M-1} \tau_{ij}^{m} \right)^{1-\sigma} \left(\tau_{ij}^{M} \right)^{-\sigma}. \tag{4}$$

Thus, if goods are substitutes, a reduction in currency conversion has a larger effect when other trading costs are smaller. This is because, as noted by Tenreyro (2001), lower transaction costs would save on trading costs—which are particularly important when trading volumes are large—and have a higher impact on trade.

To include the differential impact of exchange rate regimes based on other trading costs, we modify equation (2) to include an interaction term of currency unions and direct pegs with geographical distance:[12]

$$\log(X_{ijt}) = \beta_0 + \sum_{k=1}^{N} \beta_k Z_{ijt} + \gamma CU_{ijt} + \delta DirPeg_{ijt} + \varepsilon IndPeg_{ijt} + \zeta Vol_{ijt} + \phi(CU_{ijt} \times Dist_{ij}) + \theta(DirPeg_{ijt} \times Dist_{ij}) + \lambda_t + u_{ijt}. \tag{5}$$

If the effect of exchange rate regime is indeed determined by distance such that dyads located closer to each other benefit more, then $\hat{\gamma}$ and $\hat{\delta}$ are expected to be positive, whereas $\hat{\phi}$ and $\hat{\theta}$ are expected to be negative. The (conditional) marginal effects of currency unions and direct pegs would be computed as $(\hat{\gamma} + \hat{\phi} Dist_{ij})$ and $(\hat{\delta} + \hat{\theta} Dist_{ij})$, respectively.[13]

Trade stability

A pertinent issue in the context of exchange rate regimes and trade is if less volatile exchange rate arrangements reduce trade instability. To investigate this question, we follow Rose (2005) and estimate an equation similar to the benchmark specification but with the coefficient of variation of real bilateral trade as the dependent variable.[14] We calculate the dependent variable over a (non-overlapping) five-year period, which gives us five observations per pair for the sample. All the explanatory variables are averaged over the corresponding time period, and the following equation is estimated:

$$CV(X_{ijt}) = \beta_0 + \sum_{k=1}^{N} \beta_k \overline{Z}_{ijt} + \gamma \overline{CU_{ijt}} + \delta \overline{DirPeg_{ijt}} + \varepsilon \overline{IndPeg_{ijt}} + \zeta \overline{Vol_{ijt}} + \lambda_t + u_{it}, \tag{6}$$

where $CV(X_{ijt})$ is the ratio of the standard deviation to the mean of bilateral trade flows between countries i and j over time-period t, and bars over variables indicate period averages.[15] If currency unions and pegs improve trade stability then the estimated coefficients for γ, δ, and ε would be negative (the direct effect) and ζ would be positive (the indirect effect), implying lower trade volatility.

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[12] Other trading costs could include transportation and information costs. Limão and Venables (2001) show that in addition to geographical distance, transport costs depend significantly on trading pair's infrastructure. Since we do not have adequate time series data available on transport or communication infrastructure for the sample of countries considered, we use geographical distance as a proxy for these costs.

[13] The respective standard errors are as follows:

$\hat{\sigma}_{X,CU} = \sqrt{\mathrm{var}(\hat{\gamma}) + [\log(Dist_{ij})]^2 \mathrm{var}(\hat{\varphi}) + 2\log(Dist_{ij})\mathrm{cov}(\hat{\gamma}, \hat{\varphi})}$, and $\hat{\sigma}_{X,DirPeg} = \sqrt{\mathrm{var}(\hat{\delta}) + [\log(Dist_{ij})]^2 \mathrm{var}(\hat{\theta}) + 2\log(Dist_{ij})\mathrm{cov}(\hat{\delta}, \hat{\theta})}$.

[14] Admittedly, the gravity equation is a model of trade flows and not trade volatility. However, the estimated equation fits the data reasonably well.

[15] The variables on the right hand side are defined similar to equation (2).

B. Estimation Issues

Estimation of the gravity model raises several methodological issues that have been discussed extensively in the literature (see, for example, Baldwin (2006)). To the extent that these concerns relate to the analysis presented in this paper, we discuss our attempts to address them in the estimation, and through the sensitivity analysis of the obtained results.

An important issue in the estimation of our benchmark specification is the omitted variables bias originating from the correlation of any pro-trade omitted variables with the right-hand side variables (or the lack of independence of Z_k and the errors) in the gravity equation. The pooled Ordinary Least Squares (OLS) approach essentially assumes that there is no unobserved individual heterogeneity across countries. However, if such heterogeneity exists, and the error term is correlated with Z_k, then the OLS estimator is likely to be biased and inconsistent.

Research following Rose (2000) attempts to control for this bias by introducing country-specific idiosyncrasies in the gravity model—both for cross-sectional and panel estimations. In cross-section analysis, country fixed-effects (CFE) are used to account for Anderson and van Wincoop's (2003) "multilateral resistance" factor, according to which trade between two countries does not only depend on the characteristics of the countries, but also on the barriers between them and the rest of the world. However, given that there is a time-series element to the potential bias that is not eliminated with this procedure; Anderson and van Wincoop (2004) propose that separate country fixed-effects should be included for each year (CYFE) to take into account changes in multilateral resistance over time.

Glick and Rose (2002) argue that including CFE or CYFE may still not resolve the omitted variables problem. This is because the unobserved variables could be correlated with the bilateral characteristics of the dyads (such as the propensity to opt for a particular exchange rate regime) and the trade between them, which may bias the CFE/CYFE estimates. They therefore propose using the panel data fixed-effects estimator that adds country-pair specific effects (CPFE) to the gravity equation, thereby controlling for any strong bilateral likelihood to trade. However, the CPFE approach does not provide coefficient estimates for the time invariant variables as they are captured by the CPFE. This may have implications for estimating equation (2) since, as noted by KS (2006), any country pair that has had the same exchange rate regime (currency union or direct peg) during the sample period will not yield information in the estimated impact of the regime on bilateral trade.

Another potential source of endogeneity in our model stems from the fact that the choice of exchange rate regime may itself depend on trade links between partner countries. If this holds true then some of the large trade-creating effects of these regimes may actually be a reflection of reverse causality. Most of the literature has ignored the endogeneity concerns because of the difficulty in finding plausible instruments, but exceptions include Alesina, Barro, and Tenreyro

(2002), Barro and Tenreyro (2007), and Frankel (2008), who find that currency unions have a significant positive effect on bilateral trade even after controlling for endogeneity.[16]

In our analysis, we attempt to address both the endogeneity concerns (resulting from the omitted variable bias and reverse causality) as well as the estimation of time invariant (or with little variation) regressors using the Hausman and Taylor (HT, 1981) estimation technique. The HT estimator—based on the instrumental variable approach—yields consistent and efficient estimates in the presence of correlation between some explanatory variables and the error term.[17] To construct instruments, the HT method exploits various dimensions of the data, and suggests using the following variables: (i) deviations of the time variant exogenous variables from individual means; (ii) deviations of the time variant endogenous variables from individual means; (iii) mean (across time) of the time variant exogenous variables; and (iv) time-invariant exogenous variables.[18] The regressors that constitute the endogenous set of variables can be determined by a Hausman test, which is based on the comparison of the HT estimator with the fixed effects estimator.

The two most obvious advantages of the HT estimator are the construction of valid instruments from within the model, and using the means of exogenous time variant variables as instruments to estimate the effect of time-invariant variables. However, despite its useful features, the HT method has been less widely applied. Egger (2002), and Egger and Pfaffermayr (2003, 2004) argue that the HT method is superior to the traditional OLS, random and fixed effects methods in the context of bilateral trade models. Carrére (2008) applies it to study the endogenous link between regional trade agreements and bilateral trade flows, and finds that the HT estimator is more appropriate to the fixed effects estimator. Serlenga and Shin (2007) use the HT method to examine intra-EU trade during 1960-2001, and also find evidence that the HT method is more suitable than the fixed and random effect methods.

Finally, there are several issues relating to model misspecification that have been raised and discussed extensively in earlier literature. These include: (i) the aggregation of exports and imports as the dependent variable; (ii) possible nonlinear effects entering the gravity equation; (iii) the sample selection bias; and, (iv) the inclusion of zero-trade flows. We attempt to address all these issues in an extensive set of robustness tests, discussed in Section V.

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[16] To address endogeneity issues, Alesina, Barro, and Tenreyro (2002) and Barro and Tenreyro (2007) exploit client country decisions to peg to an anchor country to construct instruments. Frankel (2008) attempts a "natural experiment" where he examines the effect of the French franc's conversion to the euro in 1999 on the bilateral trade of CFA members with other European countries.

[17] Thus, instead of imposing an "all" (as in the fixed effects method) or "nothing" (as in the random effects method) correlation among the omitted and explanatory variables, the HT method allows for some regressors to be correlated with the omitted variables. Baltagi (2001) proposes to check the viability of the HT specification when testing for the validity of the fixed and random effects.

[18] Specifically, the endogenous time varying variables are instrumented by their deviation from individual means, and the endogenous time invariant variables are instrumented by the deviation of the exogenous time varying variables from individual means. Identification requires the number of exogenous time varying variables to be at least as large as the number of endogenous time invariant variables in the model.

IV. DESCRIPTION OF DATA

A. Exchange Rate Regime Classification

An important issue in the empirical study of exchange rate regimes is that of regime classification. Early literature used the de jure classification—the regime declared by national authorities, and published in the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions (AREAR)*. However, since the work of Obstfeld and Rogoff (1995) and Calvo and Reinhart (2002) highlighted pervasive differences in the de jure and de facto currency regimes through the "mirage of fixed rates" and the "fear of floating", respectively, the use of de jure classification in empirical exchange rate analysis has been significantly reduced.[19] Thereafter, "de facto" classifications that seek to categorize regimes according to the exchange rate or central bank behavior have been developed. The most commonly used are those of GGW (2003), Levi-Yeyati and Sturzenegger (LYS, 2003), Reinhart and Rogoff (RR, 2004), and Shambaugh (JS, 2004), which are based on measurable outcomes of exchange rate behavior.[20]

Nevertheless, any attempt to examine the differences in macroeconomic implications of the de facto regime vis-à-vis the de jure regime using the above classifications is beset with two problems. First, the sources and data coverage underlying the above classifications is different from the de jure classification, making it difficult to judge whether any difference in findings reflects substantive variation across the two classifications or simply differences in the sample and sources. Second, there is little agreement among the various de facto classifications, making it hard to know whether results are driven by genuine differences in performance across regimes or simply idiosyncrasies in the classification schemes.

To address the problems discussed above, we define the exchange rate arrangement between trading partners using the IMF's de jure and de facto classifications. This enables us to capture the stated *and* implemented policies of the central bank using data from a common source, with similar sample coverage. The IMF's de facto classification scheme—adopted since 1999—combines available information on central bank's policy framework with the actual exchange rate and international reserves movements to form a judgment about the exchange rate regime in place. In this respect, it is the only de facto classification that takes into account central bank behavior. The necessary information is compiled from different primary (for example, IMF's surveillance and technical assistance reports) and secondary (such as reports of the press and other multinational organizations) sources.[21]

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[19] The "mirage of fixed rates" refers to the fact that most countries that claim to peg their exchange rates do not firmly peg in practice. The "fear of floating" reflects the fact that most countries that claim to have a floating currency regime, intervene heavily in the market to prevent exchange rate movement.

[20] See Rogoff et al. (2004) and Shambaugh (2004) for a detailed discussion on the various regime classifications.

[21] In 2000, the IMF started publishing the de facto classification in the *AREAER* instead of the de jure classification. Its de facto classification has been extended backwards for the period 1990-2000 by Bubula and Ötker-Robe (2002). Recently, Anderson (2008) extended it further backwards, and also updated the de jure classification to 2006.

The IMF's de facto classification is also less idiosyncratic than the others. This means that—on average—for each (country-year) observation, the other de facto classifications agree more with the Fund's classification than with each other. Figure 1 compares the IMF's de facto and de jure classifications with the classifications of LYS and RR using a composite measure of similarity when regimes are ggrouped as fixed, intermediate and floating.[22] The similarity index—which is a weighted average of the consensus between the classifications across the three regimes—takes a value in the range from 0 to 1, with a value of 1 indicating perfect similarity of the classifications.[23] The comparison reveals that the IMF's classification is overall more similar to the other two—it receives an average score of about 0.75 if the de jure classification is included in the comparator category (and of about 0.72 if it is not), while LYS and RR receive overall scores of 0.66, and 0.58, respectively.[24]

Further grouping the regimes into a binary variable reflecting pegged versus non-pegged regimes, and comparing IMF's de facto classification with those of JS, LYS, and RR, we note that it is the closest to the JS classification and the least similar to LYS. About 87 percent of the observations in the IMF de facto classification are coded (as pegs or non-pegs) in the same way as in the JS classification, and the overall correlation between the two series is 0.76.[25]

Table 1. Agreement of different exchange rate regimes, 1972-2006

	Percentage matching				
	De jure	De facto	JS	LYS	RR
De jure	100.00				
De facto	84.51	100.00			
JS	80.80	87.79	100.00		
LYS	69.76	59.64	62.90	100.00	
RR	77.55	77.67	81.95	54.31	100.00
	Correlation				
	De jure	De facto	JS	LYS	RR
De jure	1.00				
De facto	0.71	1.00			
JS	0.58	0.76	1.00		
LYS	0.49	0.57	0.64	1.00	
RR	0.54	0.66	0.64	0.54	1.00

Source: Authors' calculations based on Anderson (2008).

[1] JS, LYS and RR refer to the de facto classifications of Shambaugh (2004), Levy-Yeyati and Struzenegger (2003), and Reinhart and Rogoff (2004).

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[22] See Appendix A for a discussion on the aggregation of the regimes, and their distribution across country groups.

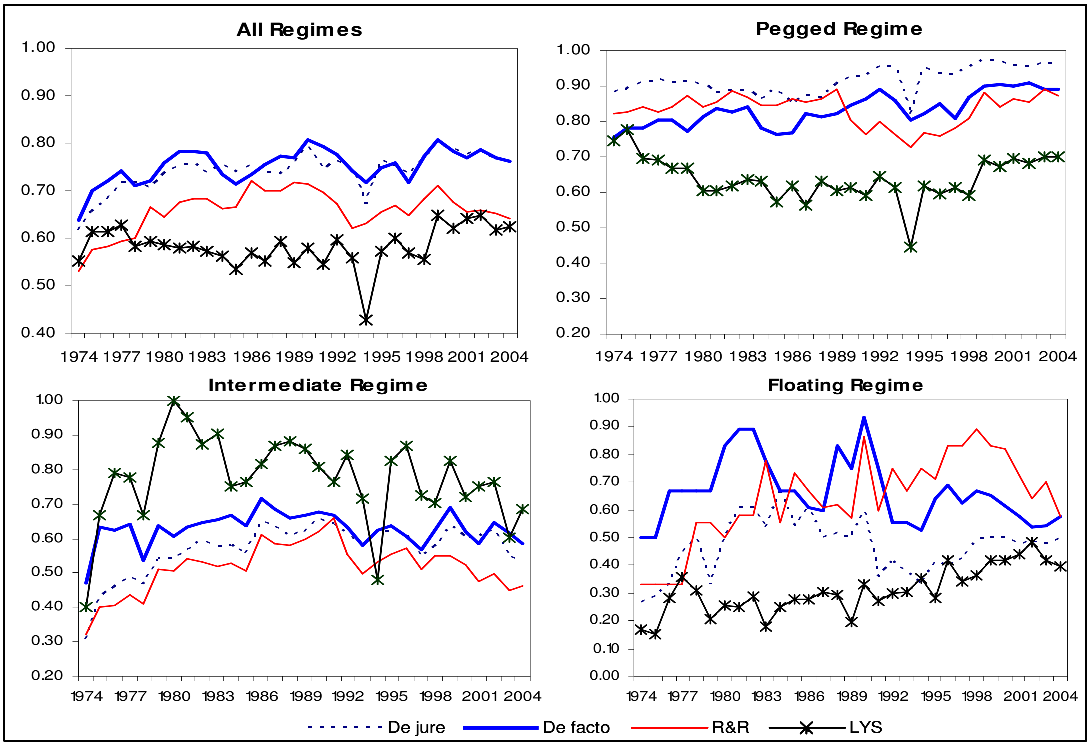
[23] To construct the index based on DJ, DF, RR and LYS (DF, RR, and LYS), each classification is assigned a value of 1 if it agrees with any of the other classifications. Hence, for every classification, a country-year observation receives a score of 1/3 (1/2) for each other classification that agrees with it. The overall index is constructed as the weighted sum of the scores for the three regime groups, with the weights equal to the proportion of pegs, intermediate, and floats in the particular classification.

[24] The JS classification is not included in the comparison as it is available as a binary variable (pegs/non-pegs).

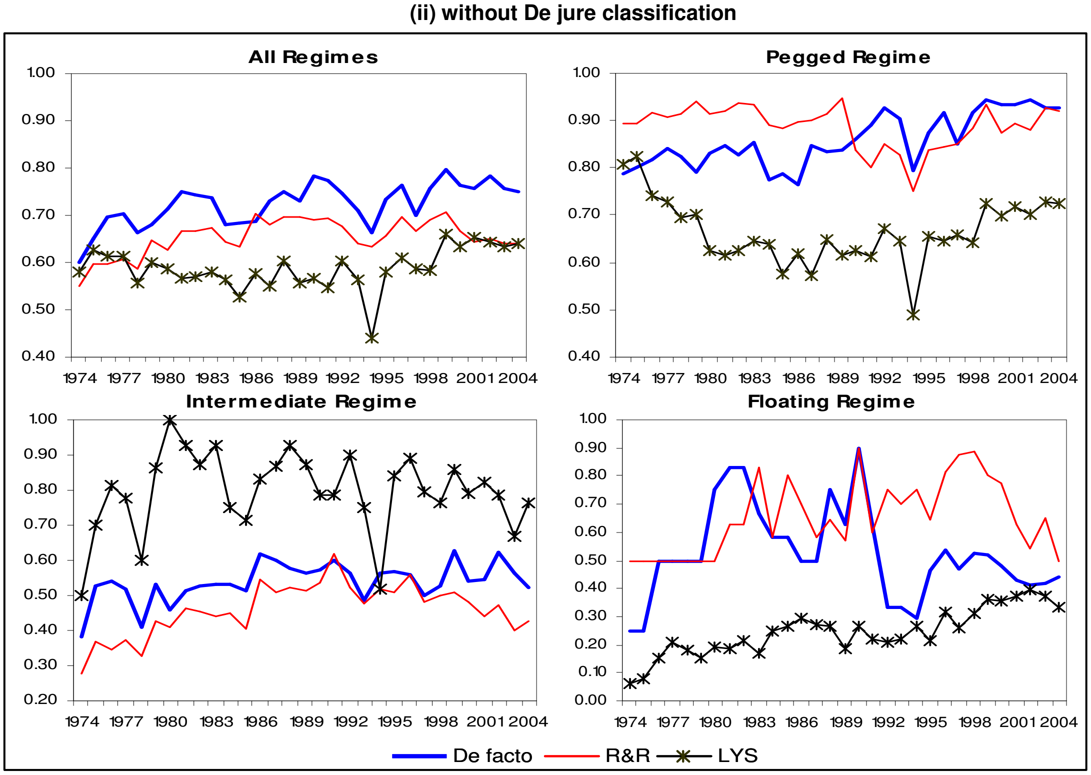
[25] For binary coding, RR's classification with codes 1-4 (inclusive) is considered as pegs; LYS's three-way classification index with code equal to 3 is treated as a peg; and JS's binary classification (jspeg) is used.

Figure 1. Similarity Index across different exchange rate regime classifications*

(i) with De jure classification



(ii) without De jure classification



Source: Authors' calculations based on Anderson (2008).
*Sample period is restricted to 1974-2004 corresponding to data availability for the other classifications.

B. Data and Summary Statistics

The data used in the empirical analysis have been compiled from multiple sources.[26] Annual bilateral trade data has been obtained from IMF's *Direction of Trade Statistics*. Information on the IMF's de jure and de facto exchange rate classifications and the anchor currencies, used to construct the binary variables for currency unions and direct pegs, has been compiled from Anderson (2008). The dummy variable for indirect pegs is defined using an algorithm to associate bilateral exchange rate relations with anchor currencies, along the lines discussed in Section III.A.[27]

Data on real GDP (in 2000 US dollars), real GDP per capita (in 2000 US dollars), population and geographical size have been taken from the World Bank's *World Development Indicators 2007*. The source of information on free trade agreements is the Regional Trade Agreements database of the World Trade Organization. The various measures of distance have been compiled from the *Centre D'Etudes Prospectives et D'Informations Internationales*, while colonial ties, common border and language are compiled from the CIA *World Factbook 2004* and Rose (2000).

We estimate the benchmark and augmented gravity specifications for a range of samples including dyads belonging to different or similar income groups, but for brevity report the results of four samples (world, industrial-industrial (Ind-Ind), industrial-nonindustrial (Ind-Nind), and nonindustrial-nonindustrial (Nind-Nind)). The first sample covers all countries for which the required data are available; the second comprises those observations where both trading partners belong to industrial countries; the third includes those dyads where one partner is an industrial country and the other is a nonindustrial country; and the fourth covers the pairs where both countries are nonindustrial.

Table 2 presents the distribution of currency unions, direct pegs, and indirect pegs in the bilateral dataset used for estimation purposes. The dataset covers 159 countries over the period 1972-2006, yielding 10,894 individual country pairs (rather than 159×158/2=12,561 because of missing observations), and 177,270 observations. Over half of the observations in the sample belong to Nind-Nind dyads but interestingly they account for only 7 percent of world trade conducted in the sample period; while the Ind-Ind pairs constitute about 5 percent of the observations, and represent over 50 percent of world trade. Almost 40 percent of the observations are Ind-Nind pairs that make up 40 percent of world trade.

In the full sample, the number of observations coded as de facto pegs is higher than de jure pegs. Of the direct de jure and de facto pegs, about 90 percent of the dyadic observations are Ind-Nind pairs. Since one direct peg can generate several indirect pegs, we have 8,092 and 16,705 indirect pegs based on the de jure and de facto classifications, respectively, the majority of which are between the Nind-Nind pairs. Further, of the 124 country pairs that have a de jure direct peg, 83 show a change in regime (both on and off a peg) with 69 switching once, eleven switching twice,

[26] See Appendix B for a description of data sources and summary statistics.

[27] We would like to thank Jean Salvati for assistance in STATA coding of the indirect peg variable.

and 3 switching more than twice—leading to a total of 101 switches during 1972-2006. Based on the de facto classification, 106 country pairs switched regimes 143 times, with 78 switching once, 19 switching twice, and 9 switching more than twice.

About 178 country pairs in the full sample share a currency. Of the total 2,121 observations coded as currency unions, about 80 percent are nonindustrial pairs, largely comprising African trading partners, and 15 percent are industrialized pairs. There are 67 country pairs that switch to enter a currency union, of which 59 are the Ind-Ind dyads. These mainly represent the EMU member countries that adopted the Euro between 1999 and 2006.[28]

Table 2. Distribution of regimes in the sample, 1972-2006[1]

	Total	Ind-Ind	Ind-NInd	NInd-NInd
Total				
Observations	177,270	9,710	75,169	92,391
No. of country pairs	10,894	350	3,515	7,029
% of world trade	100%	53%	40%	7.3%
Direct pegs (de jure)				
Observations	1,192	90	1,096	6
No. of country pairs	124	23	100	1
% of world trade	4.1%	2.0%	2.1%	0.1%
No. of switches	101	12	88	1
Direct pegs (de facto)				
Observations	1,625	100	1,487	38
No. of country pairs	143	23	118	2
% of world trade	7.9%	2.0%	5.8%	0.0%
No. of switches	145	13	131	1
Indirect pegs (de jure)				
Observations	8,092	702	1,493	5,897
No. of country pairs	1,786	153	515	1,118
% of world trade	3.9%	3.2%	0.4%	0.3%
Indirect pegs (de facto)				
Observations	16,705	862	2,237	13,606
No. of country pairs	2,622	149	607	1,866
% of world trade	6.8%	3.5%	2.3%	1.1%
Currency unions				
Observations	2,121	324	126	1671
No. of country pairs	178	66	9	103
% of world trade	4.7%	4.5%	0.2%	0.0%
No. of switches to join CU	67	59	1	7

Source: Authors' calculations.
[1] Indicates the sample used in the baseline estimations (where outliers are excluded).
Sum of pairs for Industrialised-Industrialised, Nonindustrialised-Nonindustrialised and Nonindustrialised-Industrialised equals the total (world) sample.
Direct pegs exclude pairs which are members of the same currency union.

[28] For the post-EMU period, we treat direct pegs with Euro as a peg with Germany for all countries but the CFA franc zone. By this definition, all members of the EMU (excluding Germany) would have indirect pegs with countries pegged to the Euro. For the CFA countries, we assume that they retain their peg with France. However, the results are robust to changes in the anchor countries for the Euro-pegged countries.

V. EMPIRICAL RESULTS

A. Main Results

Benchmark specification

World sample

The estimation results for equation (2) for the full sample are presented in Table 3. For completeness and comparison to the results reported in previous studies, we estimate the benchmark specification using both the de jure and de facto classifications with all the estimators discussed earlier, namely, pooled OLS, CFE, CYFE, CPFE, and HT.[29] We then follow the sequential testing procedure suggested in Baltagi, Bresson and Pirotte (2003), and conduct the HT specification tests to select between the various estimation methods.

The results for the de jure and de facto classifications are presented in columns (1)-(5) and columns (6)-(10) of Table 3, respectively. In both cases, for the OLS estimation when only time effects are included along with the traditional gravity variables, currency unions and direct pegs have a significant positive effect on bilateral trade.[30] The estimated impact of long-run exchange rate volatility is significantly negative, while indirect pegs are also found to have a negative effect. These results do not change much when the CFE are included to control for unobserved country-specific characteristics, but the significant F-test on fixed effects indicates the inappropriateness of the OLS method. The addition of CYFE in columns (3) and (8) does little to improve the fit of the model. However, the estimated effect of currency union becomes larger, and we obtain a counter-intuitive result for exchange rate volatility, which is estimated to have a significantly positive effect on bilateral trade flows.

Controlling for the CPFE as in columns (4) and (9), we observe that the estimated trade generating currency union and direct peg effects fall substantially but remain statistically significant. Nevertheless, we lose the cross-sectional information of the data, and all time invariant variables drop from the estimation. To take into account the cross-sectional dimension while allowing for the correlation of some regressors with the individual effects, we estimate equation (2) with the HT method specifying several possible sets of endogenous variables. The choice of endogenous variables rests on economic reasoning but the final set is selected based on a comparison of the HT specification (or Hausman) test for these estimations with the fixed effects estimator. The test results (as reported in the last row of Table 3) suggest that the difference between the FE and HT estimators is not significant enough to reject the appropriateness of the HT estimator when currency union, direct peg, real GDP, real GDP per

[29] We also estimate the benchmark specification with the random effects model. However, in all cases, the Hausman test—based on the differences between the fixed and random effects models—fails to confirm the hypothesis that the explanatory variables are uncorrelated with the unobserved omitted variables.

[30] The signs and magnitude of the estimated coefficients of the traditional gravity variables are plausible and in line with earlier studies. A majority of these variables are statistically significant at the 1 percent level.

capita, distance, and free trade agreement are considered as endogenous variables. Hence, the preferred HT specification reported here takes this set of variables as endogenous.[31]

The estimated trade generating effect of currency unions and direct pegs based on the HT method is quite similar to that obtained from the CPFE approach but different from the CYFE. We interpret the estimated coefficients to indicate that the membership of a currency union—on average—increases bilateral trade by about 36-39 percent.[32] This result is in line with the estimates of recent studies, which report a smaller effect than Rose (2000). Both de jure and de facto direct pegs have a significantly positive effect on bilateral trade, with the size of the estimated effect (35-39 percent) being close to that of currency unions. Considering that the estimated positive impacts of more stable exchange rate regimes are significant despite controlling for exchange rate volatility supports the notion that these regimes promote trade through channels in addition to reduced exchange rate volatility.

The estimated impact of exchange rate volatility is strongly negative. The obtained point estimate implies that increasing exchange rate volatility by one standard deviation would lead to a reduction in bilateral trade by about 5 percent.[33] Interestingly, indirect pegs between dyads are estimated to lower bilateral trade. This result, somewhat surprising, is similar to that obtained by Adam and Cobham (2007).[34] KS (2006) also report a negative estimated coefficient for indirect pegs, but find it to be insignificant. We believe that this result indicates possible trade diversion among bilateral partners toward the anchor country, and explore it further in the context of different sub-samples.

The above results hold when we estimate the benchmark specification using short-run exchange rate volatility, define indirect pegs to include more distant relationships, and include a quadratic term for exchange rate volatility in the benchmark specification (see Tables C1-C4). In the latter case, we find a strongly positive estimated coefficient, indicating a non-linear relationship between exchange rate volatility and bilateral trade.

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[31]We try several possible combinations of the regressors as endogenous variable in the HT method, but present the results for the final (selected) estimation for brevity. As suggested by Baltagi (2001), we also compute the geometric average of the canonical correlations—a measure of the squared correlations between the set of instruments and the regressors—to compare the gains in asymptotic efficiency across different sets of instruments. The average is maximized when currency union, direct peg, real GDP, real GDP per capita, distance, and free trade agreements are taken as endogenous variables.

[32] The effect of currency unions or direct pegs may include both the direct effect, and the estimated indirect effect through exchange rate volatility. Following previous literature, we identify these two effects separately and refer to the estimated direct impact only. The direct effect of currency union is obtained as $e^{0.332}$-1 = 0.394 for the de jure classification and $e^{0.306}$-1 = 0.358 for the de facto classification.

[33] The standard deviation for long run volatility in the world sample is 0.203. The impact is computed as [(exp(0.203)*(-0.240))-1] = -0.048.

[34] The estimated coefficient for Adam and Cobham's binary variable, which is equal to unity if "one country in the pair is pegged to a currency with reference to which the other's currency is managed" is significantly negative for the period 1948-98, but is insignificant for shorter samples.

Sub-samples

Table 4 presents the results of the benchmark specification with the HT method for different groups of countries. The results show a strong trade generating effect of currency unions for the Ind-Ind and Nind-Nind samples, which is almost twice as large for the latter as compared to the former. Both de jure and de facto direct pegs increase bilateral trade between Ind-Ind and Ind-Nind pairs. Interestingly, the effect of direct pegs is larger for pegs between Ind-Nind pairs, than those between industrialized dyads. Further, the latter appear to benefit more from currency unions than from direct pegs. For the Nind-Nind sample, we find a significantly negative effect of de jure direct pegs and an insignificant de facto direct peg effect. Considering the small number of (non-currency union) de jure direct pegs in the Nind-Nind sample (only 6 of a total of 92,391 observations), not much weight can be given to this result.

In contrast to KS (2006), who find a significantly negative effect of indirect pegs for the industrialized pairs, we find a small trade creating effect for this sample. This difference in result is likely to be driven by our extended dataset, which includes years immediately before (and after) the creation of the EMU, when several countries were pegged to the Deutsche Mark, leading to a much higher number of indirect pegs in our Ind-Ind sample than that of KS (2006).[35] These indirect pegs created in the run-up to the EMU are likely to have boosted trade between the potential member countries in anticipation of the formation of the monetary union. However, our results for the other two sub-samples indicate that indirect pegs reduce bilateral trade when at least one country in the dyad is a nonindustrialized country. This result makes intuitive sense since most nonindustrial countries that maintain direct pegs adopt currencies of large, institutionally advanced economies as anchors. Direct pegs could potentially encourage trade with the anchor economies because of reduced transaction costs implying lower trade with other bilateral partners.

Taken together, these findings confirm the view that there may be systematic differences in country characteristics that imply a varying effect of regimes on bilateral trade across different country groups. Thus, as suggested by Baldwin (2006), in order to draw specific inferences about the impact of exchange rate regimes on trade, it is important to go beyond a large sample comprising heterogeneous countries. If countries are too different in factors that are not accounted for in the model, then the linear approximation of the gravity model—which estimates the results for the average country in the sample—may lead to incorrect generalized inferences.

Deeds versus words

The results presented in Tables 3 and 4 suggest that, on average, de facto pegs have a slightly higher effect than de jure pegs. To investigate the extent to which "deeds" matter more than "words", we examine three further cases: (i) deeds match words—when the sample is restricted to include only those dyadic observations where both the de jure and de facto classifications

[35] Restricting the sample to 1972-99 and using CYFE (as in KS, 2006), we find an insignificant effect of indirect pegs.

Table 3. Benchmark specification results for the world sample, 1972-2006

	De jure classification					De facto classification				
Sample	World	World	World	World	World	World	World	World	World	World
Estimation	OLS	CFE	CYFE	CPFE	HT	OLS	CFE	CYFE	CPFE	HT
Specification	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
CU	0.551 ***	0.601 ***	0.707 ***	0.316 ***	0.332 ***	0.520 ***	0.597 ***	0.711 ***	0.291 ***	0.306 ***
	(0.16)	(0.17)	(0.17)	(0.08)	(0.08)	(0.16)	(0.17)	(0.17)	(0.08)	(0.08)
Direct peg	0.365 **	0.492 ***	0.495 ***	0.306 ***	0.303 ***	0.339 ***	0.562 ***	0.601 ***	0.329 ***	0.327 ***
	(0.11)	(0.11)	(0.12)	(0.11)	(0.11)	(0.10)	(0.11)	(0.11)	(0.10)	(0.10)
Indirect peg	-0.247 **	-0.285 ***	-0.294 ***	-0.152 ***	-0.149 ***	-0.355 ***	-0.322 ***	-0.248 ***	-0.228 ***	-0.228 ***
	(0.06)	(0.05)	(0.07)	(0.04)	(0.04)	(0.05)	(0.04)	(0.05)	(0.03)	(0.03)
Volatility	-0.288 ***	-0.201 ***	0.835 ***	-0.254 ***	-0.245 ***	-0.277 ***	-0.198 ***	0.802 ***	-0.249 ***	-0.240 ***
	(0.04)	(0.03)	(0.16)	(0.03)	(0.03)	(0.04)	(0.03)	(0.16)	(0.03)	(0.03)
Lrgdp	1.124 ***	0.470 ***		1.039 ***	1.193 ***	1.125 ***	0.460 ***		1.027 ***	1.175 ***
	(0.01)	(0.08)		(0.07)	(0.05)	(0.01)	(0.08)		(0.07)	(0.05)
Lrgdppc	0.019	0.731 ***		0.214 ***	0.127 **	0.015	0.743 ***		0.228 ***	0.146 **
	(0.01)	(0.08)		(0.07)	(0.06)	(0.01)	(0.08)		(0.07)	(0.06)
Ldist	-1.226 ***	-1.512 ***	-1.519 ***		-1.809 ***	-1.231 ***	-1.514 ***	-1.517 ***		-1.884 ***
	(0.02)	(0.03)	(0.03)		(0.13)	(0.02)	(0.03)	(0.03)		(0.17)
Comlang	0.498 ***	0.501 ***	0.493 ***		0.549 ***	0.501 ***	0.506 ***	0.496 ***		0.543 ***
	(0.05)	(0.05)	(0.05)		(0.05)	(0.05)	(0.05)	(0.05)		(0.05)
Comborder	0.621 ***	0.420 ***	0.405 ***		-0.015	0.624 ***	0.412 ***	0.401 ***		-0.157
	(0.14)	(0.15)	(0.15)		(0.31)	(0.14)	(0.15)	(0.15)		(0.37)
Fta	1.300 ***	0.662 ***	0.692 ***	0.248 ***	0.268 ***	1.305 ***	0.670 ***	0.694	0.262 ***	0.283 ***
	(0.11)	(0.10)	(0.10)	(0.05)	(0.05)	(0.10)	(0.10)	(0.10)	(0.05)	(0.05)
Landl	-0.313 ***	-1.258 ***			-0.155 ***	-0.324 ***	-1.270			-0.172 ***
	(0.03)	(0.29)			(0.05)	(0.03)	(0.29)			(0.06)
Island	0.110 ***	-1.007 ***			0.374 ***	0.112 ***	-1.001 **			0.388 ***
	(0.04)	(0.39)			(0.07)	(0.04)	(0.39)			(0.07)
Lareap	-0.079 ***	0.362 ***			-0.028	-0.080 ***	0.365 ***			-0.015
	(0.01)	(0.05)			(0.03)	(0.01)	(0.05)			(0.03)
Comcol	0.748 ***	0.656 ***	0.676 ***		1.076 ***	0.740 ***	0.657 ***	0.676 ***		1.062 ***
	(0.08)	(0.07)	(0.07)		(0.10)	(0.08)	(0.07)	(0.07)		(0.11)
Curcol	-0.128	-0.0524	-0.005	-0.376	-0.383	-0.137	-0.048	0.012	-0.382	-0.387
	(0.58)	(0.72)	(0.53)	(0.702)	(0.70)	(0.58)	(0.73)	(0.53)	(0.70)	(0.70)
Evercol	1.217 ***	1.348 ***	1.333 ***		1.070 ***	1.197 ***	1.326 ***	1.309 ***		1.062 ***
	(0.14)	(0.14)	(0.14)		(0.17)	(0.14)	(0.14)	(0.14)		(0.17)
Comctry	1.627 ***	1.159	1.304		2.007 ***	1.645 ***	1.161	1.303		2.043 ***
	(0.57)	(0.77)	(0.60)		(0.71)	(0.57)	(0.77)	(0.60)		(0.72)
Constant	-26.990 ***	-14.720 ***	26.977 ***	-38.550 ***	-28.640 ***	-26.860 ***	-14.480 ***	27.068 ***	-38.170 ***	-27.750 ***
	(0.37)	(2.32)	(0.97)	(2.65)	(1.71)	(0.37)	(2.31)	(0.94)	(2.64)	(2.13)
Observations	177,270	177,270	177,270	177,270	177,270	177,270	177,270	177,270	177,270	177,270
Number of pairs				10,894	10,894				10,894	10,894
R-squared	0.722	0.77	0.79			0.72	0.77	0.79		
R2-within				0.14					0.14	
R2-between				0.77					0.67	
R2-overall				0.72					0.61	
Hausman chi-2 (FE vs. RE)[1]				0.00					0.00	
Hausman chi-2 (HT vs. RE)[2]					0.00					0.00
Hausman chi-2 (FE vs. HT)[3]					0.78					0.62

Source: Authors' calculations.
Robust clustered standard errors in parentheses; Time effects included in all specifications.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
Volatility refers to long-run volatility computed over 36-month horizon.
Variables instrumented for in HT: CU, Direct peg, Lrgdp, Lrgdppc, Ldist, FTA.
[1] Hausman test applied to the difference between the within (fixed effects) and GLS (random effects) estimators.
[2] Hausman test applied to the difference between the HT estimators and GLS (random effects).
[3] Hausman test applied to the difference between the within (fixed effects) and HT estimators.

Table 4. Benchmark specification results for the sub-samples, 1972-2006

	De jure classification			De facto classification		
Sample	Ind-Ind	Ind-Nind	Nind-Nind	Ind-Ind	Ind-Nind	Nind-Nind
Estimation	HT	HT	HT	HT	HT	HT
Specification	(1)	(2)	(3)	(4)	(5)	(6)
CU	0.291 ***	0.190	0.640 **	0.319 ***	0.203	0.631 **
	(0.06)	(0.22)	(0.31)	(0.06)	(0.22)	(0.31)
Direct peg	0.180 ***	0.251 **	-0.655 ***	0.207 ***	0.283 ***	0.128
	(0.05)	(0.12)	(0.03)	(0.05)	(0.11)	(0.09)
Indirect peg	0.042	-0.027	-0.171 ***	0.087 **	-0.101 **	-0.230 ***
	(0.04)	(0.05)	(0.05)	(0.04)	(0.05)	(0.03)
Volatility	-0.031	-0.130 ***	-0.340 ***	-0.030	-0.130 ***	-0.333 ***
	(0.03)	(0.03)	(0.06)	(0.03)	(0.03)	(0.06)
Lrgdp	0.406 ***	0.665 ***	1.116 ***	0.431 ***	0.672 ***	1.128 ***
	(0.14)	(0.09)	(0.07)	(0.15)	(0.09)	(0.09)
Lrgdppc	0.832 ***	0.697 ***	0.106	0.799 ***	0.692 ***	0.0907
	(0.21)	(0.09)	(0.08)	(0.21)	(0.09)	(0.09)
Ldist	-0.374 ***	-0.609 ***	-2.226 ***	-0.382 ***	-0.680 ***	-2.321 ***
	(0.10)	(0.15)	(0.19)	(0.10)	(0.14)	(0.29)
Comlang	0.561 ***	0.432 ***	0.285 ***	0.565 ***	0.446 ***	0.259 ***
	(0.18)	(0.10)	(0.09)	(0.18)	(0.10)	(0.12)
Comborder	1.039 ***	1.085 *	-0.089	0.998 ***	0.920	-0.264
	(0.29)	(0.60)	(0.38)	(0.28)	(0.58)	(0.56)
Fta	0.228 ***	0.314 ***	0.063	0.221 ***	0.344 ***	0.0634
	(0.04)	(0.06)	(0.10)	(0.04)	(0.06)	(0.10)
Landl	-0.978 ***	-0.393 ***	-0.323 ***	-0.960 ***	-0.392 ***	-0.338 ***
	(0.19)	(0.07)	(0.09)	(0.19)	(0.07)	(0.11)
Island	-0.588 **	-0.762 ***	0.545 ***	-0.553 **	-0.733 ***	0.584 ***
	(0.27)	(0.13)	(0.08)	(0.27)	(0.13)	(0.10)
Lareap	0.054	0.137 ***	0.060	0.046	0.134 ***	0.057
	(0.05)	(0.04)	(0.05)	(0.05)	(0.04)	(0.06)
Comcol	0.873 *	1.079 ***	1.029 ***	0.861 *	1.071 ***	1.001 ***
	(0.52)	(0.23)	(0.12)	(0.51)	(0.22)	(0.14)
Curcol	-0.083 ***	-0.538		-0.084 ***	-0.540	
	(0.02)	(0.86)		(0.03)	(0.86)	
Evercol	0.749 ***	1.709 ***	-0.725 *	0.734 ***	1.688 ***	-0.815 *
	(0.22)	(0.17)	(0.43)	(0.21)	(0.16)	(0.45)
Comctry		1.121			1.131	
		(0.87)			(0.87)	
Constant	-15.760 ***	-26.420 ***	-22.990 ***	-16.150 ***	-26.010 ***	-22.430 ***
	(3.60)	(2.28)	(2.92)	(3.60)	(2.30)	(4.05)
Observations	9,710	75169	92,391	9,710	75169	92,391
Number of pairs	350	3,515	7,029	350	3,515	7,029
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	1.00	0.55	0.09	0.91	0.54	0.05

Source: Authors' calculations.
Robust clustered standard errors in parentheses; Time effects included in all specifications.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
Volatility refers to long-run volatility computed over 36-month horizon.
Variables instrumented for in HT: CU, Direct peg, Lrgdp, Lrgdppc, Ldist, FTA.
[1] Hausman test applied to the difference between the HT estimators and GLS (random effects).
[2] Hausman test applied to the difference between the within (fixed effects) and HT estimators.

Table 5. Results for deeds versus words for the world and sub-samples, 1972-2006

	Deeds match words[a]				Mirage of fixed[b]				Fear of floating[c]			
Sample	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
Estimation		HT	HT	HT	HT	HT	HT	HT	HT	HT	HT	HT
Specification		(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
CU	0.344 ***	0.295 ***	0.209	0.641 **	0.309 ***	0.272 ***	0.158	0.640 **	0.281 ***	0.297 ***	0.161	0.632 **
	(0.08)	(0.06)	(0.23)	(0.31)	(0.08)	(0.06)	(0.22)	(0.31)	(0.08)	(0.06)	(0.23)	(0.31)
Direct peg	0.416 ***	0.192 ***	0.375 **	-0.364 ***	-0.262	0.264 ***	-0.381 *	-0.651 ***	0.087	0.177 ***	0.077	0.193
	(0.13)	(0.05)	(0.14)	(0.04)	(0.18)	(0.04)	(0.21)	(0.03)	(0.08)	(0.05)	(0.09)	(0.35)
Indirect peg	-0.148 ***	0.043	-0.027	-0.171 ***	-0.150 ***	0.037	-0.031	-0.171 ***	-0.229 ***	0.082 **	-0.104 **	-0.230 ***
	(0.04)	(0.04)	(0.05)	(0.05)	(0.04)	(0.04)	(0.05)	(0.05)	(0.03)	(0.04)	(0.05)	(0.03)
Volatility[d]	-0.245 ***	-0.030	-0.130 ***	-0.340 ***	-0.244 ***	-0.031	-0.130 ***	-0.340 ***	-0.240 ***	-0.0299	-0.129 ***	-0.333 ***
	(0.03)	(0.03)	(0.03)	(0.06)	(0.03)	(0.03)	(0.03)	(0.06)	(0.02)	(0.03)	(0.03)	(0.07)
Lrgdp	1.193 ***	0.412 ***	0.663 ***	1.116 ***	1.193 ***	0.398 ***	0.660 ***	1.116 ***	1.175 ***	0.420 ***	0.673 ***	1.128 ***
	(0.05)	(0.15)	(0.09)	(0.07)	(0.05)	(0.15)	(0.09)	(0.07)	(0.05)	(0.15)	(0.09)	(0.09)
Lrgdppc	0.126 **	0.823 ***	0.697 ***	0.107	0.127 **	0.843 ***	0.702 ***	0.106	0.146 **	0.814 ***	0.693 ***	0.091
	(0.06)	(0.21)	(0.09)	(0.08)	(0.06)	(0.21)	(0.09)	(0.08)	(0.06)	(0.21)	(0.09)	(0.09)
Ldist	-1.813 ***	-0.378 ***	-0.646 ***	-2.224 ***	-1.806 ***	-0.372 ***	-0.609 ***	-2.227 ***	-1.884 ***	-0.378 ***	-0.676 ***	-2.320 ***
	(0.13)	(0.10)	(0.15)	(0.19)	(0.13)	(0.10)	(0.15)	(0.19)	(0.17)	(0.10)	(0.14)	(0.29)
Comlang	0.548 ***	0.560 ***	0.441 ***	0.286 ***	0.552 ***	0.557 ***	0.438 ***	0.285 ***	0.547 ***	0.563 ***	0.452 ***	0.259 **
	(0.05)	(0.18)	(0.09)	(0.09)	(0.05)	(0.19)	(0.09)	(0.09)	(0.05)	(0.18)	(0.09)	(0.12)
Comborder	-0.024	1.025 ***	1.039 *	-0.087	-0.008	1.059 ***	1.083 *	-0.091	-0.153	1.025 ***	0.919	-0.262
	(0.31)	(0.29)	(0.61)	(0.38)	(0.30)	(0.29)	(0.60)	(0.38)	(0.37)	(0.29)	(0.58)	(0.56)
Fta	0.266 ***	0.229 ***	0.306 ***	0.063	0.270 ***	0.232 ***	0.316 ***	0.063	0.284 ***	0.224 ***	0.342 ***	0.064
	(0.05)	(0.04)	(0.06)	(0.10)	(0.04)	(0.04)	(0.06)	(0.10)	(0.05)	(0.04)	(0.06)	(0.10)
Landl	-0.155 ***	-0.977 ***	-0.398 ***	-0.322 ***	-0.155 ***	-0.995 ***	-0.396 ***	-0.323 ***	-0.174 ***	-0.975 ***	-0.392 ***	-0.338 ***
	(0.05)	(0.19)	(0.07)	(0.09)	(0.05)	(0.19)	(0.07)	(0.09)	(0.06)	(0.19)	(0.07)	(0.11)
Island	0.375 ***	-0.582 **	-0.751 ***	0.545 ***	0.373 ***	-0.607 **	-0.768 ***	0.545 ***	0.388 ***	-0.573 **	-0.736 ***	0.584 ***
	(0.07)	(0.28)	(0.13)	(0.08)	(0.07)	(0.28)	(0.13)	(0.08)	(0.07)	(0.27)	(0.13)	(0.10)
Lareap	-0.028	0.052	0.136 ***	0.060	-0.028 ***	0.056	0.140 ***	0.060	-0.010 ***	0.049	0.135 ***	0.057
	(0.03)	(0.05)	(0.04)	(0.05)	(0.03)	(0.05)	(0.04)	(0.05)	(0.03)	(0.05)	(0.04)	(0.06)
Comcol	1.075 ***	0.869 *	1.073 ***	1.030 ***	1.077 ***	0.879 *	1.083 ***	1.029 ***	1.061 ***	0.867 *	1.072 ***	1.001 ***
	(0.10)	(0.52)	(0.23)	(0.12)	(0.10)	(0.53)	(0.23)	(0.12)	(0.11)	(0.52)	(0.23)	(0.14)
Curcol	-0.380	-0.083 ***	-0.536		-0.387	-0.082 ***	-0.543		-0.393	-0.083 ***	-0.547	
	(0.70)	(0.02)	(0.86)		(0.70)	(0.02)	(0.86)		(0.70)	(0.02)	(0.86)	
Evercol	1.054 ***	0.747 ***	1.683 ***	-0.723 *	1.102 ***	0.753 ***	1.741 ***	-0.726 *	1.100 ***	0.738 ***	1.720 ***	-0.815 *
	(0.16)	(0.22)	(0.17)	(0.43)	(0.17)	(0.22)	(0.17)	(0.43)	(0.17)	(0.22)	(0.17)	(0.45)
Comctry	2.022 ***		1.133		1.977 ***		1.096		2.013 ***		1.103	
	(0.72)		(0.87)		(0.72)		(0.87)		(0.72)		(0.87)	
Constant	-28.610 ***	-15.810 ***	-25.960 ***	-23.020 ***	-28.690 ***	-15.620 ***	-26.350 ***	-22.990 ***	-27.740 ***	-15.990 ***	-26.100 ***	-22.430 ***
	(1.72)	(3.61)	(2.34)	(2.92)	(1.71)	(3.62)	(2.28)	(2.92)	(2.11)	(3.60)	(2.26)	(4.05)
Observations	176,749	9,698	74,698	92,353	177,270	9,710	75169	92,391	177,270	9,710	75169	92,391
Number of pairs	10,892	350	3,514	7,028	10,894	350	3,515	7,029	10,894	350	3,515	7,029
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	0.76	0.85	0.59	0.09	0.79	0.99	0.35	0.09	0.64	0.99	0.00	0.03

Source: Authors' calculations.

Robust clustered standard errors in parentheses; Time effects included in all specifications; ***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.

Indirect peg refers to relation=2 in Figure A1.

[a] Sample comprises observations for which the de jure classification is the same as de facto classification.

[b] Direct peg is unity if de jure indicates a peg and de facto is a non-peg; zero otherwise.

[c] Direct peg is unity if de jure indicates a non-peg and de facto is a peg; zero otherwise.

[d] Refers to long-run volatility over the 36-month horizon.

[1] Hausman test applied to the difference between the GLS (random effects) and HT estimators.

[2] Hausman test applied to the difference between the HT estimators and GLS (random effects).

indicate the same exchange rate arrangement (peg or non-peg); (ii) the “fear of floating”—when there is a de jure non-peg, but the de facto classification suggests a direct peg; and (iii) the “mirage of fixed”—when the de jure classification indicates a direct peg but the de facto arrangement is a non-peg.

Table 5 presents the results for the three scenarios. Clearly, the positive effect of a direct peg on trade is the largest when deeds match words. Interestingly, the estimated coefficient for direct pegs is also larger than those reported for the de jure and de facto pegs in Table 3, suggesting that increased transparency of exchange rate regime choice amplifies the impact of more stable regimes. For the case where central bank behavior indicates a peg but the announced regime is a non-peg—the fear of float scenario—the estimated coefficient of the binary variable reflecting direct pegs is positive, but insignificant. Similarly, when the de jure classification is a peg but the de facto classification indicates otherwise—the mirage of fixed scenario—the effect of direct pegs is found to be negative but not statistically significant.

Looking across the sub-samples, we find that the strong effect of words complemented by deeds also holds for the Ind-Ind and Ind-Nind dyads. In both cases, we find the impact of direct pegs to be larger than those reported in Table 4. However, for the scenarios when deeds do not match words, it is difficult to draw inferences for the individual sub-samples since the number of observations coded as a de jure peg and a de facto non-peg, or, a de jure non-peg and a de facto peg, is relatively small across the three groups.

Does distance matter?

The results obtained from the estimation of equation (5) indicate that distance undermines the trade generating effect of currency unions but does not influence the impact of direct pegs.[36] The estimated coefficient of currency unions ($\hat{\gamma}$) is positive and statistically significant, while $\hat{\phi}$ is significantly negative in all specifications—indicating that the currency union effect diminishes with geographical distance.

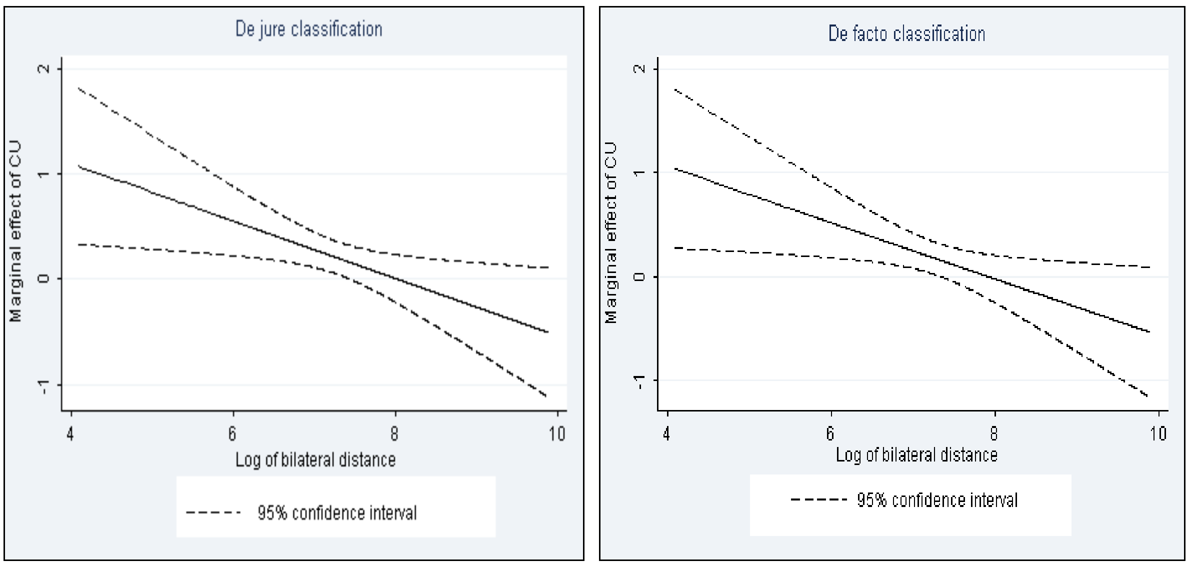
Figure 2 illustrates the marginal effect of currency unions on bilateral trade across the observed range of geographical distance for the world sample. The solid downward sloping line in both panels shows the change in the impact of currency unions as geographical distance between the trading partners increases. The significance of the effect can be assessed by considering the 95 percent confidence interval around it, with the effect being significant when the upper and lower bounds of the confidence interval are both either above or below zero.[37]

[36] We estimate equation (5) by including each interaction term individually, and then both together. In all the cases, the interaction term for direct pegs and distance is statistically insignificant. Due to multicollinearity concerns between the direct peg variable and the interaction term, we also estimate the effect of direct peg variable for different quartiles of distance. However, the estimated marginal effect of direct pegs does not vary in a systematic and meaningful way across the quartiles.

[37] On average, the currency union effect is insignificant for the world sample at a high threshold of geographical distance, with about 70 dyads sharing the same currency located at a distance larger than the threshold, most of which belong to the Ind-Nind and Nind-Nind samples.

The plot in both panels supports the hypothesis that the impact of currency unions depends negatively on bilateral geographical distance, and highlights the importance of modeling a non-linear relationship between the two variables based on other trading costs to derive a better approximation of the data. This result holds for both Ind-Ind and Nind-Nind dyads, but the distance threshold beyond which the currency union is insignificant is higher for the former than the latter. Taking distance as the proxy for other trading costs, notably, transportation and information costs, this result implies that if such costs are high then price convergence may be lower, and risks of entry sufficiently higher to deter countries from engaging in trade despite a common currency.

Figure 2. Marginal effect of currency unions on distance for the world sample



Source: Authors' calculations.

Trade stability

The results of the impact of fixed exchange rates on trade stability are presented in Table 6. Like before, we estimate equation (6) using the OLS, CFE, CYFE, CPFE and HT, but present the results for the HT method. The total number of observations is about one-fifth of that for the benchmark specification, since the dependent variable (the coefficient of variation of real bilateral trade flows) is computed over a five-year period, and the right hand side variables are averaged over the corresponding period.

The results for the world sample show that trading pairs with direct pegs experience less volatile trade. The effect is statistically significant, and of the same magnitude for both de jure and de facto classifications. Despite indicating a stronger commitment to exchange rate stability, currency unions do not appear to have a strong impact on reducing trade instability (except through reduced exchange rate volatility). This finding is consistent with those obtained by Tsangarides et al. (2008), who find an ambiguous effect of currency unions on trade stability using a much longer sample (1948-2002).

For the sub-samples considered, both direct pegs and currency unions appear to have the strongest effect of reducing trade volatility for the Ind-Nind trading partners, but not much effect in all the other cases. Indirect pegs appear to make bilateral trade more unstable, with the effect being the strongest for the non-industrialized dyads. Although the channels through which this effect sets in are less apparent, trade diversion toward anchor country could be a possible source of generating instability in links with other trading partners.

B. Sensitivity Analysis

Model specification and estimation

To check the robustness of our main set of results, we change the model specification and estimation methodology in several ways. First, we construct the dependent variable as the average of the logarithm of exports and imports (rather than the logarithm of the average) as proposed by some critics, and re-estimate the benchmark specification using both the de jure and de facto classifications. Second, we add quadratic terms for output and output per capita to equation (2) to control for possible sample nonlinearities. In both cases, the results are similar to those obtained earlier: the estimated impact of currency unions and de jure pegs remains significantly positive, while exchange rate volatility reduces bilateral trade (see Table C5).

Third, we control for the possible sample selection bias in our estimation. As discussed in Carrère (2006), unbalanced samples could be subject to a selection bias if the probability of a trading pair being included in the sample is dependent on the model error term, and, in particular, to the unobserved bilateral effects. Following Nijman and Verbeek (1992), we control for the selection bias by including variables that reflect each dyad's presence in the sample, such as: (i) the number of years the pair is present in the sample; and (ii) a dummy variable equal to unity if the pair is observed during the entire period, and zero otherwise.[38] Once again, the results are similar to those reported in Tables 3 and 4, and the estimated coefficients of the variables of interest retain their signs and significance.

Finally, we address the issue of zero-trade observations that commonly arises in bilateral datasets either because some dyads did not trade, or because of rounding errors and missing observations. Using the log-linear form of the gravity equation as in equation (2) implies including only those observations for which the dependent variable is positive. Given that trade flows between some pairs of countries—typically pairs of small countries—tends to be zero, truncation at zero may result in inconsistent estimators when ordinary least squares (OLS) are used. We check the sensitivity of our results to the inclusion of zero-trade observations by applying the Poisson pseudo maximum likelihood (PPML) approach applied by Santos Silva and Tenreyro (2006). The PPML approach takes the real value of trade as the dependent variable, and includes zero observations. The results obtained from this approach

[38] The estimated coefficients of both variables are positive and significantly different from zero, suggesting that country-pairs with higher frequency of data exhibit greater bilateral trade than those with interrupted data.

are broadly similar to those obtained earlier except for the effect of indirect pegs, which appears positive for the de jure classification, and insignificant for the de facto classification.

Measures of exchange rate volatility and distance

Using the other two volatility measures—*Vol*2 and *Vol*3—discussed earlier does not alter the main set of results significantly. Like before, fixed exchange rate regimes are found to promote bilateral trade, while, on average, indirect pegs have a dampening effect.

In addition to volatility, we check the sensitivity of our results to different definitions of distance. For example, Melitz (2003) argues that distance between the most populous cities of bilateral trading partners is a better proxy for transport costs than the commonly used definition of distance between geographic centers. Thus, we estimate equation (5) using two other bilateral distance measures: distance between the capital cities; and, distance between the largest cities of the two countries, with the inter-city distances being weighted by the share of the city in total population of the country. Neither of the two measures is found to have a significant impact on the trade-generating effect of direct pegs, but—as before—the marginal effect of currency unions diminishes with distance (see Figure C1).

Alternate sub-samples

To make sure that our results are not being driven by economies with certain characteristics, we estimate the benchmark specification for two alternate sub-samples: non-oil exporting economies and countries with a population of over 1 million. We drop oil exporting economies from the sample as determinants of oil trade are different from other trade flows.[39] However, this exercise leaves the results of the benchmark specification completely unchanged. We also drop the micro states (with population less than 1 million) from the sample on the ground that they are not representative of the average country in our sample. In fact, critics argue that the currency union effect may be driven by the inclusion of small states in the sample, which tend to gain the most from joining a currency union. We lose about one-fourth of the observations in this case, but the estimated currency union effect—although slightly smaller—remains significant at the one percent level.

[39] Oil exporters include the WEO fuel exporters (Algeria, Angola, Azerbaijan, Congo, Equatorial Guinea, Gabon, Iran, Iraq, Kuwait, Libya, Nigeria, Oman, Qatar, Saudi Arabia, Turkmenistan, United Arab Emirates, Venezuela, Yemen), Kazakhstan and Norway.

Table 6. Trade stability for the world and sub-samples, 1972-2006

	De jure classification				De facto classification			
Sample	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
Estimation		HT	HT	HT	HT	HT	HT	HT
Specification		(2)	(3)	(4)	(5)	(6)	(7)	(8)
CU	0.002	-0.001	-0.008 **	-0.005	0.002	-0.001	-0.008 **	-0.004
	(0.01)	(0.00)	(0.00)	(0.03)	(0.01)	(0.00)	(0.00)	(0.03)
Direct peg	-0.008 **	0.000	-0.010 ***	0.007	-0.008 **	0.000	-0.010 ***	-0.347
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.40)
Indirect peg	0.005 **	0.001 *	-0.002	0.010 ***	0.009 ***	0.001	0.003	0.011 ***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Volatility	0.009 *	-0.003	0.000	0.012	0.010 *	-0.003	0.000	0.012
	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.01)
Ldist	-0.043 ***	-0.004 ***	-0.023 ***	-0.044 ***	-0.043 ***	-0.004 ***	-0.023 ***	-0.044 ***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Lrgdp	0.022 ***	-0.001	0.011 ***	0.024 ***	0.021 ***	-0.001	0.011 ***	0.024 ***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Lrgdppc	0.059 ***	0.002 **	0.005	0.097 ***	0.043 ***	0.002 **	0.005	0.103 ***
	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.02)
Comlang	-0.015 ***	-0.002 **	-0.014 ***	0.006	-0.017 ***	-0.003 **	-0.014 ***	0.007
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)
Comborder	0.071 ***	0.005 **	0.027 *	0.119 ***	0.040 **	0.006 **	0.025 *	0.132 ***
	(0.02)	(0.00)	(0.02)	(0.02)	(0.02)	(0.00)	(0.01)	(0.03)
Fta	-0.004	0.000	0.002	-0.008	-0.004	0.000	0.001	-0.008
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)
Landl	-0.008 ***	0.001	-0.005 **	0.008 **	-0.009 ***	0.000	-0.005 **	0.010 *
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)
Island	-0.026 ***	-0.002	-0.003	-0.012 ***	-0.020 ***	-0.002	-0.002	-0.014 **
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)
Lareap	0.015 ***	0.001 **	0.009 ***	0.015 ***	0.015 ***	0.001 **	0.008 ***	0.015 ***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Comcol	-0.016 ***	-0.004	0.008	-0.021 ***	-0.020 ***	-0.004	0.008	-0.019 **
	(0.01)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)
Curcol	0.008	0.001 **	0.003		0.011	0.001 **	0.003	
	(0.01)	(0.00)	(0.01)		(0.01)	(0.00)	(0.01)	
Evercol	0.025 ***	0.000	0.003	0.091 ***	0.025 ***	0.000	0.003	0.092 **
	(0.01)	(0.00)	(0.00)	(0.03)	(0.01)	(0.00)	(0.00)	(0.04)
Constant	0.893 ***	0.203 ***	0.734 ***	0.546	1.007 ***	0.202 ***	0.729 ***	0.485 **
	(0.06)	(0.04)	(0.07)	(0.12)	(0.08)	(0.04)	(0.06)	(0.19)
Observations	33,203	1,790	14,047	17,366	33,203	1,790	14,047	17,366
Number of pairs	9,709	370	3,427	6,015	9,709	370	3,427	6,015
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	0.08	0.26	0.01	0.26	0.26	1.00	0.04	0.45

Source: Authors' calculations.
Dependent variable is the coefficient of variation of (log of) real trade computed over 5 years; Indpenedent variables are averaged over 5-yr. periods.
Robust clustered standard errors in parentheses.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
Volatility refers to long-run volatility computed over 36-month horizon.
Variables instrumented for in HT: CU, Direct peg, Lrgdp, Lrgdppc, Ldist, FTA.
[1] Hausman test applied to the difference between the HT estimators and GLS (random effects).
[2] Hausman test applied to the difference between the within (fixed effects) and HT estimators.

VI. Conclusion and Policy Implications

This paper provides some insights into the impact of currency unions and other fixed exchange rate arrangements on bilateral trade. Using an extended dataset on IMF's de jure and de facto classifications, and addressing omitted variable and potential endogeneity concerns, we find that more stable regimes promote bilateral trade as well as trade stability. The trade-generating effect of direct pegs is derived from both de jure and de facto classifications, suggesting that the commitment of the monetary authorities to a stable nominal anchor along with actual behavior have significant effects on trade outcomes. In addition, the impact is the largest when words are complemented by deeds, suggesting that increased transparency in regime choice anchors expectations and improves credibility, amplifying the effect of more stable regimes on trade outcomes. These results are robust to different model specifications, variable definitions, and econometric methodologies.

Our results also show that the impact of fixed exchange rate regimes differ across sub-samples, hence caution should be exercised in drawing generalized inferences. For example, the currency union effect holds for both industrialized and nonindustrialized trading pairs, but the impact is almost twice as large for the latter. Similarly, we find that industrialized and non-industrialized trading partners benefit more from direct pegs as compared to sharing a common currency. Importantly, we find that the effect of currency unions is conditional on the distance between member countries, such that countries located geographically close benefit the most. In this respect, earlier studies that discount the effect of other trading costs and assume a uniform effect of currency unions on all members tell only part of the story. This finding has important policy implications, particularly, for nonindustrialized economies, and suggests that lowering transportation and information costs through, for example, improved infrastructure, could be important elements to increase trade gains from joining currency unions.

Interestingly, we find the bilateral trade benefits from direct pegs to be—on average—at least as large as those from currency unions. Thus, notwithstanding the costs of fixed exchange rate arrangements, countries that want to expand cross-border trade activity through a more stable exchange rate regime while retaining some flexibility, could opt for conventional pegged arrangements rather than complete monetary integration. Although we find some evidence of trade diversion through direct pegs, the effect is small compared to their trade generating effect.

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APPENDIX A

Distribution of regime classification

The IMF's de jure and de facto classifications group exchange rate regimes into eight categories: exchange arrangement with no separate legal tender, currency board arrangement, conventional pegged arrangement, pegged exchange rates within horizontal bands, crawling peg, crawling band, managed float with no predetermined path for the exchange rate, and, independently floating arrangement.[40] To examine the distribution of regimes across countries, we group the first three arrangements (excluding peg to a basket) as the fixed exchange rate regime; group the next four (including peg to a basket) as the intermediate regime; and classify the last one as the floating regime.

Table 1 shows the distribution of countries across the three aggregate groups for both de jure and de facto classifications. Clearly, the classifications are not identical but quite similar in terms of the distribution of observations across the different regimes, and reveal some interesting trends. First, the share of pegs in the de facto classification is consistently higher than in the de jure classification, supporting the "hidden pegs" phenomenon of LYS (2003). Second, there appears to be a consistent decline in the share of intermediate regimes, as suggested by Eichengreen's (1994) "hollowing-out" hypothesis, but this seems to be largely driven by the advanced and emerging countries (Figure A1). In fact, the share of intermediate regimes since the late 1980s appears to be broadly stable for the developing economies.[41] Third, we find that the proportion of floating regimes in the de jure classification is lower than the de facto classification, providing support for Calvo and Reinhart's (2000) "fear of floating" hypothesis.

Table A1: Distribution based on IMF's classification of exchange rate regimes, 1972-2006*

(in percent of total observations)

	1972-79	1980-89	1990-99	2000-06	Full sample 1972-2006
			De jure classification[a]		
Pegged	44.2	38.1	29.7	37.0	36.4
Intermediate	45.5	53.1	46.6	36.9	45.9
Floating	10.3	8.8	23.7	26.2	17.7
			De facto classification[b]		
Pegged	63.9	45.4	36.3	46.5	46.3
Intermediate	33.3	50.8	50.3	36.1	43.9
Floating	2.8	3.8	13.4	17.4	9.8

Source: Authors' calculations based on Anderson (2008).

* Pegged include hard and conventional pegs; intermediate includes basket pegs, pegged within bands, and managed floats; and floating includes independent floats.

[a] De jure classification refers to the officially announced classification.

[b] De facto classification indicates the actual policy followed by the country.

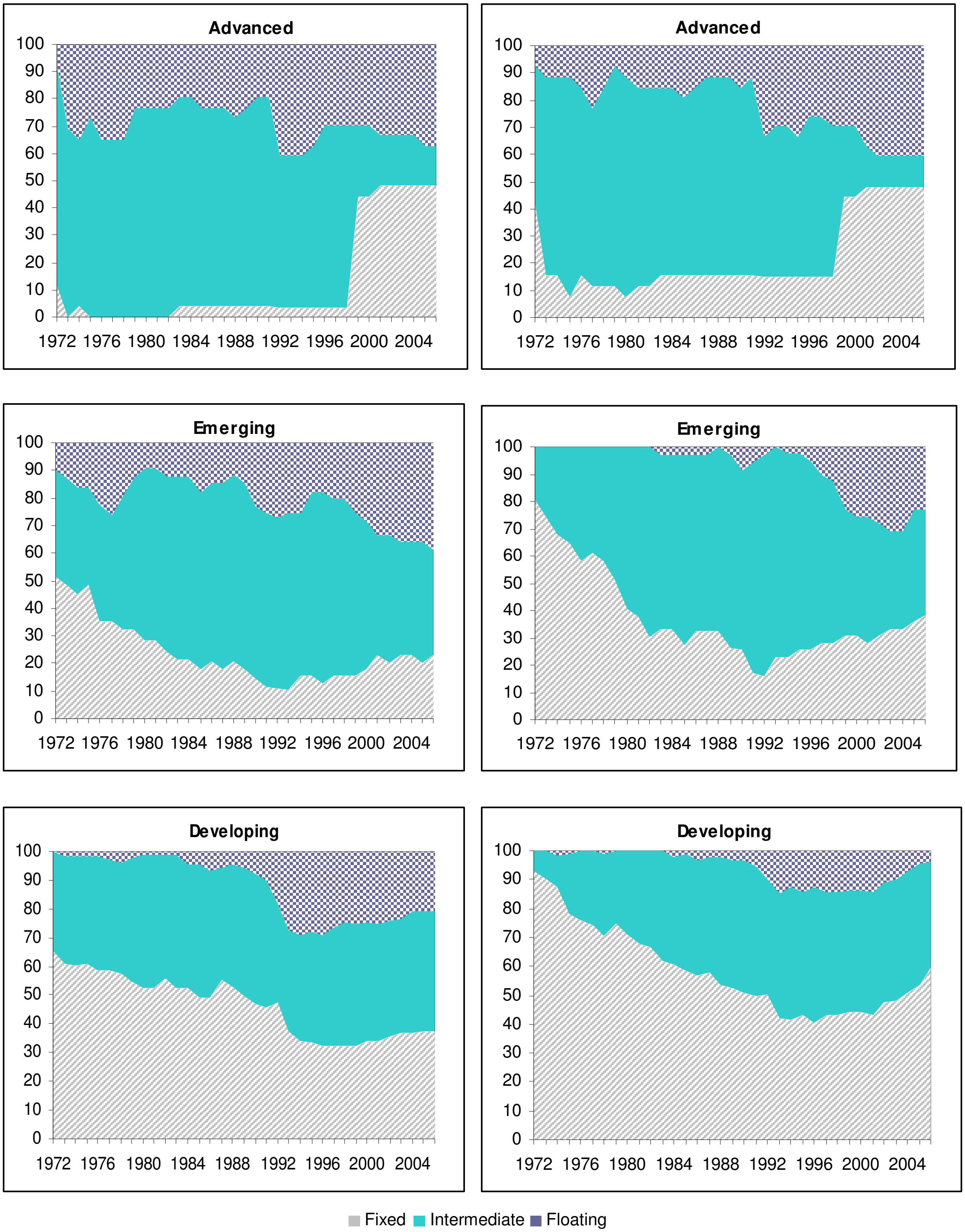
[40] From April 2009, IMF has revised the classification methodology for de facto exchange rate arrangements, which entails changes to the titles and definitions of the categories.

[41] LYS (2003) note that this is in line with the argument that exposure to strong capital inflows may be necessary for the hollowing-out pattern to develop.

Figure A1. Distribution of exchange rate regimes across country groups, 1972-2006 (in percent)

(a) IMF de jure classification

(b) IMF de facto classification



Source: Authors' calculations based on Anderson (2008).

APPENDIX B

Table B1: Variable description and data sources

Variable	Description	Source
Dependent variable		
$lrtrade_{ijt}$	Log of the average value of real bilateral trade between i and j at time t	IMF's *Direction of Trade (DoT)*; Average of exports from a to b, and b to a; and import into a from b, and to b from a. Deflated by U.S. CPI for urban consumers.
Explanatory variables		
cu_{ijt}	Binary variable which is unity if i and j share currency at time t	Anderson (2008)
Direct peg (De jure)$_{ijt}$	Binary variable which is unity if i and j are pegged to each other at time t	Anderson (2008)
Direct peg (De facto)$_{ijt}$	Binary variable which is unity if i and j are pegged to each other at time t	Anderson (2008)
Volatility	Exchange rate volatility	Information Notice System.
$Lrgdp_{ijt}$	Log of the product of real GDP of i and j at time t	World Bank's *World Development Indicators (WDI)*
$Lrgdp_{ijt}$	Log of the product of real GDP per capita of i and j at time t	*WDI*
$Ldist_{ij}$	Log of the distance between i and j	CEPII(http://www.cepii.fr/anglaisgraph/bdd/distances.htm)
$Ldist\_cap_{ij}$	Log of the distance between capital cities of i and j	CEPII(http://www.cepii.fr/anglaisgraph/bdd/distances.htm)
$Ldist\_wces_{ij}$	Log of population weighted distance between the largest cities of i and j	CEPII(http://www.cepii.fr/anglaisgraph/bdd/distances.htm)
$Lang_{ij}$	Binary variable which is unity if i and j have a common language	CIA's *World Factbook* and Rose (2000)
$Comborder_{ij}$	Binary variable which is unity if i and j share a land border	CIA's *World Factbook* and Rose (2000)
Landl	Number of landlocked countries in the country-pair (0, 1, or 2)	CIA's *World Factbook* and Rose (2000)
Island	Number of island nations in the country-pair (0, 1, or 2)	CIA's *World Factbook* and Rose (2000)
$Larea_{ij}$	Log of product of land area of i and j	WDI and CIA's *World Factbook*
$Comcol_{ij}$	Binary variable which is unity if i and j were colonies after 1945 with the same colonizer	CIA's *World Factbook* and Rose (2000)
$Curcol_{ij}$	Binary variable which is unity if i and j are colonies at time t	CIA's *World Factbook* and Rose (2000)
$Evercol_{ij}$	Binary variable which is unity if i colonized j or vice versa	CIA's *World Factbook* and Rose (2000)
$Comcty_{ij}$	Binary variable which is unity if i and j remained part of the same nation during the	CIA's *World Factbook* and Rose (2000)
Fta_{ij}	Binary variable which is unity if i and j belong to the same regional trade	WTO(http://rtais.wto.org/UI/PublicMaintainRTAHome.aspx)

Table B2: Summary statistics of selected variables, 1972-2006

Sample	World		Ind-Ind		Ind-Nind		Nind-Nind	
Observations	177,270		9,710		75,169		84,406	
Variable	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Log of real trade	14.66	3.86	19.82	2.24	15.92	3.17	13.09	3.64
Log product real GDP	47.79	2.92	51.85	2.38	48.81	2.50	46.54	2.55
Log product real GDP/capita	15.81	2.14	19.56	0.71	16.97	1.30	14.48	1.73
Currency union	0.01	0.11	0.03	0.18	0.00	0.04	0.02	0.13
De jure direct peg	0.01	0.08	0.01	0.10	0.01	0.12	0.00	0.01
De facto direct peg[1]	0.01	0.10	0.01	0.10	0.02	0.14	0.00	0.02
Long-run exchange rate volatilty	0.16	0.20	0.14	0.23	0.17	0.23	0.16	0.17
Short-run exchange rate volatilty	0.17	0.42	0.16	0.61	0.19	0.56	0.15	0.22
Log distance	8.22	0.79	7.80	1.04	8.35	0.64	8.16	0.85
FTA	0.05	0.21	0.18	0.39	0.01	0.10	0.06	0.24
Log product of areas	23.87	3.31	23.83	3.17	23.68	3.29	24.02	3.32
Number landlocked in the pair	0.32	0.52	0.18	0.40	0.30	0.50	0.34	0.54
Number islands in the pair	0.36	0.55	0.49	0.59	0.41	0.57	0.31	0.52
Common colonizer	0.09	0.29	0.02	0.13	0.04	0.20	0.15	0.35
Ever colony	0.02	0.13	0.04	0.19	0.03	0.18	0.00	0.04
Common land border	0.02	0.15	0.05	0.22	0.00	0.06	0.03	0.18

Table B3. List of countries in the sample

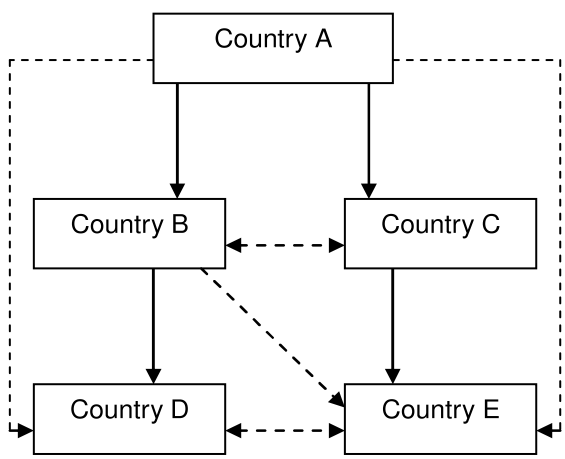
Advanced	***Emerging***		***Developing***			
Australia	Argentina	Russia	Albania	Croatia	Libya	Suriname
Austria	Brazil	Saudi Arabia	Algeria	Dominica	Lithuania	Swaziland
Belgium	Bulgaria	Slovak Republic	Angola	Equatorial Guinea	Macedonia FYR	Syrian Arab R
Canada	Chile	South Africa	Antigua and Barbuda	Estonia	Madagascar	Tajikistan
Hong Kong	China	Sri Lanka	Armenia	Ethiopia	Malawi	Tanzania
Cyprus	Colombia	Thailand	Aruba	Fiji	Mali	Togo
Denmark	Czech Rep.	Tunisia	Azerbaijan Rep. of	Gabon	Malta	Tonga
Finland	Côte d'Ivoire	Turkey	Bahamas The	Gambia The	Mauritania	Trinidad and T
France	Dominican Rep.	Ukraine	Bahrain Kingdom of	Georgia	Mauritius	Uganda
Germany	Ecuador	Uruguay	Bangladesh	Ghana	Moldova	Vanuatu
Greece	Egypt	Venezuela	Belarus	Grenada	Mongolia	Vietnam
Iceland	El Salvador	Zimbabwe	Belize	Guatemala	Mozambique	Zambia
Ireland	Hungary		Benin	Guinea-Bissau	Nepal	
Israel	India		Bhutan	Guyana	Niger	
Italy	Indonesia		Bolivia	Haiti	Papua New Guinea	
Japan	Jordan		Botswana	Honduras	Paraguay	
Luxembourg	Korea		Brunei Darussalam	Iran I.R. of	Rwanda	
Netherlands	Malaysia		Burkina Faso	Jamaica	Samoa	
New Zealand	Mexico		Burundi	Kazakhstan	Senegal	
Norway	Morocco		Cambodia	Kenya	Seychelles	
Portugal	Nigeria		Cameroon	Kuwait	Sierra Leone	
Singapore	Oman		Cape Verde	Kyrgyz Republic	Slovenia	
Spain	Pakistan		Central African Rep.	Lao PDR	Solomon Islands	
Sweden	Panama		Chad	Latvia	St. Kitts and Nevis	
Switzerland	Peru		Macao	Lebanon	St. Lucia	
United Kingdom	Philippines		Congo Republic of	Lesotho	St. Vincent & Grens.	
United States	Poland		Costa Rica	Liberia	Sudan	

Table B4: List of anchor countries and currencies

Anchor country	Anchor currency
France	French franc
Germany	Deutsche Mark
Belgium	Belgian franc
Portugal	Portuguese escudo
Spain	Spanish peseta
Pound	Pound sterling
Australia	Australian dollar
New Zealand	New Zealand dollar
Singapore	Singapore dollar
India	Indian rupee
Russia	Russian ruble
South Africa	South African rand

Source: Anderson (2008).

Figure B1: Direct and indirect peg relations across countries*



A-B=Direct peg (relation=1)
A-C=Direct peg (relation=1)
B-D=Direct peg (relation=1)
C-E=Direct peg (relation=1)

B-C=Indirect peg (relation=2)
A-D=Indirect peg (relation=3)
A-E=Indirect peg (relation=3)
B-E=Indirect peg (relation=4)
C-D=Indirect peg (relation=4)
D-E=Indirect peg (relation=5)

*Adapted from KS (2006).

APPENDIX C

Table C1: Benchmark specification results with short-run volatility for the world, 1972-2006

	De jure classification				De facto classification			
Sample	World	World	World	World	World	World	World	World
Estimation	OLS	CFE	CPFE	HT	OLS	CFE	CPFE	HT
Specification	(1)	(2)	(4)	(5)	(7)	(8)	(10)	(11)
CU	0.514 ***	0.569 ***	0.264 ***	0.283 ***	0.483 ***	0.566 ***	0.238 ***	0.255 ***
	(0.16)	(0.16)	(0.07)	(0.08)	(0.16)	(0.16)	(0.08)	(0.08)
Direct peg	0.341 ***	0.452 ***	0.29 ***	0.285 ***	0.327 ***	0.533 ***	0.311 ***	0.307 ***
	(0.11)	(0.11)	(0.10)	(0.10)	(0.10)	(0.11)	(0.10)	(0.06)
Indirect peg	-0.255 ***	-0.299 ***	-0.153 ***	-0.149 ***	-0.357 ***	-0.329 ***	-0.232 ***	-0.231 ***
	(0.06)	(0.05)	(0.04)	(0.04)	(0.05)	(0.04)	(0.03)	(0.02)
Volatility	-0.27 ***	-0.163 ***	-0.192 ***	-0.188 ***	-0.266 ***	-0.166 ***	-0.193 ***	-0.188 ***
	(0.04)	(0.03)	(0.02)	(0.02)	(0.04)	(0.03)	(0.02)	(0.02)
Lrgdp	1.123 ***	0.51 ***	1.063 ***	1.204 ***	1.123 ***	0.5 ***	1.05 ***	1.188 ***
	(0.11)	(0.08)	(0.07)	(0.05)	(0.01)	(0.08)	(0.07)	(0.02)
Lrgdppc	0.018	0.702 ***	0.197 ***	0.115 *	0.014	0.713 ***	0.212 ***	0.132 ***
	(0.01)	(0.08)	(0.07)	(0.06)	(0.01)	(0.08)	(0.07)	(0.03)
Ldist	-1.226 ***	-1.515 ***		-1.692 ***	-1.23 ***	-1.517 ***		-1.734 ***
	(0.02)	(0.03)		(0.13)	(0.02)	(0.03)		(0.16)
Comlang	0.501 ***	0.501 ***		0.558 ***	0.504 ***	0.506 ***		0.554 ***
	(0.05)	(0.05)		(0.05)	(0.05)	(0.05)		(0.08)
Comborder	0.611 ***	0.398 ***		0.190	0.615 ***	0.391 ***		0.114
	(0.14)	(0.15)		(0.30)	(0.14)	(0.15)		(0.36)
Fta	1.277 ***	0.644 ***	0.254 ***	0.273 ***	1.282 ***	0.651 ***	0.268 ***	0.287 ***
	(0.11)	(0.10)	(0.05)	(0.05)	(0.10)	(0.10)	(0.05)	(0.03)
Landl	-0.314 ***	-0.672 **		-0.144 ***	-0.325 ***	-0.685 **		-0.158 ***
	(0.03)	(0.33)		(0.05)	(0.03)	(0.33)		(0.05)
Island	0.116 ***	-0.289		0.342 ***	0.117 ***	-0.286		0.345 ***
	(0.04)	(0.28)		(0.07)	(0.04)	(0.28)		(0.08)
Lareap	-0.077 ***	0.342 ***		-0.04	-0.078 ***	0.344 ***		-0.030
	(0.01)	(0.05)		(0.03)	(0.01)	(0.05)		(0.02)
Comcol	0.756 ***	0.661 ***		1.112 ***	0.749 ***	0.663 ***		1.106 ***
	(0.08)	(0.07)		(0.10)	(0.08)	(0.07)		(0.11)
Curcol	-0.113	-0.053	-0.365	-0.373	-0.120	-0.048	-0.372	-0.378
	(0.58)	(0.73)	(0.70)	(0.70)	(0.58)	(0.73)	(0.70)	(0.23)
Evercol	1.21 ***	1.343 ***		1.053 ***	1.188 ***	1.321 ***		1.044 ***
	(0.14)	(0.15)		(0.17)	(0.14)	(0.15)		(0.24)
Comctry	1.645 ***	1.192		1.971 ***	1.664 ***	1.197		1.995
	(0.57)	(0.77)		(0.72)	(0.57)	(0.77)		(2.76)
Constant	-26.96 ***	-17.04 ***	-39.44 ***	-29.73 ***	-26.83 ***	-16.8 ***	-39.02 ***	-29.12 ***
	(0.37)	(1.63)	(2.64)	(1.72)	(0.37)	(1.63)	(2.63)	(1.50)
Observations	177,270	177,270	177,270	177,270	177,270	177,270	177,270	177,270
Number of pairs			10,901	10,901			10,901	10,901
R-squared	0.72	0.77			0.72	0.77		
R2-within			0.14				0.14	
R2-between			0.67				0.67	
R2-overall			0.61				0.61	
Hausman chi-2 (FE vs. RE)[1]			0.00				0.00	
Hausman chi-2 (HT vs. RE)[2]				0.00				0.00
Hausman chi-2 (FE vs. HT)[3]				0.95				0.93

Source: Authors' calculations.
Robust clustered standard errors in parentheses; Time effects included in all specifications
Indirect peg refers to relation=2 in Figure A1.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
Volatility refers to short-run volatility computed over 12-month horizon.
Variables instrumented for in HT: CU, Direct peg, Lrgdp, Lrgdppc, Ldist, FTA.
[1] Hausman test applied to the difference between the within (fixed effects) and GLS (random effects) estimators.
[2] Hausman test applied to the difference between the HT estimators and GLS (random effects).
[3] Hausman test applied to the difference between the within (fixed effects) and HT estimators.

Table C2: Benchmark specification results with short-run volatility for sub-samples, 1972-2006

	De jure classification			De facto classification		
Sample	Ind-Ind	Ind-Nind	Nind-Nind	Ind-Ind	Ind-Nind	Nind-Nind
Estimation	HT	HT	HT	HT	HT	HT
Specification	(1)	(2)	(3)	(4)	(5)	(6)
CU	0.324 ***	0.218	0.611 *	0.356 ***	0.231	0.599 ***
	(0.06)	(0.22)	(0.33)	(0.03)	(0.17)	(0.23)
Direct peg	0.194 ***	0.238 **	-0.670 ***	0.221 ***	0.271 ***	0.078
	(0.05)	(0.11)	(0.03)	(0.05)	(0.05)	(1.09)
Indirect peg	0.039	-0.038	-0.168 ***	0.096 ***	-0.108 ***	-0.232 ***
	(0.04)	(0.05)	(0.05)	(0.02)	(0.03)	(0.02)
Volatility	-0.0284	-0.161 ***	-0.199 ***	-0.026	-0.161 ***	-0.201 ***
	(0.02)	(0.02)	(0.04)	(0.02)	(0.02)	(0.03)
Lrgdp	0.419 ***	0.681 ***	1.119 ***	0.447 ***	0.687 ***	1.130 ***
	(0.14)	(0.09)	(0.07)	(0.04)	(0.03)	(0.05)
Lrgdppc	0.808 ***	0.683 ***	0.116	0.770 ***	0.679 ***	0.102 **
	(0.21)	(0.09)	(0.08)	(0.05)	(0.03)	(0.04)
Ldist	-0.385 ***	-0.563 ***	-2.215 ***	-0.393 ***	-0.628 ***	-2.316 ***
	(0.10)	(0.14)	(0.19)	(0.09)	(0.16)	(0.25)
Comlang	0.594 ***	0.421 ***	0.291 ***	0.600 ***	0.432 ***	0.263 **
	(0.18)	(0.09)	(0.09)	(0.20)	(0.12)	(0.13)
Comborder	1.034 ***	1.196 **	-0.072	1.000 ***	1.048	-0.256
	(0.28)	(0.58)	(0.38)	(0.34)	(0.68)	(0.51)
Fta	0.227 ***	0.322 ***	0.060	0.216 ***	0.351 ***	0.060
	(0.04)	(0.06)	(0.10)	(0.01)	(0.05)	(0.07)
Landl	-0.940 ***	-0.374 ***	-0.310 ***	-0.914 ***	-0.374 ***	-0.325 ***
	(0.18)	(0.07)	(0.09)	(0.18)	(0.08)	(0.08)
Island	-0.575 **	-0.765 ***	0.542 ***	-0.530 ***	-0.739 ***	0.582 ***
	(0.27)	(0.13)	(0.08)	(0.16)	(0.11)	(0.11)
Lareap	0.051	0.126 ***	0.058	0.041	0.124 ***	0.056
	(0.05)	(0.04)	(0.05)	(0.03)	(0.02)	(0.04)
Comcol	0.860 *	1.073 ***	1.048 ***	0.844	1.066 ***	1.019 ***
	(0.52)	(0.23)	(0.12)	(0.57)	(0.22)	(0.14)
Curcol	-0.089 ***	-0.528		-0.090	-0.529 ***	
	(0.02)	(0.86)		(0.13)	(0.19)	
Evercol	0.723 ***	1.701 ***	-0.655	0.702 *	1.682 ***	-0.750
	(0.21)	(0.17)	(0.43)	(0.37)	(0.26)	(1.13)
Comctry		1.110			1.118	
		(0.86)			(2.46)	
Constant	-15.790 ***	-27.110 ***	-23.360 ***	-16.240 ***	-26.730 ***	-22.720 ***
	(3.56)	(2.29)	(2.93)	(1.17)	(1.52)	(2.61)
Observations	9,781	75,123	92,366	9,781	75,123	92,366
Number of pairs	351	3,517	7,033	351	3,517	7,033
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	0.67	0.881	0.166	0.723	0.464	0.119

Source: Authors' calculations.
Robust clustered standard errors in parentheses; Time effects included in all specifications.
Indirect peg refers to relation=2 in Figure A1.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
Volatility refers to long-run volatility computed over 36-month horizon.
Variables instrumented for in HT: CU, Direct peg, Lrgdp, Lrgdppc, Ldist, FTA.
[1] Hausman test applied to the difference between the HT estimators and GLS (random effects).
[2] Hausman test applied to the difference between the within (fixed effects) and HT estimators.

Table C3: Benchmark specification with different levels of indirect pegs, 1972-2006

	De jure classification				De facto classification			
	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
CU	0.337***	0.278***	0.188	0.640**	0.287***	0.298***	0.202	0.562*
	(0.08)	(0.06)	(0.22)	(0.31)	(0.08)	(0.06)	(0.23)	(0.32)
Direct peg	0.304***	0.172***	0.252**	-0.652***	0.323***	0.195***	0.282***	0.124
	(0.11)	(0.05)	(0.12)	(0.03)	(0.10)	(0.05)	(0.11)	(0.08)
Indirect peg	-0.112***	0.0142	0.0502	-0.153***	-0.202***	0.0424	-0.0917**	-0.190***
	(0.03)	(0.04)	(0.04)	(0.04)	(0.03)	(0.04)	(0.04)	(0.03)
Volatility[a]	-0.245***	-0.03	-0.132***	-0.340***	-0.239***	-0.0296	-0.131***	-0.328***
	(0.03)	(0.03)	(0.03)	(0.06)	(0.03)	(0.03)	(0.03)	(0.06)
Observations	177,270	9,710	75,169	92,391	177,270	9,710	75,169	92,391
Number of pairid	10,894	3,515	3,515	7,029	10,894	3,515	3,515	7,029
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	0.723	0.821	0.63	0.117	0.966	0.87	0.994	0.271
	De jure classification				**De facto classification**			
	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
CU	0.285***	0.310***	0.216	0.609*	0.238***	0.331***	0.229	0.529
	(0.07)	(0.06)	(0.22)	(0.33)	(0.07)	(0.07)	(0.22)	(0.34)
Direct peg	0.286***	0.184***	0.239**	-0.672***	0.304***	0.205***	0.270***	0.075
	(0.10)	(0.05)	(0.11)	(0.03)	(0.10)	(0.05)	(0.10)	(0.08)
Indirect peg	-0.118***	0.011	0.033	-0.150***	-0.204***	0.044	-0.0972**	-0.192***
	(0.03)	(0.04)	(0.04)	(0.04)	(0.03)	(0.04)	(0.04)	(0.03)
Volatility[b]	-0.188***	-0.029	-0.161***	-0.199***	-0.187***	-0.028	-0.161***	-0.198***
	(0.02)	(0.02)	(0.02)	(0.04)	(0.02)	(0.02)	(0.02)	(0.04)
Observations	177,270	9,781	75,123	92,366	177,270	9,781	75,123	92,366
Number of pairs	10,901	351	3,517	7,033	10,901	351	3,517	7,033
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	1.00	1.00	0.99	0.21	0.98	1.00	0.07	0.31

Source: Authors' calculations.
Estimation results are obtained from the HT method; Robust clustered standard errors in parentheses; Time effects included in all specifications.
Other control variables include Lrgdp, Lrgdppc, Ldist, Fta, Comlang, Comborder, Island, Landl, Lareap, Comcol, Curcol, Evercol, and Comctry.
Indirect peg refers to relation=2, 3, 4, and 5 in Figure A1.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
[a] Refers to long-run volatility over the 36-month horizon.
[b] Refers to short-run volatility over the 12-month horizon.
[1] Hausman test applied to the difference between the GLS (random effects) and HT estimators.
[2] Hausman test applied to the difference between the HT estimators and GLS (random effects).

Table C4: Augmented specification with quadratic volatility, 1972-2006

	De jure classification				De facto classification			
	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
CU	0.341***	0.299***	0.173	0.646**	0.315***	0.325***	0.185	0.635**
	(0.09)	(0.06)	(0.22)	(0.31)	(0.08)	(0.06)	(0.23)	(0.31)
Direct peg	0.299***	0.183***	0.248**	-0.654***	0.323***	0.208***	0.280***	0.149*
	(0.11)	(0.05)	(0.12)	(0.03)	(0.10)	(0.05)	(0.11)	(0.09)
Indirect peg	-0.153***	0.042	-0.03	-0.174***	-0.229***	0.087**	-0.102**	-0.231***
	(0.04)	(0.04)	(0.05)	(0.05)	(0.03)	(0.04)	(0.05)	(0.03)
Volatility[a]	-0.605***	-0.164**	-0.534***	-0.514***	-0.599***	-0.158**	-0.533***	-0.512***
	(0.07)	(0.08)	(0.07)	(0.13)	(0.07)	(0.08)	(0.07)	(0.13)
$(\text{Volatility})^2$	0.296***	0.093**	0.308***	0.164*	0.295***	0.088**	0.307***	0.168*
	(0.04)	(0.05)	(0.05)	(0.09)	(0.04)	(0.04)	(0.05)	(0.09)
Observations	177,270	9,710	75,169	92,391	177,270	9,710	75,169	92,391
Number of pairs	10,894	350	3,515	7,029	10,894	350	3,515	7,029
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	0.43	0.92	0.16	0.09	0.21	1.00	0.01	0.05
	De jure classification				**De facto classification**			
	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
CU	0.277***	0.322***	0.213	0.594***	0.250***	0.354***	0.226	0.581*
	(0.08)	(0.03)	(0.17)	(0.22)	(0.07)	(0.06)	(0.22)	(0.33)
Direct peg	0.282***	0.194***	0.237***	-0.686	0.303***	0.220***	0.269***	0.067
	(0.07)	(0.05)	(0.05)	(1.09)	(0.09)	(0.05)	(0.10)	(0.08)
Indirect peg	-0.149***	0.037**	-0.04	-0.169***	-0.230***	0.094**	-0.108**	-0.233***
	(0.02)	(0.02)	(0.03)	(0.03)	(0.02)	(0.04)	(0.05)	(0.03)
Volatility[b]	-0.326***	-0.116*	-0.233***	-0.347***	-0.329***	-0.098	-0.233***	-0.350***
	(0.03)	(0.06)	(0.04)	(0.06)	(0.05)	(0.06)	(0.05)	(0.08)
$(\text{Volatility})^2$	0.0803***	0.0424	0.039*	0.092***	0.082***	0.035	0.039	0.093**
	(0.02)	(0.03)	(0.02)	(0.03)	(0.02)	(0.03)	(0.03)	(0.04)
Observations	177,270	9,781	75,123	92,366	177,270	9,781	75,123	92,366
Number of pairs	10,901	351	3,517	7,033	10,901	351	3,517	7,033
Hausman chi-2 (HT vs. RE)[1]	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Hausman chi-2 (FE vs. HT)[2]	0.01	1.00	0.44	0.11	0.00	1.00	0.02	0.06

Source: Authors' calculations.
Estimation results are obtained from the HT method; Robust clustered standard errors in parentheses; Time effects included in all specifications.
Other control variables include Lrgdp, Lrgdppc, Ldist, Fta, Comlang, Comborder, Island, Landl, Lareap, Comcol, Curcol, Evercol, and Comctry.
Indirect peg refers to relation=2 in Figure A1.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
[a] Refers to long-run volatility over the 36-month horizon.
[b] Refers to short-run volatility over the 12-month horizon.
[1] Hausman test applied to the difference between the within (fixed effects) and GLS (random effects) estimators.
[2] Hausman test applied to the difference between the HT estimators and GLS (random effects).

Table C5: Results for sensitivity analysis, 1972-2006

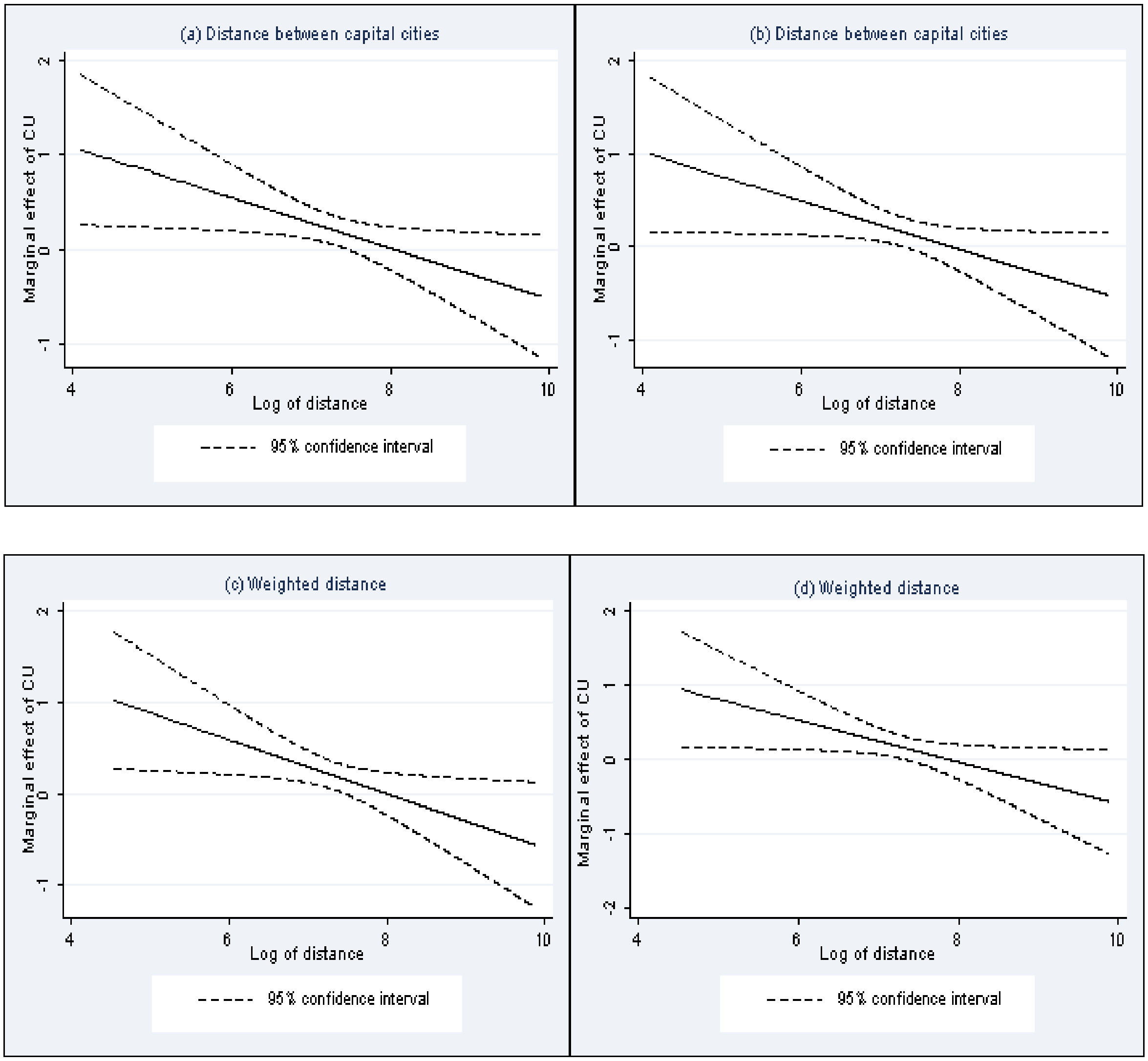
	De jure classification					De facto classification				
Model specification and estimation										
	Log (X_{ij})[1]	Quadratic[2]	Sample selection[3]	Sample selection[4]	PPML[5]	Log (X_{ij})[1]	Quadratic[2]	Sample selection[3]	Sample selection[4]	PPML[5]
CU	0.333***	0.332***	0.316***	0.329***	0.196***	0.302***	0.307***	0.292***	-0.227***	0.088
	(0.08)	(0.07)	(0.08)	(0.08)	(0.06)	(0.08)	(0.07)	(0.08)	(0.03)	(0.09)
Direct peg	0.151	0.369***	0.306***	0.303***	0.091**	0.224***	0.388***	0.329***	-0.241***	-0.147
	(0.10)	(0.11)	(0.110)	(0.110)	(0.04)	(0.08)	(0.10)	(0.10)	(0.03)	(0.14)
Indirect peg	-0.0733**	-0.145***	-0.150***	-0.149***	0.079**	-0.186***	-0.213***	-0.227***	0.304***	0.015
	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.03)	(0.03)	(0.03)	(0.08)	(0.04)
Volatility[a]	-0.205***	-0.256***	-0.250***	-0.246***	-0.164***	-0.202***	-0.251***	-0.246***	0.327***	-0.156***
	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.03)	(0.03)	(0.10)	(0.03)
Observations	177,270	177,270	177,270	177,270	192,906	177,270	177,270	177,270	177,270	192,906
Number of pairs	10,894	10,894	10,894	10,894	10,894	10,894	10,894	10,894	10,894	10,894
Measure of volatility and alternate sub-samples										
	Vol2[6]	Vol3[7]	Excl. small states[8]	Excl. oil exporters[9]		Vol2[6]	Vol3[7]	Excl. small states[8]	Excl. oil exporters[9]	
CU	0.364***	0.275***	0.285***	0.291***		-0.229***	0.249***	0.254***	0.248***	
	(0.08)	(0.07)	(0.08)	(0.09)		(0.03)	(0.07)	(0.08)	(0.09)	
Direct peg	0.306***	0.281***	0.300***	0.278***		0.342***	0.306***	0.338***	0.274***	
	(0.11)	(0.11)	(0.12)	(0.09)		(0.09)	(0.10)	(0.11)	(0.08)	
Indirect peg	-0.149***	-0.150***	-0.133***	-0.125***		0.326***	-0.234***	-0.211***	-0.226***	
	(0.04)	(0.04)	(0.04)	(0.04)		(0.10)	(0.03)	(0.03)	(0.03)	
Volatility[a]	-0.134***	0.0242	-0.281***	-0.265***		-0.132***	0.0278	-0.276***	-0.259***	
	(0.02)	(0.02)	(0.03)	(0.03)		(0.02)	(0.02)	(0.03)	(0.03)	
Observations	177,270	177,270	136,060	111,552		177,270	177,270	136,060	111,552	
Number of pairs	10,891	10,887	7,940	6364		10,891	10,887	7,940	6364	

Source: Authors' calculations.
Estimation results are obtained from the HT method; Robust clustered standard errors in parentheses; Time effects included in all specifications.
Other control variables include Lrgdp, Lrgdppc, Ldist, Fta, Comlang, Comborder, Island, Landl, Lareap, Comcol, Curcol, Evercol, and Comctry.
Indirect peg refers to relation=2 in Figure A1.
***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.
[a] Refers to long-run volatility over the 36-month horizon.
[1] Dependent variable is defined as average of log of exports and imports.
[2] Estimated equation includes quadratic terms for lrgdp and lrgdppc.
[3] Estimated equation includes a variable reflecting the maximum number of years a country-pair is in the sample.
[4] Estimated equation includes a binary variable that is unity if the dyad is present in the sample throughout, and zero otherwise.
[5] Estimates from the Poisson Pseudo Maximum Likelihood approach where the dependent variable is real bilateral trade.
[6] Defines volatility as the standard deviation of the first difference of logarithms of the real exchange rate.
[7] Defines volatility as a linear transformation of the coefficient of variation of real exchange rates.
[8] Countries with a population of less than 1 million are excluded from the sample.
[9] Oil exporters are excluded from the sample.

Figure C1. Sensitivity analysis of the marginal effect of CU

(1) De jure classification

(2) De facto classification



Source: Authors' calculations.