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The Empirics of Exchange Rate Regimes and Trade: Words vs. Deeds

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The Empirics of Exchange Rate Regimes and Trade: Words vs. Deeds[†]

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Abstract

This paper examines the impact of exchange rate regimes on bilateral trade while differentiating the effects of "words" and "deeds". Our findings—based on an extended database for de jure and de facto exchange rate classifications—show that while fixed exchange rate regimes increase trade, there is no systematic difference in the effects of policy announcements versus actions to maintain exchange rate stability. The trade generating effect of more stable exchange rate regimes is however more pronounced when words and actions are aligned, both in the short and long-run. Policy credibility therefore plays an important role in determining the effects of de jure and de facto exchange rate arrangements such that deviations between the two could be costly. In addition, we find evidence that (i) the impact of hard pegs such as currency unions is broadly similar to that of conventional pegs; (ii) the currency union and direct peg effects evolve over time; and (iii) the effects of more stable regimes are heterogeneous across country groups.

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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 and scrutiny of researchers and policy makers with a series of financial crisis in the late 1990s. Most of the ensuing research focused on the influence of exchange rate regimes on economic growth and inflation, but the seminal work of Rose (2000), which investigates the effect of monetary unions on bilateral trade, has generated considerable interest in determining the effects of exchange rate regimes on international trade (see, for example, Klein and Shambaugh, 2006; Adam and Cobham, 2007; and Egger, 2008).

While relevant studies almost unanimously find that exchange rate regimes with lower uncertainty and transaction costs—namely, conventional pegs and currency unions—are significantly more pro-trade than flexible regimes, these analyses focus on de facto exchange rate classifications.[1] This approach is based on the premise that such classifications offer an improved characterization of the exchange rate regime in place since actual outcomes are likely to matter more than policy commitments. In fact, since pervasive differences were highlighted by earlier research—notably, Obstfeld and Rogoff (1995) and Calvo and Reinhart (2002)—between the officially announced exchange rate regimes and those followed in practice, the use of the former in empirical analysis has been significantly reduced.

This paper questions the presumed irrelevance of the de jure classification, and argues that to the extent that central banks' commitments affect market expectations, the de jure exchange rate arrangement—which captures these commitments—may also have an important impact on international trade. Disparate economic agents depend on signals of government's policy intentions, and the official announcement of the exchange rate regime provides one such signal. Thus, for example, the announcement of a stable exchange rate regime could anchor inflation expectations, reduce exchange rate risk and uncertainty, and boost trade, particularly in the short-run.[2] The effect may of course be magnified if the announcement is backed up by actions, lowering actual transaction costs for traders. In the long run, however, official past behavior may also send a signal, and persistent deviations from policy commitments could undermine government credibility and the effect of official declarations. The de jure peg is thus likely to be associated with higher trade if the monetary authority demonstrates a history of following through its words thereby sending a stronger signal of its commitments to market participants.

Viewed in this light, official declarations and actions could be seen as reflecting different aspects of exchange rate stability, and a full understanding of the impact of exchange rate arrangements on cross-border trade activity requires assessing both "words" and "deeds". The objective of this paper is therefore to empirically revisit the relationship between exchange rate regimes and bilateral trade, and systematically investigate if differences exist between the trade generating effects of official announcements on exchange rate policy relative to actual outcomes. While

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[1] For example, Klein and Shambaugh (2006) use the de facto classification developed by Shambaugh (2004), while Adam and Cobham (2007) and Egger (2008) use Reinhart and Rogoff's classification (2004).

[2] In a recent study, Guisinger and Singer (2009) find that the de jure fixed exchange rate regime has an important effect on inflation, with inflation being the most contained in the presence of de jure *and* de facto pegs.

doing so, we also examine the extent to which the alignment between words and deeds matters, and if long-run policy credibility—defined in terms of past deviations of deeds from words—has any influence on the effect of commitments on trade flows.

Importantly, the empirical analysis presented here addresses some of the important econometric concerns—particularly those pertaining to the treatment of omitted variables—raised in previous literature in the context of currency unions and bilateral trade. Applying recent developments in the estimation of bilateral trade flow models, known as gravity models, and focusing on various subsamples in addition to the world sample, we put forward improved quantitative estimates obtained through a range of estimation methods including controlling for dyadic fixed effects (with and without time varying country specific effects), and the Hausman Taylor approach, which permits the estimation of time invariant variables. We also modify our model to allow for the effects of currency unions (CUs) and pegs to evolve over time, and explore the dynamic properties of exchange rate regimes in relation to trade.

In addition, we use a novel dataset of the International Monetary Fund's (IMF) de jure and de facto exchange rate regime classifications compiled by Anderson (2008), which offers two notable advantages. First, it is the only available de facto classification which assesses central bank behavior (in addition to supplementary indicators such as exchange rate movements); and second, it has the same cross-country and time coverage for both de jure and de facto classifications, ensuring that any differences in results are not driven by sample differences.[3] We combine this dataset with recent bilateral trade data covering 159 countries over 1972–2006, which includes the post-European Monetary Union (EMU) period that has not been taken into account by most previous studies analyzing trade and exchange rate regimes.

Our findings suggest that, on average, *both* de jure and de facto pegs promote bilateral trade—through channels in addition to reduced exchange rate volatility. The trade generating effect of policy commitments and actions to maintain exchange rate stability is similar (35-39 percent), but is amplified when they are aligned, that is, words are backed by deeds. Further, we find evidence that (i) while countries belonging to a CU trade significantly more with each other than with comparable countries that do not share a currency, the effect is broadly in the same order of magnitude as direct pegs; (ii) exchange rate stability created with trading partners as a result of pegging to an anchor currency may promote trade, but the effect appears to be conditional on the geographical proximity of the trading partner; (iii) the bilateral trade benefits from CU and direct pegs evolve over time—while CUs have clear anticipatory effects, the effect of direct pegs could be persistent up to, on average, three years after they have been put in place; and (iv) the effect of fixed exchange rate regimes varies across subsamples with, for example, nonindustrialized dyads benefiting relatively more from CUs than the industrialized trading pairs.

This study contributes to the relevant literature in several dimensions. First, while a few studies, notably Klein and Shambaugh (KS, 2006), and Adam and Cobham (2007), investigate the importance of exchange rate regimes for trade, and others examine the macroeconomic implications of policy announcements and actions on exchange rates (for example, Genberg and

[3] Anderson (2008) harmonizes the chronological coverage of the de jure and de facto classifications by extending the former up to 2006, and the latter backwards up to 1972.

Swoboda, 2005; and Guisinger and Singer, 2009), to our best knowledge, this is the first attempt to systematically analyze the significance of words and deeds for international trade. Second, we employ a comprehensive and updated de jure and de facto classification from the same source—thus ensuring a consistent comparison—that also covers recent years and includes the formation of the EMU. Finally, our analysis provides more reliable estimates based on improved econometric specifications and methods that control for potential endogeneity biases.

In what follows, Section II outlines the empirical strategy adopted in the paper, and discusses relevant estimation issues. Section III describes the data in detail. Section IV presents the estimation results and the sensitivity analysis. Section V concludes.

II. Empirical Strategy

A. Analytical Framework

In line with recent literature, we employ the workhorse gravity model of bilateral trade flows to investigate the effect of exchange rate regimes on trade. The gravity model 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 (for example, Anderson, 1979; Bergstrand, 1985; and Deardorff, 1998). In its simple form, the gravity equation can be expressed as:

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

where X_{ij} represents the exports from country i to j, Y is total domestic output, P_i and P_j are the overall price indices in country i and j, respectively, T_{ij} are the iceberg trading costs (such that $X'_T < 0$), and σ is the elasticity of substitution between products ($\sigma > 1$).[4]

Benchmark specification

Traditionally, T_{ij} in equation (1) includes transportation costs that are proxied by geographical attributes (such as bilateral distance, access to sea, and contiguity). In recent years, other factors that may affect trade costs, for example, common language, historical ties, free trade agreements, tariffs, and non-tariff barriers have also been included. To the extent that exchange rate policy choices influence currency conversion costs, exchange rate volatility as well as uncertainty, trading costs would also depend on the exchange rate regime in place, making its inclusion in T_{ij} appropriate.

Thus, to examine the trade effects of exchange rate regimes, we augment the traditional gravity equation, and include variables for hard pegs (CUs), conventional (soft) pegs, and any exchange

[4] 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.

rate links created with trading partners as a consequence of pegging with an anchor currency. Our benchmark specification therefore takes the following form:

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

where X_{ijt} denotes bilateral trade between countries i and j in year t; Z is a vector consisting of traditional time varying and invariant trade determinants;[5] 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 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;[6] *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 (Table A4). 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 two countries that are pegged to different base countries, but their base countries are pegged to the same anchor country.[7] 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 channels other than reduced volatility (such as a reduction in transaction costs, increased transparency, and competition). The construction of the n-horizon real exchange rate volatility measure, *Vol*, follows Ghosh, Gulde, and Wolf (GGW, 2003), and is done 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. Next, we take the average of the absolute values over n months to obtain a measure corresponding to that particular year, given by:

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[5] The time variant variables are: the (log of) product of real gross domestic product (GDP) and real GDP per capita of the trading pair; and a binary variable equal to one if the pair shares a free trade agreement. The time invariant variables are: the (log of) product of land areas of the pair, and the distance between them, whether the countries are landlocked or island, and binary variables equal to one if the pair shares colonial ties, language, and border.

[6] By this definition, two countries (B and C) that are pegged to the same anchor (A) are classified as having an indirect peg with each other. Similarly, if another country, D, is pegged to B, then D would also have indirect pegs with A and C, and so forth. See Figure A1 for a diagrammatic illustration of possible indirect peg relations.

[7] Specifically, the first definition of indirect peg includes relation=2 between pairs (which is equivalent to the "sibling" relationship of KS) in Figure A1. The second definition includes indirect relations=2, 3, 4 and 5.

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

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. We also construct two other measures of volatility to verify the robustness of our results to different definitions of the variable, as follows:

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

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

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. $Vol2$ 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 12 and 36 months to measure short and long-run volatility, respectively. $Vol3$ represents a linear transformation of the coefficient of variation of real exchange rates, and is also computed over the short and long horizons.

Words versus deeds

While estimating equation (2) is important to assess the relative importance of de jure vis-à-vis de facto pegs, and for comparing the trade generating effects of CUs versus conventional pegs, we are also interested in knowing if the alignment of words with deeds has any additional impact on bilateral trade. For this purpose, we consider the matrix of four possible scenarios arising from similarities and discrepancies between de jure and de facto exchange rate arrangements: (i) words match deeds on pegs, that is, both de jure peg and de facto arrangements indicate a peg; (ii) "mirage of fixed", that is, the de jure arrangement is a peg but de facto is a nonpeg; (iii) "fear of float", that is, the de jure arrangement is a nonpeg but de facto is a peg; and (iv) words match deeds on nonpegs, that is, both de jure and de facto arrangements indicate nonpegs. We create binary variables to represent these cases and take the fourth scenario as the reference category to modify equation (2) as follows:

$$\log(X_{ijt}) = \beta_0 + \sum_{k=1}^{N} \beta_k Z_{ijt} + \gamma CU_{ijt} + \delta_1 Deedsmatchwords_{ijt} + \delta_2 Mirageoffixed_{ijt} + \delta_3 Fearoffloat_{ijt} + \varepsilon IndPeg_{ijt} + \zeta Vol_{ijt} + \lambda_t + u_{ijt}, \quad (3)$$

where *Deedsmatchwords*, *Mirageoffixed*, and *Fearoffloat* are dummy variables equal to one if both de jure and de facto classifications indicate a peg; de jure is a peg while de facto is a nonpeg; and de jure is a nonpeg but de facto indicates a peg, respectively, and are equal to zero otherwise.

The extent to which words and deeds matter for each other could be assessed by a comparison of the estimated δ_1, δ_2 and δ_3 from equation (3). If policy commitments matter for de facto exchange rate stabilization, then the de facto peg supported by words should have a larger trade generating effect than the de facto peg not supported by words (that is, $\hat{\delta}_1 > \hat{\delta}_3$). By this account, the estimated δ_1 would also be larger than the estimated δ obtained from equation (2), which represents the average impact of de facto pegs. Similarly, if there exists any costs of deviating from the announced policy to maintain a peg, then the de jure peg not backed by actions would have a smaller effect vis-à-vis the scenario when commitments are kept (that is, $\hat{\delta}_1 > \hat{\delta}_2$).

Nevertheless, as discussed earlier, the consensus between policy commitments and actions may not only be important in the short-run but also in the long-run. By observing exchange rate movements, market participants could detect defections from official proclamations, and the government risks losing credibility (thereby creating greater uncertainty) over time, if it reneges too often on its commitments. Consequently, all else being constant, the effect of words on exchange rate stability is likely to be lower in such cases than if the government maintains a good track record of following through its official commitments.

To investigate the importance of long-run policy credibility, we construct two measures based on the share of mismatches between de jure and de facto classifications over the prior three and five years. These measures take values in the range of 0 and 1 such that if a country does not abide by its committed exchange rate regime in all of the previous three or five years, it receives the score of 1—which indicates weak credibility—while the country which does not defect at all receives the score of zero. We interact our credibility measures with the de jure peg in equation (2) to test whether the impact of words and deeds in time period *t* is determined by past government behavior on exchange rate policy. If weaker policy credibility lowers the effectiveness of signaling, then the estimated coefficient of the interaction term (ψ) is expected to be negative in the equation below:

$$\log(X_{ijt}) = \beta_0 + \sum_{k=1}^{N} \beta_k Z_{ijt} + \gamma CU_{ijt} + \delta DirectPeg_{ijt} + \varepsilon IndPeg_{ijt} + \zeta Vol_{ijt} + \psi DirectPeg_{ijt} * Credibility_{ijt} + \theta Credibility_{ijt} + \lambda_t + u_{ijt}. \quad (4)$$

Dynamic effects

Estimates of exchange rate regimes based on a static specification such as equation (2) ignore the possibility that the trade generating effects of more stable exchange rate arrangements may phase in over time instead of jumping to a new long-run equilibrium as soon as a CU or direct peg is in place. This could be the outcome of, for example, entry and exit decisions of firms in response to official announcements or actions on exchange rate—where the decision taken in the current period may affect output and trade in subsequent time periods. In addition, stable exchange rate regimes could also have anticipatory effects. This is particularly true for CUs where the commitment is typically made a few years in advance, which may lead traders to strengthen links with other member countries and build networks in the run up to CU formation.[8]

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[8] In the context of regional trade agreements, Magee (2007) finds evidence of significant anticipatory effects—on average, up to four years before an RTA is formed.

To take into account the dynamic effects of exchange rate regimes on trade, we extend our benchmark specification in two ways. First, we include binary variables indicating years before adopting a common currency or a conventional peg to capture the anticipatory effects of these regimes. Second, we add lags of CU and direct peg variables to examine any persistence in the impact. The estimated equation thus takes the following form:

$$\log(X_{ijt}) = \beta_0 + \sum_{k=1}^{N} \beta_k Z_{ijt} + \sum_{s=-5}^{5} \gamma_s CU_{ij(t-s)} + \sum_{s=-5}^{5} \delta_s DirPeg_{ij(t-s)} + \varepsilon IndPeg_{ijt} + \varsigma Vol_{ijt} + \lambda_t + u_{ijt}. \qquad (5)$$

Equation (5) measures the impact of CUs and direct pegs five years prior to their adoption and up to five years after they start (with $s = 0$ representing the year of adoption). The cumulative effect of these arrangements is hence given by the sums of estimated γ and δ, respectively.

B. Estimation Issues

Estimation of the gravity model raises several methodological issues that have been discussed extensively in the literature, foremost being the potential endogeneity of regressors, essentially arising from their correlation with the error term u_{ijt} in equations (2)-(4). The two important sources of this endogeneity are omitted variables and reverse causality (or simultaneity). 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.

Omitted variable bias

The omitted variable bias may originate from the correlation of any pro-trade omitted variables with the explanatory variables in the gravity equation. For example, the error term in equation (2) may be representing unobserved political and institutional variables, which affect trade between two countries and are not accounted for in the model, but may also be correlated with the decision to adopt a particular exchange rate regime. 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" terms—the price indices P_i and P_j in equation (1)—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. The CYFE capture any time varying country-specific shocks to trade flows, as well as other factors that are not included in the model due to lack of data or measurement difficulties (for example, infrastructure, factor endowments, and institutions).

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. The CPFE, however, does not provide coefficient estimates for the time invariant variables. 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.

In our analysis, we address the endogeneity concern resulting from the omitted variable bias and 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.[9] To construct instruments, the HT method exploits the panel dimension of the data, and instruments the endogenous time varying variables by the deviation from their individual means, and the endogenous time invariant variables by the deviation of the exogenous time varying variables from their individual means.[10]

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 Serlenga and Shin (2007) use the HT method to examine intra-EU trade during 1960–2001. Both studies find evidence that the HT method is more suitable than the fixed and random effect methods.

Simultaneity bias

Another potential source of endogeneity stems from the fact that the choice of exchange rate regime may not be exogenous, but depend itself 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 studies ignore endogeneity concerns because of the difficulty in finding plausible instruments, but exceptions include Alesina, Barro, and Tenreyro (2002) and Barro and Tenreyro (2007), who exploit client country decisions to peg to an anchor country to construct instruments. Frankel (2008) addresses endogeneity by conducting a “natural

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

[10] Identification requires the number of exogenous time varying variables to be at least as large as the endogenous time invariant variables. The regressors that constitute the set of endogenous variables can be determined by a Hausman test, which is based on the comparison of the HT estimator with the within (fixed effects) estimator.

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. Similarly, KS (2006) use information on the share of pegs to potential reference currencies in neighboring countries to construct their instrument. These studies find that the significantly positive effect of fixed exchange rate regimes remains even after controlling for simultaneity, and in some cases becomes larger in magnitude.

While endogeneity may be an important issue in cross-sectional studies, an advantage of using the panel specification is that it could be addressed through the inclusion of unobserved dyad specific effects. Taking into account the dyad fixed effects captures the impact of all time-invariant factors (such as historical, cultural, political, and geographical ties) that are specific to the trading pair but are likely to have an impact on trade as well as on the choice of exchange rate arrangement between them. This makes the assumption of exogenous exchange rate arrangements—which in this context implies that countries do not base their exchange rate policy choices in response to random shocks to trade—much more plausible.[11] Nevertheless, to address any concerns that the exchange rate regime responds to changes in trade due to time-varying bilateral effects not controlled for in the regression, we also estimate equation (2) using the fixed effects-Generalized Method of Moments (GMM) and the system-GMM estimators in the sensitivity analysis.[12]

Model specification

Finally, several issues relating to misspecification of the gravity model have been discussed extensively in earlier literature (see, for example, Baldwin (2006)). These include those pertaining to the: (i) construction of the dependent variable; (ii) possible nonlinear effect of the income variable on trade; (iii) sample selection bias; (iv) sensitivity of the results to the sample; and (v) the inclusion of zero-trade flows. We attempt to address all these issues in an extensive set of robustness tests, discussed in Section IV.B.

III. 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

[11] In the context of currency unions, Rose (2000) argues that endogeneity is not a relevant concern as "trade considerations seem irrelevant when a country decides whether to join or leave a common currency area."

[12] The fixed effects (with lagged dependent variable) GMM estimator may give biased estimates due to correlation between the error term and the lagged dependent variable, but the systems GMM resolves this inconsistency. We estimate both for comparison purposes.

classification in empirical exchange rate analysis has been significantly reduced.[13] Thereafter, de facto classifications that seek to categorize regimes based on movements in the exchange rate or international reserves have been developed—the best known of which include GGW (2003), Levi-Yeyati and Sturzenegger (LYS, 2003), Reinhart and Rogoff (RR, 2004), and Shambaugh (JS, 2004).[14]

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 however beset with two problems. First, the sources and data coverage underlying the above classifications are different from IMF's de jure classification, making it difficult to judge whether any difference in findings reflect 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 these problems, 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 where 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. The classification is extended backwards for the period 1990–2000 by Bubula and Ötker-Robe (2002), and further backwards up to 1972 by Anderson (2008).

The IMF's de facto classification also has the benefit of being 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 (DF) and de jure (DJ) classifications with the classifications of LYS and RR using a composite measure of similarity when regimes are grouped as fixed, intermediate and floating.[15] The constructed "similarity" index—which is a weighted average of the consensus between the classifications across the three regimes—takes a value between 0 and 1, with a value of 1 indicating perfect similarity of the classifications. Specifically, 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

[13] The "mirage of fixed" and "fear of floating" refer to the facts that some countries that claim to peg do not do so in practice, and those that claim to have a float, intervene heavily to stabilize the exchange rate, respectively.

[14] See Rogoff et al. (2004) and Shambaugh (2004) for a review of various exchange rate regime classifications.

[15] Until the end of 2008, the IMF's classification groups 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. 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.

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 regimes, with the weights being equal to the proportion of pegs, intermediate, and floats in the particular classification.

A comparison of the indices reveals that the IMF's classification has overall greater similarity with 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. The JS classification is not included in the similarity indices as it is available as a binary variable (pegs versus nonpegs) only. To include it in the comparison, we group the other exchange rate regimes (IMF, LYS, and RR) into binary variables, and compute the correlation matrix. The IMF's de facto classification is found to be the closest to the JS classification and the least similar to LYS (Table 1). About 87 percent of the observations in the IMF de facto classification are coded (as pegs or nonpegs) in the same way as in the JS classification, and the overall correlation between the two series is 0.76.[16]

Table 2 compares the distribution of countries across the fixed, intermediate and floating regimes based on the IMF de jure and de facto classifications over the period 1972–2006. Clearly, the classifications are not identical but the similarity has increased over time. For example, in the 1970s, about 44 percent of the country-year observations are coded as de jure pegs and 64 percent as de facto pegs—a difference of about 20 percentage points in the fixed regime classification. However, during the 1990s, this difference dropped to 16 percentage points and to a further 10 percentage points in 2000–06. In recent years, the dissimilarity between the classifications is negligible for the intermediate regime, indicating that the discrepancy in the de jure and de facto fixed classifications stems largely from de jure floaters heavily stabilizing the exchange rates and being identified as de facto pegs.

The temporal comparison of the classifications reported in Table 2 also reveals three other 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" hypothesis 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, which seems to be largely driven by the advanced and emerging countries (Figure 2). In fact, the share of intermediate regimes appears to be broadly stable for the developing economies since the late 1980s under both de jure and de facto classifications. Third, the share of de facto floating regimes is lower than the de jure floats throughout, particularly for the emerging markets and developing economies, providing support for Calvo and Reinhart's (2000) "fear of floating" hypothesis.

B. Data and Summary Statistics

The exchange rate regime classification data described above is available in country-year format. Using the information on anchor currencies—also obtained from Anderson (2008)—we

[16] For binary coding, RR's classification with codes 1-4 is considered as pegs; LYS's classification with code equal to 3 is treated as a peg; and JS's binary classification is used.

construct bilateral binary variables for CUs and direct pegs, and combine them with the annual bilateral trade data obtained from the IMF's *Direction of Trade Statistics*. The binary 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.[17]

The other data required for estimation purposes has been compiled from multiple sources.[18] 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 obtained 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 3 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, 107 show a change in regime (both on and off a peg), with a total of 194 switches in our sample. The number of switches to a de jure peg is 71, while the number of exits is 123, with several country-pairs switching regimes more than once. Based on the de facto classification, 121 country pairs switched regimes 251 times, with 107 switches to a peg and 144 exits from it.

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

[18] See Appendix A for a description of data sources and summary statistics.

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 nonindustrialized 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.[19]

Table 4 presents the distribution of dyads across the various exchange rate arrangements based on the official announcements and actual exchange rate behavior. Of the four possibilities—de jure peg-de facto peg; de jure peg-de facto nonpeg; de jure nonpeg-de facto peg; and de jure nonpeg-de facto nonpeg—the majority of observations fall in the last category, and the least where the central bank announces a conventional peg but does not maintain it (the "mirage of fixed rates" scenario). It is interesting to note that the mean (long-run) exchange rate volatility is higher for the de jure peg-de facto peg case, relative to where the declared regime is a nonpeg, but the country manages its exchange rate (the "fear of floating" scenario). If policy announcements do not matter, the behavior of exchange rates should be broadly similar in both cases. The fact that average exchange rate volatility is almost twice as large in the former case supports the observation of Genberg and Swoboda (2005), and reinforces the argument that words could matter significantly.

IV. Empirical Results

A. Benchmark specification

World sample

The estimation results for equation (2) for the full sample are presented in Table 5. 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.[20] 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 5, respectively. In both cases, for the OLS estimation when only time effects are included along with the other gravity variables, CUs and direct pegs have a significantly positive effect on bilateral trade. The signs and magnitude of the estimated coefficients of the traditional gravity variables are plausible and in line with earlier studies, and a majority of these variables are statistically significant at the 1 percent level. The estimated impact of long-run exchange rate volatility is significantly negative, while indirect pegs are also

[19] 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.

[20] 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.

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 CUs 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 CU 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 5) suggest that the difference between the CPFE and HT estimators is not significant enough to reject the appropriateness of the HT estimator when CU, direct peg, real GDP, real GDP per capita, distance, and free trade agreement are considered as endogenous variables. Hence, the HT specifications reported here take this set of variables as endogenous.[21]

The estimated trade generating effect of CUs 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 CU—on average—increases bilateral trade by about 36-39 percent.[22] 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 CUs. 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.[23]

The estimated impact of exchange rate volatility is strongly negative. The obtained point estimate implies that increasing exchange rate volatility by one standard deviation leads to a

[21] 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.

[22] The effect of CUs or direct pegs may include both the direct effect, and the estimated indirect effect through exchange rate volatility. Following previous literature, we identify the two effects separately and refer to the estimated direct impact only. The direct effect of CU is obtained as $e^{0.33}$-1 = 0.39 and $e^{0.31}$-1 = 0.36 for the de jure and de facto classifications, respectively. Removing the CU, direct and indirect peg variables from the equation makes no different to the coefficient of volatility, while removing volatility has a small effect on the magnitude of regime coefficients but not on their significance.

[23] Our significantly positive estimated coefficient for de jure pegs is in contrast to KS (2006), who—despite considerable similarities between the de jure classification and their de facto classification—find it to be insignificant in their sensitivity analysis. To make sure, this difference is not driven by the longer time dimension of our sample, we restrict the sample to 1972–99, but still obtain the same result.

reduction in bilateral trade by about 5 percent.[24] 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).[25] KS (2006) also report a negative estimated coefficient for indirect pegs, but find it to be insignificant. To investigate this result further, we split the indirect peg variable based on the anchor currencies generating the indirect pegs—that is, instead of a composite indirect peg variable, we include four binary variables in equation (3) indicating the indirect peg generated through the currencies of the United States, United Kingdom, France and Germany.

Table 6 (panel [a]) presents the results obtained from this exercise, which show that indirect pegs generated through all currencies except for the US dollar, increase bilateral trade. While most of the indirect pegs in the sample are generated through the US dollar, it is worth noting that they largely comprise nonindustrialized dyads that are geographically located far apart (for example, those between trading partners in East Asia and Latin America). This is in contrast to the indirect pegs generated through, for example, the Deutsche Mark and the French Franc that are mostly between trading partners in Europe and Africa, respectively, with comparatively smaller distances. Geographical location could play an important role in determining the effect of indirect pegs as the benefits associated with greater exchange rate stability vis-à-vis the partner countries may diminish if other trading costs, such as transportation and information, increase.

To examine if the effect of indirect pegs is indeed conditional on distance, we include an interaction term between indirect pegs and distance in equation (3) and reestimate the model for both de jure and de facto classifications.[26] It is interesting to see in Table 6 (panel [b]) that in both cases the effect of indirect peg turns positive while that of the interaction term is negative, supporting our hypothesis that geographical distance dampens the effect of exchange rate stability created through indirect relationships.[27] However, the indirect peg effect is significant for the de facto peg only, indicating that for indirect relationships, the actual outcome of exchange rate policy is more important.

The results presented here 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 B1-B4). 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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[24] 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.

[25] 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.

[26] Instead of the interaction term with the indirect peg variable, we also include separate interaction terms of distance with indirect peg variables for the reference currencies, but the results remain the same—that is, the effect of the indirect peg generated through the US dollar is no longer significantly negative.

[27] We also test the effect of distance for direct pegs, but do not find it to be statistically significant.

Subsamples

Table 7 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 CUs 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 CUs 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. However, considering the small number of (non-CU) 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 preceding the creation of the EMU, when several countries were pegged to the Deutsche Mark.[28] 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 union. However, our results for the other two subsamples indicate that indirect pegs reduce bilateral trade when at least one country in the dyad is a nonindustrialized country. This result confirms the observation made earlier that the negative indirect peg effect for the world sample is likely driven by nonindustrialized countries that are geographically located far apart. Repeating the same exercise as above, and including an interaction term for indirect pegs and distance, shows that the indirect peg effect for nonindustrialized countries is indeed dependent on distance, with dyads located farther away reaping lower benefits from exchange rate stability (Table 6, panel [b]).

The findings reported in Tables 6 and 7 also support the view that there may be systematic differences in country characteristics that imply a varying effect of regimes on bilateral trade across countries groups. This underscores the importance of going beyond one large sample comprising heterogeneous countries in order to draw specific inferences about the impact of exchange rate regimes on trade, as suggested by Baldwin (2006).

B. Words versus deeds

The results presented through Tables 5-7 suggest that, on average, de facto pegs have a slightly higher effect than de jure pegs. To investigate if this effect is augmented when words coincide with actual policy behavior, we estimate equation (3) and report the results in Table 8. Clearly, the positive effect of direct pegs on trade is the largest when deeds match words as in column (1). Interestingly, the estimated coefficient for direct pegs is also larger than those reported for the de jure and de facto pegs in Table 5, 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 nonpeg—the fear of float scenario—the estimated coefficient of the binary variable reflecting direct pegs is positive, but smaller than

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[28] Restricting the sample to 1972–99 and using CYFE (as in KS, 2006), the effect of indirect pegs is insignificant.

when deeds are aligned with words. In contrast, 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 statistically weak.

Looking across the subsamples, we find that the strong effect of words complemented by deeds also holds for the Ind-Ind and Ind-Nind dyads. In both cases, the impact of direct pegs are larger than those reported in Table 7. For the scenarios when deeds do not match words, it is difficult to draw meaningful inferences for the individual subsamples since the number of observations coded as a de jure peg and a de facto nonpeg, or, a de jure nonpeg and a de facto peg, is relatively small across the three groups.

The results for policy credibility, reported in Table 9, indicate that the effectiveness of direct pegs, particularly, de jure pegs, depends not only on commitments maintained in the short-run but also on the reputation of meeting policy announcements in the long-run. The interaction term between de jure direct peg and the credibility variable is significantly negative, suggesting that deviations from words in preceding (five) years are costly for bilateral trade and reduce the positive impact of official proclamations on maintaining exchange rate stability.[29] Thus, for example, the trade generating effect of de jure pegs is estimated to be, on average, about 18 percentage points lower for a country that defected once in the previous five years (credibility measure=0.2) vis-à-vis a country with an unsullied reputation for keeping its words (credibility measure=0). This effect is the most pronounced for nonindustrial countries where weaker credibility appears to substantially dilute the benefits achieved from pegging.

C. Dynamic effects

The estimation results for the dynamic specification outlined in equation (5) are presented in Table 10. The CU and direct peg coefficients are jointly significant, but most are insignificant individually because of collinearity between the variables. The collinearity is however not an issue when the interest primarily lies in the cumulative effect of variables (Magee, 2007). Although we estimate different specifications for equation (5), including anticipatory and lagged effects up to five years for both CUs and direct pegs, their total impact on bilateral trade appears to occur by 3 years as there is no significant change in the cumulative effect of more stable regimes beyond that point.

The cumulative effect for both CU and direct pegs, presented in the last rows of Table 10, is larger than the average effect in the static specification. Specifically, the cumulative effect is in the range of 54-58 percent for CUs, and about 45 and 60 percent for de jure and de facto pegs, respectively. CUs have a clear anticipatory effect on trade flows—there is a significant increase in trade during the three years leading up to the CU—which is much larger in magnitude than for direct pegs. This result makes intuitive sense since, as mentioned earlier, the formation of a common currency area is announced in advance, enabling market participants to react in

[29] The results are similar if we use the credibility measure based on behavior in the past three years, and are not reported here for brevity.

anticipation.[30] It should be noted, however, that in our sample this result is likely driven by EMU creation since the other major CUs (for example, the CFA Franc zone) were created much before the start of the sample and their anticipatory effects are not captured in the estimation.

The estimation also indicates some persistence in the trade creating effects of stable regimes, which is, on average, larger for de facto pegs than de jure pegs. This suggests that the actual exchange rate behavior affects bilateral trade in the current and subsequent periods, with the full impact being realized after a few years. The phasing in of the trade generating effect of direct pegs is also supported by an alternate dynamic specification where we include a variable representing the inverse of the years since adopting a peg. This variable is equal to one for the first year of the peg, 0.5 for the second year, and so forth, and asymptotically approaches zero. The estimated coefficient for this variable is significantly negative and, when combined with the estimate for direct pegs, shows that the effect of the latter in the first year is negligible. However, as the duration of the peg increases, the trade generating effect also increases—becoming about 34 and 39 percent in the fifth and tenth years, respectively.

D. Sensitivity Analysis

The results presented above verify the robustness of our estimates to various estimation methods (for example, OLS, CFE, CYFE, CPFE and HT). However, several other concerns pertaining to model specification, methodology, variable definitions, and sample coverage raised in earlier literature may be relevant to our analysis. In what follows, we attempt to address these concerns through a battery of sensitivity checks.

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 B5).

Third, we control for the possible sample selection bias in gravity model 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

[30] While in principle, significantly positive estimated coefficients for variables indicating pre-CU years could be interpreted as a sign of potential endogeneity, we believe that in this context, they simply reflect anticipatory effects (rather than the formation of a CU in response to higher trade in the few years prior to common currency adoption).

variable equal to unity if the pair is observed during the entire period, and zero otherwise.[31] Once again, the results are similar to those reported earlier, and the estimated coefficients of the variables of interest retain their signs and significance.

In a recent paper, Baier and Bergstand (2007) note that in a panel setting, the multilateral price variables (P^i and P^j) are likely to be time varying. Controlling for them through CPFE may not fully account for the omitted variable problem hence they therefore propose to include both the CPFE as well as the CYFE to control for possible correlation between the unobserved omitted and time invariant bilateral variables, and between the omitted and time-variant variables, respectively. Next, we estimate the benchmark specification controlling for both the dyadic effects and the country year fixed effects, and denote this estimation approach as country-year and county-pair fixed effects (CYPFE). The results support the earlier findings and we obtain a similar trade generating effect for both de jure and de facto direct pegs. However, the size of the estimated effect is almost half of that obtained in Table 5, and smaller than the CU effect.

Finally, we address the issue of zero-trade observations that commonly arises in bilateral datasets either because some dyads do not trade, or because of rounding errors and missing observations. Using the log-linear form of the gravity equation as in equation (3) 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 OLS is used. We check the sensitivity of our results to the inclusion of zero-trade observations by applying the Poisson pseudo maximum likelihood (PPML) approach proposed by Santos Silva and Tenreyro (2006). The PPML estimator provides consistent estimates for the gravity model in the presence of heteroskedastic errors, and takes the real value of trade as the dependent variable thereby allowing countries with no trade to be included in the analysis. The results obtained from this approach—presented in Table B5—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.

Simultaneity bias

The results for GMM estimation, which address potential endogeneity concerns, are presented in Table B5. We first estimate the static fixed effects model with GMM using the lagged exchange rate arrangements as instruments, and then include the lagged dependent variable as a regressor to estimate with the FE-GMM and system-GMM approaches. The system-GMM estimator—proposed by Blundell and Bond (1998)—transforms the model by taking first differences to eliminate the fixed effects and supplements it with the levels equation, using lagged levels as instruments for the former equation and lagged differences as instruments for the latter.

The reported p-values for the Hansen test statistic indicate acceptance of the null hypothesis that the over identifying restrictions are valid for both the levels and differences equations. The p-value for the second order serial correlation test-statistic supports the absence of such correlation, as is required to obtain consistent estimates from system-GMM. The estimation results show that

[31] 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.

the strong positive impact of fixed exchange rate arrangements on bilateral trade is robust to GMM estimation and the dynamic panel specification of the model. In all cases, the coefficients for CU and direct pegs are significantly positive, but their magnitude is relatively larger—and closer to what was obtained earlier through the HT method—for the system-GMM estimator.

Measures of exchange rate volatility and distance

Using the other two volatility measures—*Vol*2 and *Vol*3—described earlier does not alter the main set of results. Like before, fixed exchange rate regimes are found to promote bilateral trade, while, on average, exchange rate volatility has 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 (3) 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 CUs or direct pegs.

Alternate subsamples

To make sure that our results are not being driven by economies with certain characteristics, or because of the inclusion of a particular time period in our sample, we estimate the benchmark specification for three alternate subsamples: (i) non-oil exporting economies; (ii) countries with a population of over 1 million; and (iii) exlcuding the immediate post-Bretton Woods period (1975–2006). We drop oil exporting economies from the sample as determinants of oil trade are different from other trade flows.[32] 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. The results for the sample restricting the time dimension to the 1975–2006 period are also close to those obtained earlier, and show a strongly positive effect of fixed exchange rate regimes on trade.

Trade stability

In addition to the impact on the level of bilateral trade, we also examine if the (de jure and de facto) exchange rate regimes associated with lower exchange rate volatility reduce trade instability. To this end, we follow Rose (2005) and estimate an equation similar to the

[32] 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.

benchmark specification but with the coefficient of variation of real bilateral trade as the dependent variable. We calculate the dependent variable over a (non-overlapping) five-year period, which gives us five observations per pair for the sample, while all the explanatory variables are averaged over the corresponding time period.

As above, we estimate this equation using the OLS, CFE, CYFE, CPFE and HT, but present the results for the HT method in Table B6. 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, CU 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).

V. Conclusion and Policy Implications

This paper provides additional insights into the impact of exchange rate arrangements on bilateral trade while arguing that actual policy behavior as well as commitments to a stable nominal anchor could matter. Using an extended dataset on IMF's de jure and de facto classifications, and addressing potential endogeneity concerns, we find support for this hypothesis and show that both de jure and de facto fixed exchange rate regimes strongly promote bilateral trade. The impact is however more pronounced when words are complemented by deeds, indicating that increased transparency in regime choice anchors expectations and improves credibility, amplifying the effect of fixed regimes on trade outcomes. These results are robust to different model specifications, variable definitions, and econometric methodologies.

The results also reveal that the impact of fixed exchange rate regimes differ across subsamples, 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. In addition, we find evidence that the effect of indirect pegs—generated through relations as a result of pegging with an anchor currency—is conditional on the distance between trading partners such that countries, particularly nonindustrialized countries, located geographically close to each other benefit the most. This finding suggests that lowering transportation and information costs through, for example, improved infrastructure, could be important elements to increase trade gains from maintaining stable exchange rates.

Further, the bilateral trade benefits from CU and direct pegs appear to evolve over time. CUs have clear anticipatory effects with bilateral trade rising between member countries in the years immediately preceding the adoption of the common currency, while the effect of direct pegs could be persistent up to, on average, three years after they have been put in place. The cumulative trade generating effects of CUs and conventional pegs are however broadly similar. Thus, notwithstanding the costs of fixed exchange rate arrangements, countries aspiring to expand cross-border trade activity through more stable exchange rate regimes while retaining some flexibility could opt for conventional pegs rather than complete monetary integration.

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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).

Table 2: 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.

Table 3. 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	100.0%	52.6%	40.1%	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	194	28	165	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	251	28	221	2
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	1,671
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	70	59	1	10

Source: Authors' calculations.

1 Indicates the sample used in the baseline estimations (where outliers are excluded).

Sum of pairs for Ind-Ind, Nind-Nind and Nind-Ind equals the total (world) sample.

Direct pegs exclude pairs which are members of the same currency union.

Table 4: Distribution of dyads across exchange rate arrangements, 1972-2006

		De facto classification: Peg	De facto classification: Non-peg
De jure classification	Peg	**Words match deeds** Observations: 1,148 Exchange rate volatility: Mean: 0.11 Min: 0.01, Max: 1.59 (Log of) Real bilateral trade: Mean: 19.61	**Mirage of fixed** Observations: 44 Exchange rate volatility: Mean: 0.08 Min: 0.01, Max: 1.59 (Log of) Real bilateral trade: Mean: 19.35
	Non-peg	**Fear of float** Observations: 477 Exchange rate volatility: Mean: 0.06 Min: 0.01, Max: 1.04 (Log of) Real bilateral trade: Mean: 20.39	**Words match deeds** Observations: 175,601 Exchange rate volatility: Mean: 0.17 Min: 0.01, Max: 1.81 (Log of) Real bilateral trade: Mean: 14.61

Source: Authors' calculations.

Table 5. 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 (by dyad) 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 6. Results for indirect pegs based on anchor currencies, 1972-2006

(a) Indirect pegs for anchor currencies

	De jure classification				De facto classification			
Sample	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
Specification	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CU	0.372***	0.297***	0.190	0.656**	0.424***	0.330***	0.189	0.647**
	(0.08)	(0.06)	(0.22)	(0.31)	(0.08)	(0.06)	(0.22)	(0.31)
Direct peg	0.305***	0.177***	0.250**	-0.647***	0.328***	0.191***	0.280***	0.132
	(0.11)	(0.05)	(0.12)	(0.03)	(0.10)	(0.05)	(0.11)	(0.09)
Indirect peg (US Dollar)	-0.195***	-0.026	-0.060	-0.192***	-0.283***	-0.139*	-0.228***	-0.247***
	(0.04)	(0.08)	(0.06)	(0.05)	(0.03)	(0.08)	(0.06)	(0.04)
Indirect peg (GBP Sterling)	0.873***		0.535***	1.167***	0.826***		0.515**	1.098***
	(0.27)		(0.20)	(0.41)	(0.26)		(0.20)	(0.40)
Indirect peg (French Franc)	0.981***			1.107***	0.963***			1.105***
	(0.24)			(0.25)	(0.24)			(0.25)
Indirect peg (Deutsche Mark)	0.030	0.055	-0.013	-0.075	0.173***	0.119***	0.191**	0.111
	(0.04)	(0.04)	(0.07)	(0.22)	(0.05)	(0.04)	(0.09)	(0.15)
Volatility[a]	-0.245***	-0.031	-0.130***	-0.339***	-0.237***	-0.033	-0.131***	-0.331***
	(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	350	3,515	7,029	10,894	350	3,515	7,029
Hausman chi-2 (FE vs. HT)[1]	0.834	0.938	0.401	0.140	0.645	1.000	0.023	0.112

(b) Augmented specification w ith interaction term

	De jure classification				De facto classification			
Sample	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
Specification	(5)	(6)	(7)	(8)	(5)	(6)	(7)	(8)
CU	0.347***	0.305***	0.188	0.635**	0.372***	0.332***	0.191	0.627**
	(0.08)	(0.06)	(0.22)	(0.31)	(0.08)	(0.06)	(0.22)	(0.31)
Direct peg	0.302***	0.172***	0.250**	-0.649***	0.326***	0.198***	0.280***	0.127
	(0.11)	(0.05)	(0.12)	(0.03)	(0.10)	(0.05)	(0.11)	(0.09)
Indirect peg	0.270	0.940***	0.525	0.273	1.139***	1.097***	1.632***	0.959***
	(0.26)	(0.33)	(0.35)	(0.34)	(0.22)	(0.32)	(0.40)	(0.27)
Indirect peg*Ldist	-0.050	-0.120***	-0.064	-0.053	-0.160***	-0.135***	-0.203***	-0.138***
	(0.03)	(0.04)	(0.04)	(0.04)	(0.03)	(0.04)	(0.05)	(0.03)
Volatility[a]	-0.243***	-0.033	-0.130***	-0.334***	-0.237***	-0.032	-0.129***	-0.329***
	(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	350	3,515	7,029	10,894	350	3,515	7,029
Hausman chi-2 (FE vs. HT)[1]	0.155	0.997	0.366	0.152	0.209	1.000	0.018	0.112

Source: Authors' calculations.

Estimation results are obtained from the HT method; robust clustered standard errors in parentheses; time effects included in all specifi

Other control variables include Lrgdp, Lrgdppc, Ldist, Fta, Comlang, Comborder, Island, Landl, Lareap, Comcol, Curcol, Evercol, and C

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] Hausman test applied to the difference between the within (fixed effects) and HT estimators.

Table 7. Benchmark specification results for the subsamples, 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 (by dyad) 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 8. Results for deeds versus words for the world and subsamples, 1972-2006

Sample	World	Ind-Ind	Ind-Nind	Nind-Nind
Estimation	HT	HT	HT	HT
Specification	(1)	(2)	(3)	(4)
CU	0.340***	0.294***	0.207	0.586*
	(0.08)	(0.06)	(0.22)	(0.32)
Words match deeds[a]	0.392***	0.189***	0.340***	-0.420***
	(0.12)	(0.05)	(0.13)	(0.03)
Mirage of fixed[b]	-0.166	0.275***	-0.302	-0.918***
	(0.22)	(0.04)	(0.24)	(0.04)
Fear of float[c]	0.223**	0.209***	0.181*	409.1
	(0.10)	(0.06)	(0.10)	(412.10)
Indirect peg	-0.148***	0.043	-0.027	-0.166***
	(0.04)	(0.04)	(0.05)	(0.05)
Volatility[d]	-0.245***	-0.031	-0.131***	-0.335***
	(0.03)	(0.03)	(0.03)	(0.06)
Lrgdp	1.192***	0.404***	0.664***	0.788***
	(0.05)	(0.14)	(0.09)	(0.13)
Lrgdppc	0.128**	0.832***	0.699***	0.360***
	(0.06)	(0.21)	(0.09)	(0.13)
Ldist	-1.810***	-0.161	-0.606***	-2.642***
	(0.13)	(0.16)	(0.15)	(0.26)
Comlang	0.547***	0.444**	0.429***	0.375***
	(0.05)	(0.20)	(0.09)	(0.11)
Comborder	-0.0173	1.452***	1.088*	-4.357***
	(0.31)	(0.39)	(0.60)	(1.07)
Fta	0.269***	0.228***	0.318***	0.0278
	(0.05)	(0.04)	(0.07)	(0.10)
Landl	-0.155***	-0.936***	-0.393***	-0.738***
	(0.05)	(0.19)	(0.07)	(0.15)
Island	0.373***	-0.784***	-0.765***	0.482***
	(0.07)	(0.30)	(0.13)	(0.09)
Lareap	-0.028	0.044	0.137***	0.271***
	(0.03)	(0.05)	(0.04)	(0.09)
Comcol	1.077***	0.951*	1.080***	1.014***
	(0.10)	(0.53)	(0.23)	(0.14)
Curcol	-0.38	-0.085***	-0.536	
	(0.70)	(0.02)	(0.86)	
Evercol	1.060***	0.794***	1.700***	1.861**
	(0.17)	(0.21)	(0.167)	(0.892)
Comctry	2.016***		1.128	
	(0.72)		(0.867)	
Constant	-28.62***	-17.09***	-26.41***	-12.58***
	(1.71)	(3.35)	(2.277)	(4.626)
Observations	177,270	9,710	75,169	92,391
Number of pairid	10,894	350	3,515	7,029
Hausman chi-2 (FE vs. HT)[1]	0.89	1.00	0.73	1.00

Source: Authors' calculations.

Robust clustered standard errors (by dyads) 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] Binary variable is unity if both de jure and de facto are pegs; zero otherwise.

[b] Binary variable 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.

Table 9. Results for long-run policy credibility for the world and subsamples, 1972-2006

	De jure classification				De facto classification			
Sample	World	Ind-Ind	Ind-Nind	Nind-Nind	World	Ind-Ind	Ind-Nind	Nind-Nind
Specificaion	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CU	0.146**	0.193***	-0.055	0.255	0.109*	0.221***	0.255	-0.054
	(0.07)	(0.06)	(0.24)	(0.30)	(0.07)	(0.07)	(0.31)	(0.24)
Direct peg	0.362*	0.873	0.255	-0.165	0.323*	0.486***	0.489***	0.231
	(0.20)	(0.98)	(0.19)	(0.28)	(0.17)	(0.08)	(0.05)	(0.16)
Indirect peg	-0.145***	0.046	0.098	-0.149**	-0.200***	0.0818**	-0.183***	0.040
	(0.05)	(0.04)	(0.09)	(0.06)	(0.03)	(0.04)	(0.04)	(0.09)
Credibility[a]	0.232	0.355**	0.161	-0.862*	0.206	0.302*	-0.505***	0.080
	(0.15)	(0.18)	(0.15)	(0.45)	(0.18)	(0.18)	(0.17)	(0.19)
Direct peg*Credibility	-0.674**	-1.307	-0.563*	-0.146	-0.361	-0.680***	-0.956**	-0.180
	(0.30)	(1.38)	(0.30)	(0.48)	(0.23)	(0.21)	(0.49)	(0.24)
Volatility[b]	-0.250***	-0.041	-0.134***	-0.380***	-0.250***	-0.040	-0.383***	-0.134***
	(0.03)	(0.03)	(0.03)	(0.07)	(0.03)	(0.03)	(0.07)	(0.03)
Lrgdp	1.317***	0.282*	0.494***	1.342***	1.305***	0.295*	1.335***	0.514***
	(0.07)	(0.16)	(0.11)	(0.09)	(0.07)	(0.16)	(0.10)	(0.11)
Lrgdppc	0.003	0.956***	0.823***	-0.099	0.018	0.939***	-0.103	0.806***
	(0.08)	(0.23)	(0.11)	(0.11)	(0.08)	(0.23)	(0.11)	(0.11)
Ldist	-1.780***	-0.062	-0.360**	-2.970***	-1.775***	-0.052	-3.275***	-0.482***
	(0.14)	(0.19)	(0.17)	(0.21)	(0.14)	(0.19)	(0.28)	(0.16)
Comlang	0.584***	0.374	0.333***	0.151	0.585***	0.375	0.058	0.360***
	(0.05)	(0.24)	(0.11)	(0.11)	(0.05)	(0.24)	(0.13)	(0.11)
Comborder	0.014	1.678***	1.795***	-1.464***	0.029	1.690***	-2.023***	1.503**
	(0.32)	(0.46)	(0.68)	(0.42)	(0.32)	(0.46)	(0.55)	(0.64)
Fta	0.305***	0.185***	0.266***	0.097	0.316***	0.180***	0.101	0.277***
	(0.05)	(0.04)	(0.07)	(0.13)	(0.05)	(0.04)	(0.13)	(0.07)
Landl	-0.140***	-0.896***	-0.434***	-0.428***	-0.147***	-0.889***	-0.486***	-0.431***
	(0.05)	(0.20)	(0.08)	(0.10)	(0.05)	(0.20)	(0.12)	(0.08)
Island	0.409***	-1.042***	-0.913***	0.692***	0.397***	-1.034***	0.777***	-0.853***
	(0.09)	(0.36)	(0.16)	(0.10)	(0.09)	(0.36)	(0.11)	(0.15)
Lareap	-0.0968***	0.075	0.198***	-0.018	-0.0906**	0.070	0.002	0.192***
	(0.04)	(0.06)	(0.05)	(0.06)	(0.04)	(0.06)	(0.07)	(0.05)
Comcol	1.089***	0.989	1.245***	0.993***	1.097***	0.981	0.911***	1.221***
	(0.12)	(0.61)	(0.27)	(0.15)	(0.12)	(0.61)	(0.16)	(0.26)
Curcol	-0.823	-0.0944***	-1.816		-0.828	-0.0929***		-1.814
	(1.00)	(0.02)	(1.57)		(1.00)	(0.02)		(1.57)
Evercol	0.862***	0.883***	1.779***	-1.610***	0.860***	0.850***	-1.848***	1.750***
	(0.17)	(0.26)	(0.18)	(0.61)	(0.17)	(0.26)	(0.68)	(0.18)
Constant	-31.27***	-14.56***	-23.63***	-22.19***	-31.10***	-14.85***	-19.59***	-23.14***
	(1.52)	(3.71)	(2.46)	(2.86)	(1.68)	(3.70)	(3.80)	(2.48)
Observations	126,728	7,966	57,911	60,851	126,728	7,966	60,851	57,911
Number of pairid	9,194	344	3,378	5,472	9,194	344	5,472	3,378
Hausman chi-2 (FE vs. HT)[1]	1.00	1.00	1.00	0.11	1.00	1.00	0.22	0.73

Source: Authors' calculations.

Robust clustered (by dyad) 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] Credibility reflects the share of mismatches between de jure and de facto classifications over the prior five years (with 0 and 1 reflecting strongest and weakest credibility, respectively).

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

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

Table 10. Results for dynamic specification for the world sample, 1972-2006

	De jure classification		De facto classification	
Specification	(1)	(2)	(3)	(4)
CU$_{(t+3)}$	0.317***		0.314***	
	(0.07)		(0.07)	
CU$_{(t+2)}$	-0.067		-0.096	
	(0.07)		(0.07)	
CU$_{(t+1)}$	0.004		0.028	
	(0.07)		(0.07)	
CU	0.347***	0.301***	0.317***	0.282***
	(0.08)	(0.10)	(0.08)	(0.10)
CU$_{(t-1)}$	-0.011		0.000	
	(0.07)		(0.07)	
CU$_{(t-2)}$	-0.110		-0.107	
	(0.09)		(0.09)	
CU$_{(t-3)}$	-0.018		-0.028	
	(0.12)		(0.11)	
Direct peg$_{(t+3)}$	0.027		0.043	
	(0.06)		(0.09)	
Direct peg$_{(t+2)}$	0.051		0.129**	
	(0.06)		(0.06)	
Direct peg$_{(t+1)}$	-0.034		-0.096*	
	(0.07)		(0.05)	
Direct peg	0.124	0.360***	0.134	0.414***
	(0.09)	(0.13)	(0.08)	(0.12)
Direct peg$_{(t-1)}$	0.049		0.081*	
	(0.06)		(0.05)	
Direct peg$_{(t-2)}$	-0.013		0.040	
	(0.04)		(0.03)	
Direct peg$_{(t-3)}$	0.163***		0.148***	
	(0.05)		(0.05)	
1/ (No. of years in CU)		-0.058		-0.059
		(0.14)		(0.14)
1/ (No. of years direct peg)		-0.321**		-0.424***
		(0.14)		(0.13)
Indirect peg	-0.173***	-0.146***	-0.219***	
	(0.04)	(0.04)	(0.03)	
Volatility[a]	-0.248***	-0.243***	-0.246***	-0.238***
	(0.03)	(0.03)	(0.03)	(0.03)
Observations	131,839	177,270	131,839	177,270
Number of pairid	9,208	10,894	9,208	10,894
Cumulative CU effect[b]	0.46		0.43	
Anticipatory CU effect[c]	0.25		0.25	
Post CU effect[d]	0.21		0.18	
Cumulative direct peg effect[e]	0.37		0.48	
Anticipatory direct peg effect[f]	0.05		0.08	
Post direct peg effect[g]	0.32		0.40	
Hausman chi-2 (FE vs. HT)[1]	0.21	0.75	0.10	0.51

Source: Authors' calculations.

Estimation results are from the HT method; robust clustered (by dyads) standard errors in parentheses; time effects included in all specifications.

***, ** and * indicate significance at the 1%, 5% and 10% significance levels, respectively.

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.

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

[b] Sum of CU$_{t+3}$, CU$_{t+2}$, CU$_{t+1}$, CU$_{t-1}$, CU$_{t-2}$, CU$_{t-3}$.

[c] Sum of CU$_{t+3}$, CU$_{t+2}$, CU$_{t+1}$.

[d] Sum of CU$_{t-1}$, CU$_{t-2}$, CU$_{t-3}$.

[e] Sum of Direct peg$_{t+3}$, Direct peg$_{t+2}$, Direct pegt$_{+1}$, Direct peg$_{t-1}$, Direct peg$_{t-2}$, Direct peg$_{t-3}$.

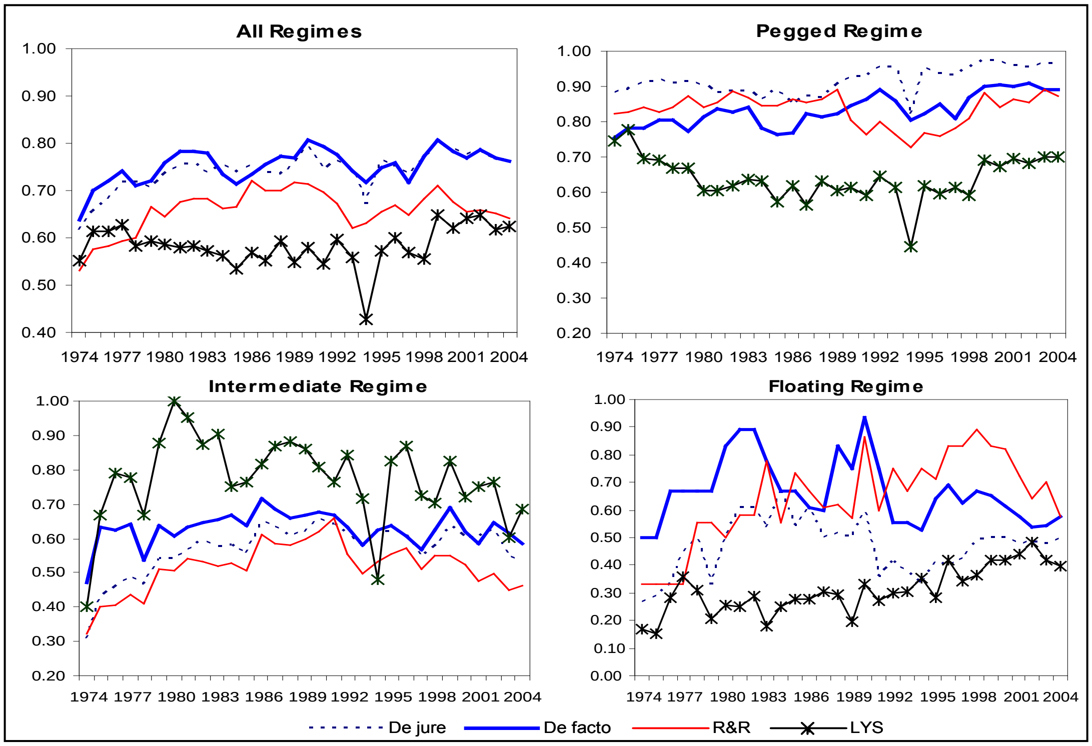
[f] Sum of Direct peg$_{t+3}$, Direct peg$_{t+2}$, Direct pegt$_{+1}$.

[g] Sum of Direct peg$_{t-1}$, Direct peg$_{t-2}$, Direct peg$_{t-3}$.

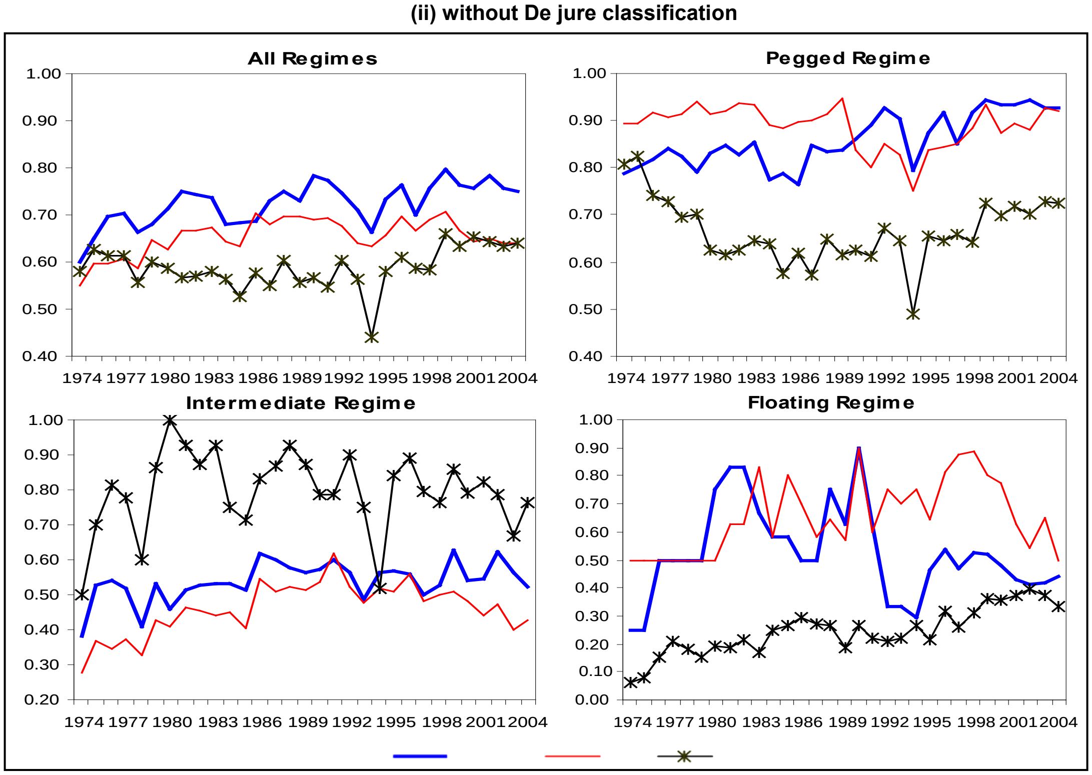
[1] Hausman test applied to the difference between the within (fixed effects) and HT estimators.

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

(i) with De jure classification



(ii) without De jure classification

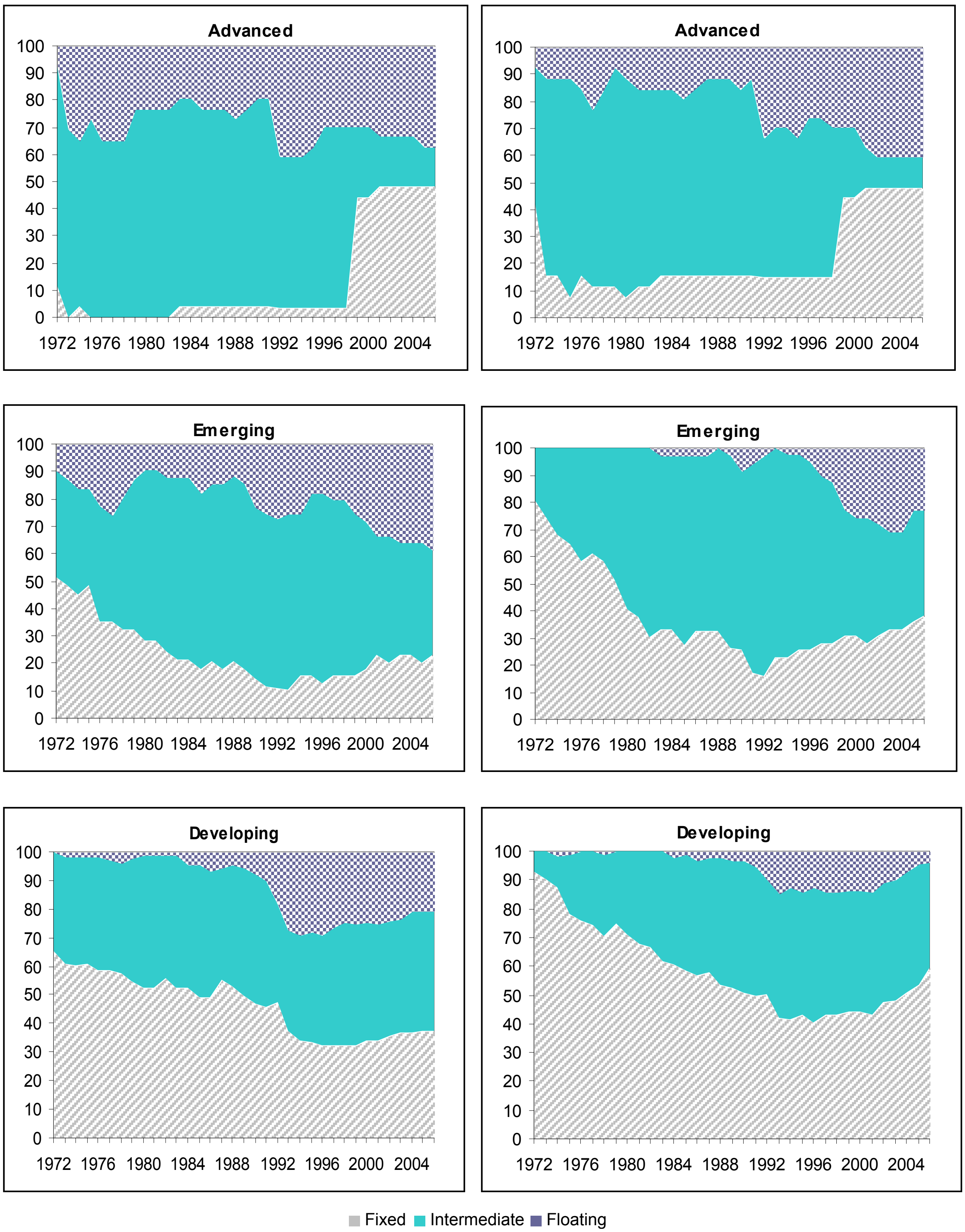


Source: Authors' calculations based on Anderson (2008).
*Sample period is restricted to 1974-2004 corresponding to data availability for the other classifications.

Figure 2. 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 A

Table A1: 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 A2: 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 A3. List of countries in the sample

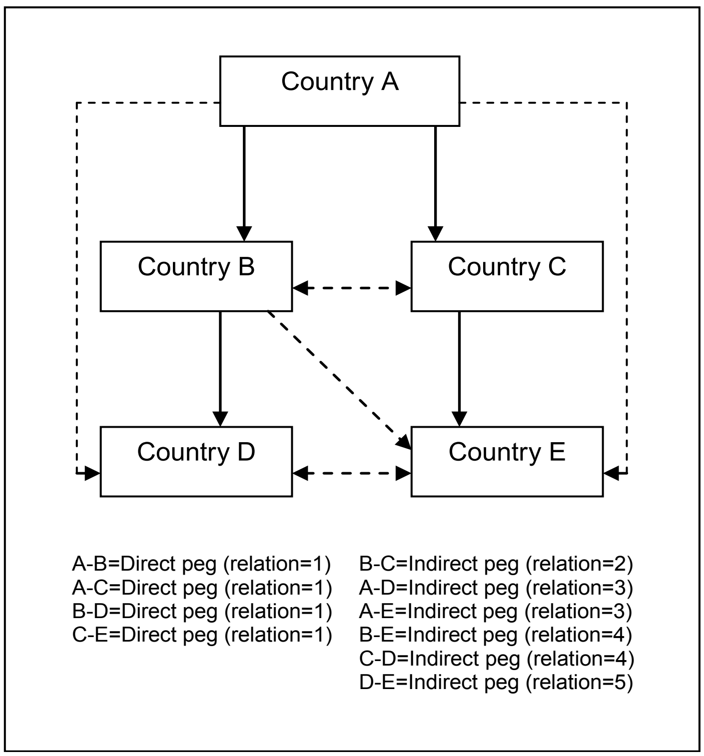
Advanced	***Emerging***		***Developing***			
Australia	Argentina	Russia	Albania	Croatia	Libya	Suriname
Austria	Brazil	Saudi Arabia	Algeria	Dominica	Lithuania	Swaziland
Belgium	Bulgaria	Slovak Republic	Angola	Equatorial Guinea	Macedonia FYR	Syrian Arab R
Canada	Chile	South Africa	Antigua and Barbuda	Estonia	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 A4: 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 A1: Direct and indirect peg relations across countries*



*Adapted from KS (2006).

Appendix B

Table B1: 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 B2: Benchmark specification results with short-run volatility for subsamples, 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 B3: 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 B4: 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)
(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)
(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 B5: Results for sensitivity analysis, 1972-2006

	De jure classification								De facto classification							
Model specification, estimation and simultaneity bias																
	Log (X_{ij})[1]	Quadratic[2]	Sample selection[3]	Sample selection[4]	PPML[5]	FE-GMM[6]	FE-GMM[7]	SYS-GMM[8]	Log (X_{ij})[1]	Quadratic[2]	Sample selection[3]	Sample selection[4]	PPML[5]	FE-GMM[6]	FE-GMM[7]	SYS-GMM[8]
Lagged trade							0.456***	0.215*							0.455***	0.272**
							(0.01)	(0.11)							-0.01	(0)
CU	0.333***	0.332***	0.316***	0.329***	0.196***	0.190***	0.089***	0.526*	0.302***	0.307***	0.292***	-0.227***	0.088	0.155***	0.072**	0.589**
	(0.08)	(0.07)	(0.08)	(0.08)	(0.06)	(0.05)	(0.03)	(0.31)	(0.08)	(0.07)	(0.08)	(0.03)	(0.09)	-0.05	-0.03	(0.262)
Direct peg	0.151	0.369***	0.306***	0.303***	0.091**	0.255***	0.132***	0.369**	0.224***	0.388***	0.329***	-0.241***	-0.147	0.280***	0.151***	0.444*
	(0.10)	(0.11)	(0.110)	(0.110)	(0.04)	(0.06)	(0.04)	(0.15)	(0.08)	(0.10)	(0.10)	(0.03)	(0.14)	-0.05	(0.03)	(0.175)
Indirect peg	-0.0733**	-0.145***	-0.150***	-0.149***	0.079**	-0.158***	-0.121***	0.0323	-0.186***	-0.213***	-0.227***	0.304***	0.015	-0.231***	-0.145***	0.452
	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.09)	(0.02)	(0.33)	(0.03)	(0.03)	(0.03)	(0.08)	(0.04)	-0.02	(0.02)	(0.244)
Volatility[a]	-0.205***	-0.256***	-0.250***	-0.246***	-0.164***	-0.179***	-0.099***	-0.0155	-0.202***	-0.251***	-0.246***	0.327***	-0.156***	-0.177***	-0.099***	-0.057
	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.01)	(0.06)	(0.02)	(0.03)	(0.03)	(0.10)	(0.03)	(0.02)	(0.01)	(0.07)
Observations	177,270	177,270	177,270	177,270	192,906	169,335	159,466	165,671	177,270	177,270	177,270	177,270	192,906	169,335	159,466	165,671
Number of pairs	10,894	10,894	10,894	10,894	10,894	10,310	9,773	10,302	10,894	10,894	10,894	10,894	10,894	10,310	9,773	10,302
AR(1) p-value								0.00								0.00
AR(2) p-value								0.51								0.26
Hansen p-value						0.05	0.08	0.10						0.05	0.07	0.12
Hansen (level) p-value								0.63								0.16
Hansen (diff) p-value								0.07								0.10

Measure of volatility and alternate sub-samples	Vol2[9]	Vol3[10]	Excl. small states[11]	Excl. oil exporters[12]	1975-2006	Vol2[9]	Vol3[10]	Excl. small states[11]	Excl. oil exporters[1]	1975-2006
CU	0.364***	0.275***	0.285***	0.291***	0.246***	-0.229***	0.249***	0.254***	0.248***	0.218***
	(0.08)	(0.07)	(0.08)	(0.09)	(0.08)	(0.03)	(0.07)	(0.08)	(0.09)	(0.08)
Direct peg	0.306***	0.281***	0.300***	0.278***	0.270**	0.342***	0.306***	0.338***	0.274***	0.292**
	(0.11)	(0.11)	(0.12)	(0.09)	(0.13)	(0.09)	(0.10)	(0.11)	(0.08)	(0.11)
Indirect peg	-0.149***	-0.150***	-0.133***	-0.125***	-0.155***	0.326***	-0.234***	-0.211***	-0.226***	-0.233***
	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.10)	(0.03)	(0.03)	(0.03)	(0.04)
Volatility[a]	-0.134***	0.0242	-0.281***	-0.265***	-0.231***	-0.132***	0.0278	-0.276***	-0.259***	-0.228***
	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)
Observations	177,270	177,270	136,060	111,552	170,432	177,270	177,270	136,060	111,552	170,432
Number of pairs	10,891	10,887	7,940	6,364	10,892	10,891	10,887	7,940	6,364	10,892

Source: Authors' calculations.
Estimation results are obtained from the HT method; Robust clustered (by dyad) 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] Fixed effects GMM estimation with lagged values of CU as instruments.
[6] Fixed effects GMM estimation with lagged trade.
[7] System-GMM dynamic panel estimation.
[8] Estimates from the Poisson Pseudo Maximum Likelihood approach where the dependent variable is real bilateral trade.
[9] Defines volatility as the standard deviation of the first difference of logarithms of the real exchange rate.
[10] Defines volatility as a linear transformation of the coefficient of variation of real exchange rates.
[11] Countries with a population of less than 1 million are excluded from the sample.
[12] Oil exporters are excluded from the sample.

Table B6. Trade stability for the world and subsamples, 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	HT
Specification	(1)	(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 (by dyad) 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.