Evidence on the impact of exchange rate regimes on foreign direct investment flows ☆

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Abstract

This paper empirically investigates the impact of different exchange rate regimes upon foreign direct investment (FDI) flows. Using panel data from 27 OECD and non-OECD high income countries for the period 1980 through 2003, and drawing from three different classifications of exchange rate regimes, we find that a currency union is the policy framework most conducive to cross-border investment. Being a member of EMU also appears to spur greater FDI flows with countries floating their currency vis-à-vis the default regime of a ‘double-float’. For country-pairs fixing or pegging their currency to each other, the effect on FDI flows is found to be qualitatively similar but weaker and less robust across alternative model specifications. The effect on FDI flows of country-pairs’ regime combinations involving one country fixing its currency and the other floating or being a member of a currency union, is statistically indistinguishable from that of floating currency country-pairs.

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1. Introduction

Foreign direct investment (FDI) has become even more important than trade as a catalyst for economic development (Fig.1) and now constitutes the single largest source of capital flows. Even though FDI flows have suffered a sharp decline as a result of the more unstable economic climate following 9/11, over the period 1980 to 2003, real world FDI flows experienced almost a 450% increase.[1]

< Figure 1 near here >

These trends have prompted much research on the determinants of FDI (Culem, 1988; Billington, 1999; Chakrabarti, 2001; Sun et al., 2002; Portes and Rey, 2005; De Vita and Abbott, 2007; etc.), its growth enhancing effects (Balasubramanyam et al., 1996; Aitken et al., 1997; Borensztein et al., 1998; Alfaro et al., 2004; Durham, 2004; De Vita and Kyaw, 2007; etc.) and, more recently, the role of firm heterogeneity in influencing the choice of foreign entry mode (Helpman et al., 2004; Nocke and Yeaple, 2007). However, much less attention has been paid to whether the choice of exchange rate regime affects FDI flows between countries. This is particularly striking when it is considered that the examination of the effect of different exchange rate regimes on trade has already produced a voluminous literature (see Rose, 2000; Rose and van Wincoop, 2001; Frankel and Rose, 2002; Glick and Rose, 2002; Rose and Stanley, 2005; Klein and Shambaugh, 2006a to name but a few)[2]. The notable exception is the recent empirical work by Schiavo (2007) who, using a linear gravity-type model on a sample of OECD countries, investigates the impact of EMU on FDI flows for the period 1980-2001. Working within standard neoclassical investment theory and the Dixit and Pindyck's (1994) option value framework of investment under uncertainty, Schiavo argues that the higher the volatility of the exchange rate, the higher the probability that an investment

[1] In 1980 real World FDI flows were US$ 107,848 millions, rising to US$ 601,710 millions in 2003. Source: World Bank, World Development Indicators.

[2] In addition to effects on trade, recent research has uncovered empirical regularities between exchange rate regimes and other macroeconomic variables such as growth (Bailliu et al., 2003; Levy-Yeyati and Sturzenegger, 2003) and national prices (Broda, 2006). By demonstrating "*the extent to which exchange rate regimes really do matter for exchange rate outcomes*" (p.2), Klein and Shambaugh (2006b) provide a sound empirical basis to explain the observed trade and wider macroeconomic consequences of exchange rate regimes.

opportunity be deferred. From this, he goes on to suggest that the elimination of volatility stemming from a currency union "*gives a non-negative impulse to cross-border investment*" (ibid, p.8). Along the lines of the work by Rose (2000) on exchange rate regimes and bilateral trade flows, his gravity model, which revolves around GDP as the 'mass' variable, is augmented with a dummy for a currency union and, as a further test of 'deeper integration', with an additional dummy that switches on when one of the two countries is a member of EMU. Schiavo's OLS and Tobit estimation results seem to confirm the hypothesis that EMU has resulted in larger FDI flows not only between EMU members but also with the rest of the world, though inferences drawn from the EMU coefficient need, by his own admission, to be taken with caution due to the very short length of time (1999-2001) for which EMU data are used.

Our study aims to extend and improve upon Schiavo's work in three fundamental respects. First, currency unions represent only one possible exchange rate regime within the feasible policy set. As Klein and Shambaugh (2006a) point out, often the realistic choice facing policy makers is not whether to abandon the national currency and join a currency union but whether or not to fix or peg their currency to that of one of the major industrialised countries. Additional questions then arise from consideration of the effects on FDI flows of emerging combinations of regimes between pairs of countries, with one country being in a currency union and the other country floating its currency or adopting a fixed exchange rate (CU-FLT and CU-FIX, respectively), one country adopting a fixed exchange rate and the other floating (FIX-FLT), both countries fixing their exchange rate (FIX-FIX), or the case of two countries with a fixed exchange rate policy but targeting different pegs (DFIX). Accordingly, we present what to our knowledge is the first set of estimates of the impact of a wide menu of exchange rate regime combinations on bilateral FDI flows between country-pairs from a relatively large panel. In terms of the categorisation of exchange rate regime data, our contribution is also distinguished by the comparative use of the 'natural' or *de facto* exchange rate regime classifications recently developed by Reinhart and Rogoff (2004) and Shambaugh (2004), in addition to the *de jure* classification published by the International

Monetary Fund (IMF) in its *Annual Report on Exchange Rate Arrangements and Restrictions* (various issues).

Second, we believe that in assessing the impact of an exchange rate regime on FDI flows, it is important to consider which specific alternative regime that effect is benchmarked against. Schiavo's (2007) study examines the impact of a currency union on FDI flows against a composite default regime that aggregates all other alternatives. Our wide menu of regime policy options, on the other hand, allows us to compare the specific effect of each country-pair regime combination *vs.* the single case of a floating currency country-pair, i.e. against the more plausible alternative of a 'double float'.

An additional virtue of our approach is that rather than assuming weak exogeneity of the regressors, we explicitly control for simultaneity bias. Controlling for the presence of endogeneity between explanatory and dependent variables appears to be particularly important in our study since FDI flows, factors influencing regime choice (see Juhn and Mauro, 2002) and variables typically entering gravity-type equations (e.g. GDP per capita), may be simultaneously determined and in some cases causation is likely to run both ways.[3] In this paper, instrumental variable estimation of a dynamic panel model within a system generalised methods of moments (SYS-GMM) framework not only exploits the time series variation in the data while accounting for unobserved country specific effects, it also allows us to control for both a possible correlation between the regressors and the error term, and endogeneity bias.

Our results show that common membership of a currency union is the regime framework most conducive to FDI flows, though we also show that this effect is not solely attributable to the elimination of exchange rate risk. EMU membership also appears to spur greater FDI flows with extra-EMU countries floating their currencies vis-à-vis the level of FDI occurring between country-pairs with flexible exchange rates. FDI flows between country-pairs fixing or pegging their currency are found to be higher than those occurring

[3] For example, the bilateral causation between FDI and various measures of national output is now well documented in the literature (see Ericsson and Irandoust, 2001; and Chowdhury and Mavrotas, 2006).

between floating currency country-pairs, though this effect is weaker than that emerging from a currency union and does not prove robust to alternative model specifications. Other regime combinations (FIX-FLT; FIX-CU; DFIX) are not found to be significantly more pro-FDI than the default regime of a 'double-float'.

Section 2 describes the empirical model and dataset that we use. Section 3 explains the methodology in more detail. Section 4 presents and discusses the estimation results. Section 5 summarizes the main findings and outlines profitable avenues for future research.

2. Model and data

We use an unbalanced panel of 27 OECD and non-OECD high income countries over the period 1980 to 2003[4], yielding 8,014 country-year observations, across 345 country-pairs.[5] Over the full sample, there are 7,270 observations used for estimation.[6] Drawing from standard variables typically entering the gravity equation[7], our parsimonious baseline model is expressed (in long-run form) as:

$$\begin{aligned} fdi_{ijt} &= \delta_0 + \delta_1\, tbt_{ijt} + \delta_2\, y_{it} + \delta_3\, y_{jt} + \delta_4\, RXRVOL_{ijt} + \delta_5\, dis_{ij} + \alpha_6\, LANG_{ij} + \alpha_7\, COL_{ij} \\ &+ \alpha_8\, COMLAN_{ij} + \alpha_9\, FTA_{ijt} + \alpha_{10}\, CU - CU_{ijt} + \varepsilon_{ijt} \end{aligned} \tag{1}$$

[4] The countries are Australia, Austria, Belgium-Luxembourg, Canada, Denmark, Finland, France, Germany, Greece, Hong Kong, Iceland, Ireland, Israel, Italy, Japan, South Korea, Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland, United Arab Emirates, United Kingdom, United States.

[5] Due to the presence of a freely falling exchange rate in the R+R regime classification, to avoid distortions, we exclude observations from 1980 to 1983 for Iceland; 1980 to 1986 for Israel; and 1998 for Korea.

[6] A number of observations are lost since we include two lagged values of the dependent variable when estimating our short-run model by SYS-GMM, from which the long-run coefficients in equation (1) are derived. The full data set is available from the authors upon request.

[7] We also include measures of trade openness and real exchange rate volatility in the right hand side of equation (1). We later extend the model by adding additional regressors that have been found in previous literature to have explanatory power in the determination of FDI.

Where fdi_{ijt} is the log of total bi-lateral real FDI flows between countries i and j at period t. Total FDI is the sum of inward and outward FDI flows, calculated from the OECD's International Direct Investment Statistics database. This database provides data for reporting and partner countries (denoting recipient and investing countries). For most of the country-pairs, two values are reported for the same flow. For example, there are observations estimating inward FDI flows into France from Germany, that are different from the outward FDI flows reported for Germany to France. While in theory the two values should be identical, in practice we take their maximum.[8] tbt_{ijt} is the log of total real bi-lateral trade for the country-pair, computed as the sum of exports and imports. $y_{it}\left(y_{jt}\right)$ denotes the log of real per capita GDP [9] for country *i* (country *j*) and $RXRVOL_{ijt}$ is a measure of exchange rate volatility, calculated as the annual standard deviation of the monthly percentage changes in the real bi-lateral exchange rate. dis_{ij} is the log of geographic distance[10], $LANG_{ij}$ is a dummy variable that equals 1 when both countries share a common official language, while COL_{ij} and $COMLAN_{ij}$ are dummies that switch on to indicate the former or current existence of a colonial relationship and, land adjacency, respectively. To capture goods market integration, FTA_{ijt} equals one when both countries have a free trade agreement. The $CU\text{-}CU_{ijt}$ dummy variable equals unity when both countries, *i* and *j*, are members of the same currency union in year *t*.

In equation (1), the coefficient for $CU\text{-}CU_{ijt}$ compares the level of FDI to a composite default regime that consists of: two countries fixing their exchange rate to each other (FIX-FIX)[11]; both countries floating their currencies (FLT-FLT); one country floating its

[8] An alternative approach would have been to use the average. However, for many of the non-OECD countries there was significant under-reporting of FDI data and, as a result, in some of these cases only one observation was available.

[9] Like Micco et al. (2003) we use per capita GDP though, to avoid over-parameterization, we do not add a separate variable for population, for which information is already fed into our model also through our weighted distance measure. As a robustness check, we later re-estimate the model using GDP.

[10] Rather than using simple distance, our measure is based on bilateral distances between the biggest cities of the two countries, with those inter-city distances being weighted by the share of the city in the overall country's population.

[11] This includes direct pegs and indirect pegs when the two currencies are tied indirectly through a peg to the same third country's currency.

currency and the other in a currency union (CU-FLT); a country in a currency union and the other fixing its currency (CU-FIX); one country fixing its currency and the other floating (FIX-FLT); and two countries following a fixed exchange rate policy but to different pegs (DFIX).

We then build on this baseline model by changing the composition of equation (1) to compare the level of FDI under each country-pair regime combination against a single alternative. Specifically, by including all of the individual exchange rate regime dummies on the right hand-side of equation (1), we can establish the difference in the level of total FDI flows of each regime combination against the 'double float' case as the sole benchmark.

< Table 1 near here >

We draw from three different exchange rate regime classifications to construct the regime dummies. Our first classification is that published by the IMF in its *Annual Report on Exchange Rate Agreements and Restrictions* (various issues). Traditionally, the IMF's classification was compiled solely on the basis of countries' announced regimes, thus failing to capture the extent to which actual policies conformed to countries' declared commitment. Since 1999 the IMF moved to a hybrid classification scheme that combines data on the actual behaviour of the exchange rate with information on the policy framework.[12] Despite this, concerns arising from the IMF's *de jure* or hybrid classification have prompted researchers to develop alternative classification schemes that attempt to characterize more accurately countries' *de facto* regimes. For comparative purposes and as a robustness test we, therefore, also use two of such alternative classification schemes. The one developed by Reinhart and Rogoff (2004) (henceforth, R+R), which incorporates data on market determined, dual/parallel exchange rates, and the one developed by Shambaugh (2004), which is based on the behaviour of countries' official exchange rates.[13] Table 1 shows the exact correspondence between the original categories of the three classification schemes and those we derived from

[12] For details, see IMF (1999).
[13] Other attempts to base classifications on the *de facto* regimes include those by Ghosh et al. (1997) and Levy-Yeyati and Sturzenegger (2003). However, as noted by Husain et al. (2005), these alternative classification schemes tend to suffer from reduced samples with approximately one-third of the observations missing or unclassifiable.

them to inform our menu of exchange rate regimes.[14] Further information on the data series used and their sources is provided in the Data Appendix.

In this paper we are deliberately cautious in declaring a priori expectations for the various exchange rate regime dummies because the core of conventional trade theory has nothing to say on FDI and the multinational firm (Ethier, 1995) and, as recently noted by Ricci (2006) *"location theory normally does not consider the effects of exchange rate regimes on the location choices of firms"* (p. 52). Indeed, a formal analytical framework of investment behaviour under the full menu of exchange rate regimes has yet to be developed, and the few theoretical contributions on the relationship between FDI and exchange rate volatility (Froot and Stein, 1991; Campa, 1993; Dixit and Pindyck, 1994; Goldberg and Kolstad, 1995; Blonigen, 1997) have postulated different outcomes depending on the assumptions employed in relation to the risk preferences of foreign investors, cost reversibilities, entry conditions and the timing of production decisions. This leaves the arduous task of resolving the question of the impact of exchange rate regimes on FDI flows to the empirical research. Our aim here, therefore, is to start filling this vacuum by letting the data tell part of the story.

3. Methodology

To explain the econometric methodology we employ, firstly let us consider a simple AR(1) model with unobserved individual-specific effects:

$$y_{it} = \beta\, y_{i,t-1} + \chi_i + \nu_{it} \qquad \text{for } i = 1, \ldots, N \text{ and } t = 2, \ldots, T \qquad (2)$$

where $\chi_i + \nu_{it} = \mu_{it}$ has the standard error component structure:[15]

$$E[\chi_i] = 0;\quad E[\nu_{it}] = 0;\quad E[\chi_i\, \nu_{it}] = 0 \qquad \text{for } i = 1, \ldots, N; \text{ and } t = 2, \ldots, T \qquad (3)$$

Assuming serially uncorrelated transient errors

$$E\left[\nu_{it}\, \nu_{iq}\right] = 0 \qquad \text{for } i = 1, \ldots, N \text{ and } q \neq t \qquad (4)$$

[14] The extent of the 'compression' of the fine codes of the original regime classifications' categories within the categories we adopt to construct our regime dummies, was chosen to ensure empirical tractability whilst allowing a sufficiently wide and informative menu of feasible policy options to be examined.

[15] The individual effects are categorised at the country-pair level to account for unobservable factors that could explain two countries' greater or lesser tendency to undertake cross-border investment.

and that the initial conditions y_{i1} are predetermined

$$E\left[y_i \nu_{it}\right] = 0 \quad \text{for} \quad i = 1, \ldots, N \text{ and } t = 2, \ldots, T \tag{5}$$

the following m = ½ (T – 1) (T – 2) moment restrictions are obtained

$$E\left[y_{i,t-q} \Delta\nu_{it}\right] = 0 \qquad \text{for } t = 3, \ldots, T \text{ and } q \geq 2 \tag{6}$$

or, more compactly

$$E\left(Z_i' \Delta\nu_i\right) = 0 \tag{7}$$

where Z_i is the (T – 2) x m matrix and $\Delta\nu_i$ is the (T-2) vector $\left(\Delta\nu_{i3}, \Delta\nu_{i4}, \ldots, \Delta\nu_{iT}\right)'$. As proposed by Arellano and Bond (1991), these are moment restrictions exploited by the standard linear first-differenced GMM estimator, which entails the use of lagged levels as instruments for the equations in first-differences. This yields a consistent estimator of β when N approaches infinity and T is fixed.

However, there are shortcomings with this first-differenced estimator. Alonso-Borrego and Arellano (1999), Blundell and Bond (1998) and Blundell et al. (2000) show that if the series are highly persistent or if the variance of the individual specific effect is large relative to the variance of the remainder of the error term, then lagged levels make weak instruments for the regression equation in differences. Instrument weakness, in turn, affects the asymptotic and small-sample performance of the first-differenced GMM estimator. Asymptotically, the variance of the coefficients increases while, in small samples, instrument weakness may produce biased estimates. To reduce the imprecision and potential bias associated with the standard GMM estimator, the system GMM (henceforth, SYS-GMM) model of Arellano and Bover (1995) and Blundell and Bond (1998) is estimated.[16]

SYS-GMM imposes the additional assumption that:

$$E\left(\chi_i \Delta y_{i2}\right) = 0 \quad \text{for } i = 1, \ldots, N \tag{8}$$

This assumption, which requires the stationarity of the process generating the initial conditions y_{i1}, yields (T-2) further linear moment conditions:

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[16] Blundell et al. (2000) report dramatic gains in estimate efficiency from SYS-GMM when the series used are highly persistent.

$$E(\mu_{it}\ \Delta y_{i,t-1}) = 0 \qquad \text{for } i = 1, \dots, N \text{ and } t = 3, 4, \dots, T \qquad (9)$$

The additional T-2 further moment conditions allow the construction of a GMM estimator that uses both sets of moment restrictions, (6) and (9), in a stacked system of lagged first-differences in the series as instruments for T-2 equations in levels, in addition to lagged levels as instruments for T-2 equations in first-differences.[17]

The full set of second-order conditions available now becomes

$$E\left(Z_i^{+'}\ \mu_i^{+}\right) = 0 \qquad (10)$$

where $\mu_i^{+} = \left(\Delta\nu_{i3},\dots,\Delta\nu_{iT},\dots,\mu_{i3},\dots,\mu_{iT}\right)'$.

This procedure seems particularly appropriate in our context, both conceptually and for its statistical virtues. First, FDI might adjust slowly to changes in the regressors, reflecting persistence in the series over time.[18] Lagged values of the dependent variable can be included in the estimated equation to account for speed of adjustment.[19] Second, FDI and one or more of the regressors may be simultaneously determined, forcing a correlation between the estimated residuals and the regressors. With SYS-GMM every regressor is instrumented, so potential issues of endogenity bias are overcome, provided that instrumental validity is ensured through testing. Third, including both level and first difference equations in a stacked system allows us to investigate whether time-invariant variables, such as distance and adjacency, play an important role in the determination of total FDI, something not possible from the standard GMM estimator.

The specific linear dynamic model that we use for estimation[20] can be defined as:

[17] For an illustration of the structure of the matrix of instruments when the full set of orthogonality conditions is exploited, see Blundell and Bond (1998) and Blundell et al. (2000).

[18] Using two lags of the dependent variable across various specifications, our estimates indicate a mean coefficient for the AR(1) component of 0.42 and an average value of 0.28 for the AR(2) component.

[19] Estimating a short-run model by fixed or random effects would produce biased estimates (see, for example, Nickell, 1981), due to the correlation between the estimated residuals and the autoregressive term. SYS-GMM overcomes this problem while establishing whether a dynamic relationship exists between the dependent variable and the regressors.

[20] All regressions were estimated using xtabond2 for Stata (Roodman, 2005).

$$y_{ijt} = \alpha_0 + \sum_{k=1}^{p} \alpha_k \, y_{ijt-k} + \sum_{l=0}^{q} \beta_l \, x_{ijt-l} + \eta_i + \lambda_t + \nu_{ijt}$$

$$i=1,\ldots,n \qquad t=1,\ldots..T \qquad (11)$$

where y_{ijt} is total FDI flows, $y_{ij1}, \ldots\ldots y_{ijp}$ represent the autoregressive structure, and $x_{ij0}, \ldots\ldots x_{ijq}$ are the current and lagged values of the matrix of regressors that could be strictly exogenous, predetermined or endogenous with respect to ν_{ijt}, the error term.[21] η_i are individual effects that estimate differences in the mean level of FDI across country-pairs. λ_t are time specific effects that capture the effect of common disturbances and/or spatial correlation across the units of the panel. Transformation of equation (11) leads to a set of T_i equations being estimated across the country pairs. Provided that T>3, then for every period and with a lag length of q, (T_i-q) first-difference equations are estimated using (T_i-q) lagged level instruments, with (T_i-q) level equations estimated using (T_i-q+1) first-difference instruments.[22] As T grows in size, the computational requirements of SYS-GMM rise significantly, resulting in a trade-off between gains in efficiency and greater bias in the estimates, due to over-fitting of the estimated equation with too many instruments. We therefore restrict the instrument matrix so that only one lag is used for the first difference equation, while for the level equation, the contemporaneous first difference and its lag are used.

Consistency of the SYS-GMM estimates requires evidence of significant first-order serial correlation[23] but no higher order serial correlation. For this purpose, we employ the (ARp) Arellano-Bond statistic (Arellano and Bond, 1991). Moreover, to verify the validity of

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[21] Endogenity rules out x_{ijt} being correlated with future shocks, ν_{ijt+1}, ν_{ijt+2},….., though x_{ijt} can be correlated with ν_{ijt} and earlier shocks. Predetermination implies no correlation between the regressors and contemporaneous shocks, though x_{ijt} could be correlated with lagged shocks. Strict exogenity requires x_{ijt} to be uncorrelated with all past, present and future shocks.

[22] We use a two-step version of the SYS-GMM procedure with finite-sample adjustment to the covariance matrix that corrects for estimate bias (Windmeijer, 2005). The estimated standard errors are robust to heteroscedasticity.

[23] By construction, from the first-difference equation, $\Delta\nu_{ijt} = \nu_{ijt} - \nu_{ijt-1}$ should correlate with $\Delta\nu_{ijt-1} = \nu_{ijt-1} - \nu_{ijt-2}$ due to the common element ν_{ijt-1}.

the chosen instruments, the Hansen's (1982) J-test for over-identifying restrictions is adopted. This test statistic is robust to problems of autocorrelation and heteroskedasticity that can arise from using the Sargan's test (Sargan, 1958) with SYS-GMM.[24]

Another important consideration is how many lags of the dependent variable and regressors should be included. We found that two lags of the dependent variable were required to ensure an absence of second order serial correlation. Lagged values of the time varying regressors were also included when their estimated coefficients were found to be statistically significant. Using estimates of the coefficients of the current and lagged regressors and those from the autoregressive terms, we were able to obtain estimates of the long-run effect of exchange rate regimes on FDI flows and approximate the associated standard errors using the delta method (see Feiveson, 1999).

4. Results

The first set of estimates relating to our baseline model (1) is presented in Table 2. Here, the only exchange rate dummy included is CU-CU, the effects of which are measured against 'all other regimes' as the default benchmark. The CU-CU dummy displays a positive sign and is significant at the 5% level, across all regime classification specifications. Over the sample of observations, FDI flows were from 24.73% ($\exp^{0.221} - 1$) to 30.3% ($\exp^{0.265} - 1$) higher for EMU members compared to all other country-pairs.[25] Though qualitatively consistent with Schiavo's findings, our estimated CU-CU dummy suggests a much more modest increase in cross-border investment flows than that he reported (up to 320%). We attribute this difference to the more selective composition of our sample[26], and different methodology employed.

[24] The J-statistic, which is the minimized value of the two-step GMM criterion function, has an asymptotic chi-squared distribution, where the number of degrees of freedom equals the number of instruments used.
[25] This range of estimates arises since Reinhart and Rogoff (2004) classify Greece as a *de facto* member of EMU from 1999, whereas the IMF and Shambaugh classifications show a start date of 2001.
[26] We do not include Hungary, Poland and the Czech Republic since information on their *de facto* regimes is not available for all years. Mexico is also dropped since it is not a high income country like the rest of our sample, which was deliberately kept fairly homogenous to minimise possible bias stemming from linear approximation of the 'true' model's nonlinearity. However, we also include Israel, Korea, Singapore and the United Arab Emirates.

< Table 2 and Figure 2 near here >

Table 3 shows the estimation results using our full menu of exchange rate regime combinations against the default regime of a 'double float'. The model works well on a number of different dimensions. The plots of the residuals (see Figure 2) suggest that the estimated errors are indeed white noise. The Hansen's J-test for instrumental validity, as well as exclusion and 2^{nd} order serial correlation tests do not reject the chosen econometric specification, which appears to fit the data well, explaining over two-thirds of the variation in FDI flows. To check for the presence of undesirable correlation patterns among the variables, a correlation matrix was also estimated. As shown in Table 4, since none of the estimated coefficients have values in excess of ±0.70, the results indicate the absence of any multicollinearity problems.

< Tables 3 and 4 near here >

From Table 3 it is evident that the effect of a currency union emerges as the strongest among the combinations of policy options examined, though other regime combinations also appear to promote FDI. Specifically, a currency union (CU-CU) is found to increase FDI flows between member countries vis-à-vis floating currency country-pairs across all regime classification specifications. In two of the three cases (IMF and Shambaugh's specifications), the magnitude of this effect is now greater than that obtained from our baseline model, though still considerably lower than the one reported by Schiavo (2007). Like Rose (2000), we interpret a positive and significant currency union coefficient in conjunction with an insignificant exchange rate volatility coefficient as evidence that the effect of a currency union in promoting FDI does not merely stem from the elimination of exchange rate risk. Other forces such as the removal of informational barriers (see Mody et al., 2003; Portes and Rey, 2005) may be at work in driving international investment decisions.

Interestingly, we also find that EMU membership promotes FDI flows with countries floating their currency (CU-FLT dummy), with an estimated increase in FDI flows of the order of 15-23% vis-à-vis the default regime of floating currency country-pairs. This result can be rationalised by the attractiveness of EMU for extra-EMU investors, both as a means for tariff jumping FDI and as a platform for exports throughout the Customs Union. For country-pairs fixing or pegging their currency to each other (FIX-FIX), the impact on FDI flows appears to be qualitatively similar to that of a currency union but weaker, as one would expect, given the greater uncertainty and larger transaction costs that the FIX-FIX regime combination entails compared to common membership in a currency union. However, it should be noted that this effect is only significant in two out of the three regime classification specifications. Finally, the effect of the CU-FIX, FIX-FLT and DFIX dummies, is found to be statistically indistinguishable from that of the default regime of a 'double float' in most cases.

Although in this paper our interest centres upon the impact of exchange rate regimes on FDI flows, it is worth noting that most of the coefficients of the control variables are statistically significant, with sensible economic interpretations. Our estimated coefficient of trade openness (tbt) is positive and significant in all three cases. This may be due to the fact multinational organisations are attracted to open economies by virtue of their instrinsic export potential and more stable economic climate. Geographic distance (dis) appears to play an insignificant effect on FDI flows while cultural proximity, proxied by common language (LANG)[27], seems to encourage cross-border investment considerably. Görg and Wakelin (2002) and De Vita and Abbott (2007) report a similar effect of the language dummy on US and UK FDI respectively. A former or current colonial relationship (COL) does not appear have any significant impact on FDI flows and there is only weak evidence in support of an effect of a free trade agreement (FTA) and a common land border (COMLAN). Consistent with the standard gravity theory prediction and previous empirical findings (Culem, 1988; Billington, 1999; Chakrabarti, 2001), the income coefficients of country *i* and country *j* (y_i and y_j, respectively) are both positive and significant, signalling the importance of an expanding market for producers' goods in the supply of FDI. [28]

[27] In our sample, the common language dummy applies to the English language (for the USA, the UK, Canada, Australia and New Zealand), to French (for Belgium-Luxemburg, Canada, France and Switzerland) and to German (for Austria and Germany).

[28] To further examine our preferred model we also added trend terms. As shown in Figure 1, World FDI flows have grown significantly over the course of our chosen sample period. Given that all of the CU-CU observations are located in the last five years, the positive and statistically significant CU-CU

< Tables 5 and 6 near here >

Still keen on gaining a deeper understanding of the forces at work in the data generation process, we re-examined the stylised facts of mean FDI flows by country and by regime type. In order to put our sample's geographical distribution of FDI flows in perspective, Table 5 provides a brief summary of the size of the economies examined, their mean FDI flows, and their share as a percentage of world FDI flows. From Table 5 it is evident that the USA and the UK, the dominant FLT-FLT country-pairs, collectively account for just over 30% of world FDI flows and almost 40% of our sample's aggregate FDI flows. Table 6, which presents summary statistics of mean FDI flows by regime and across classification schemes, also shows that a very different picture emerges when the USA and the UK are excluded from the FLT-FLT country-pairs. In 18 cases (IMF), 22 cases (R+R) and 14 cases (Shambaugh) out of the 24 years considered, mean FDI flows of floating currency country-pairs were higher compared to fixed exchange rate country-pairs, though this conclusion can only be made in 0 cases (IMF and Shambaugh) and 8 cases (R+R) when the USA and the UK are excluded from the sample.

< Table 7 near here >

Against this backdrop, to establish the extent to which USA and UK data are driving our results, we decided to 'play with' the FLT-FLT default regime and re-estimate the model by excluding such countries from the sample (Table 7). Remarkably, the results of this exercise are broadly consistent with those reported in Table 3. With the exception of the DFIX coefficient, which is now significant and negatively signed in all three cases, most of the estimated regime dummies prove fairly robust to this new sample selection (though in the R+R's classification specification the CU-FIX estimated coefficient is now significantly negative). The reliability of this new set of results is corroborated by the economically sensible changes to the coefficients of some of our control variables. For example, having

estimates could be picking up the overall rise in FDI. However, we found that the estimates for the trend term and its squared values were not statistically significant. Other features of our data set were: (i) occasionally the presence of negative values due to disinvestment; and (ii) an absence of FDI flows between some country-pairs. We therefore re-estimated our model with only positive values and found that the conclusions of Table 3 were broadly confirmed.

now excluded USA and UK data from the sample, it is not surprising that the LANG coefficient has become insignificant at the 5% level across all three regime classification models. The fact that, in the absence of the substantial FDI flows between the geographically distant UK and USA, the coefficient for COMLAN is now positive and significant in all cases, also supports economic intuition.

< Table 8 near here >

Finally, to test further the sensitivity of our results to alternative specifications, we subjected our preferred model (Table 3) to a series of perturbations. First, we replaced real per capita GDP with real GDP (the output measure used in Schiavo, 2007). Second, to alleviate the potential problem of omitted variable bias, we added two additional regressors. One is a long-term measure of exchange rate volatility to account for persistence or mean reversion in exchange rate movements. Following Clark et al. (2004), this measure is constructed as the standard deviation of the log of monthly changes in the real exchange rates using observations from the preceding five years. The other additional regressor is included to capture the direct effect of informational flows that may not have already been picked up by geographic proximity (dis) and the transparency inherent in trade openness (tbt). FDI usually requires significant fixed costs in acquiring knowledge of the local environment and coordinating activities with suppliers and distributors, thus a well developed informational infrastructure may play a role in the FDI location decision (see also Mody et al., 2003). Following Schiavo (2007), our proxy is based on (the product of) the number of fixed telephone lines per 1,000 people, for country *i* (inf_{it}) and country *j* (inf_{jt}).

As shown in Table 8, looking first at the specifications based on the IMF and Shambaugh's regime classifications, the results of this exercise do not significantly change the conclusions reached heretofore. While still significant and consistently signed, the income coefficients are of smaller magnitude than those obtained previously and so are the CU-CU and CU-FLT estimated dummies. The FIX-FIX estimated dummy is also of smaller magnitude according to the Shambaugh regime classification specification (and only significant at the 10% level) while using the IMF specification the coefficient turns insignificant. The estimated informational flow coefficients show a remarkable similarity across regime classification specifications, suggesting that improvements in telecommunication and informational infrastructure induce greater bilateral FDI flows. Long-run volatility (LRRXRVOL) is statistically insignificant across all specifications.

However, some anomalies now emerge in the estimated coefficients based upon the R+R regime classification specification, which does not prove robust to these extensions and perturbations. The CU-CU dummy is now negatively signed and insignificant, the FIX-FIX estimated coefficient is still negatively signed but now significant, and CU-FLT has changed from positive and significant (at the 10% level) to negative and insignificant.

We attribute the greater sensitivity of the estimates obtained drawing from the R+R classification, and their misalignment vis-à-vis the results obtained from the other two regime classifications, to two concomitant factors. First, although the R+R and Shambaugh's regime classifications are both *de facto* schemes, the former exhibits a much higher degree of disagreement with the IMF's regime classification. For example, in the subsets of CU-CU and FIX-FIX country-pair observations, the R+R classification differs from the IMF on 8% and 70% of the observations respectively, while in Shambaugh's classification the disagreement only applies to 0% and 53% of the observations. Particularly noticeable country-specific examples of divergence include the observations for Greece, 1999 and 2000, recorded as FIX in the IMF and Shambaugh's regime classifications and as CU in the R+R scheme, and the observations for Switzerland, from 1982 onwards, recorded as FLT in the IMF and Shambaugh's regime classifications but as FIX in the R+R scheme. Second, the IMF and Shambaugh's classification schemes are, by construction, able to capture year-to-year instability in a way which, being based on five year windows, the R+R classification cannot. While the R+R classification may perform better in the context of a very long time horizon with relatively smooth and stable data series, it appears inappropriate in dealing with the high degree variability that FDI data, being based on a series of discrete events or 'episodes', exhibit year by year and across countries.

5. Conclusion

In this paper we investigated the impact of different exchange rate regimes on FDI flows using panel data from 27 OECD and non-OECD countries over the period 1980 through 2003. Drawing from three different exchange rate regime classification schemes, we presented the first set of SYS-GMM estimates of the impact of a wide menu of exchange rate regime combinations on FDI flows between 345 country-pairs while controlling for a standard set of gravity and non-gravity type explanatory variables, correlation issues and endogeneity bias.

Overall, the main contribution of this paper is to demonstrate that by exerting greater or lesser stability of exchange rates, exchange rate regimes do affect FDI flows. Our findings suggest that a currency union constitutes the regime type most conducive to cross border investment. Being a member of EMU also appears to increase FDI flows with extra-EMU countries floating their currency compared to the level of FDI occurring between country-pairs with flexible exchange rates. Probably due to the greater uncertainty, for countries fixing or pegging their currency to each other, the increase in FDI flows vis-à-vis the default regime of a ‘double-float’ is less than that experienced through EMU membership but we find this effect to be the most sensitive to alternative model specifications and, as a result, the most ambiguous. The effect on FDI flows of other regime combinations is consistently found to be statistically indistinguishable from that of floating currency country-pairs.

Although our results pass a battery of econometric checks and, in two out of the three exchange rate regime classification specifications, prove robust to several extensions and perturbations to the model (changes to the composition of the sample, alternative measures of national output and additional regressors), in interpreting our findings, two caveats must be borne in mind. First, despite a relatively large dataset covering 27 countries over a relatively long time period, during this sample the only currency union in place was EMU, with no recorded episodes of dissolved common currency linkages among the 345 country-pairs examined. We were, therefore, unable to compare FDI flows before and after a currency union regime change. Second, although compared to Schiavo’s study our dataset benefits from two additional years of EMU data, an even longer time span of EMU data availability would help provide a more reliable empirical basis for policy analysis. Profitable avenues for future research would also entail re-running the model using FDI data disaggregated by type of investment (M&A and greenfield) and by sector. In addition to testing the general findings of our aggregate study when all the sectors are pooled, the latter exercise would offer sector-specific estimates to shed light on whether ICIR (imperfect competition and increasing returns) like effects (for example, the impact of uncertainty on market structure) may play a role in how different exchange rate regimes affect FDI flows. This, however, awaits more data.

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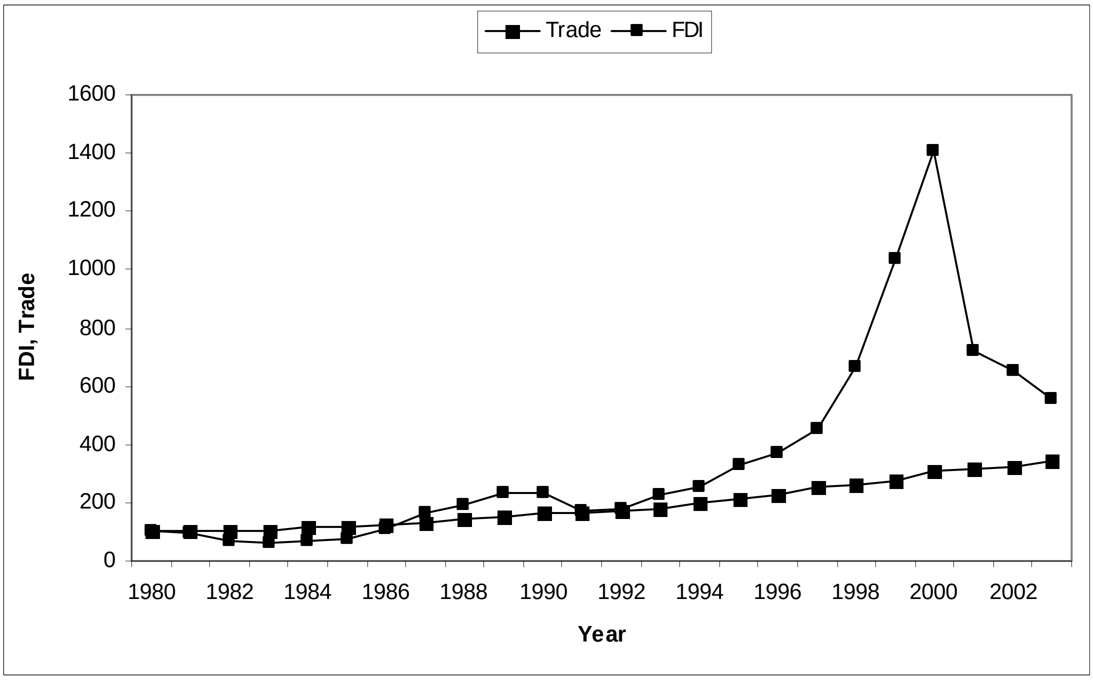
Figure 1: World Total FDI Flows and Total Trade (base year, 1980=100)



Source: World Bank, World Development Indicators.

Table 1: Classifications of exchange rate regimes

	Reinhart and Rogoff's classification	IMF's classification	Shambaugh's classification	Our classification for exchange rate dummies
1	No separate legal tender.	Currency union	Currency union	Currency union
2	Pre announced peg or currency board arrangement.	Currency board/ Currency peg within horizontal band of ±1%.	De facto currency peg. Either 1. no fluctuation at all; 2. movement within 1% band or 3. movement within 2% band; 4. One time devaluation in 1 month with 0% change in the remaining 11 months.	Fixed exchange rate
3	Pre announced horizontal band that is narrower than or equal to ±2%.			
4	De facto peg.			
5	Pre announced crawling peg.	Currency peg within crawling band of ±1%.		
6	Pre announced crawling band that is narrower than or equal to ±2%.			
7	De facto crawling peg.			
8	De facto crawling band that is narrower than or equal to ±2%.			
9	Pre announced crawling band that is wider than or equal to ±2%.	Currency peg within crawling band of at least±1%.		
10	De facto crawling band that is narrower than or equal to ±5%.			
11	De facto moving band that is narrower than or equal to ±2%			
12	Managed floating	Managed floating/ Independently floating	Currency float	Currency float
13	Freely floating			
14	Freely falling	N/A	N/A	N/A
15	Dual market in which parallel market data is missing			

Notes: Classifications are obtained from Reinhart and Rogoff (2004), Shambaugh (2004) and various issues of the IMF's Annual Report on Exchange Rate Arrangements and Restrictions.

Table 2: Baseline model: CU-CU against 'all other regimes' as the benchmark

Variable	R+R	IMF	Shambaugh
constant	4.600[a]	4.830[a]	4.830[a]
	(3.84)	(4.19)	(4.19)
tbt_{ijt}	0.061[a]	0.060[a]	0.060[a]
	(3.66)	(3.67)	(3.67)
y_{it}	0.175[a]	0.163[a]	0.163[a]
	(2.96)	(2.87)	(2.87)
y_{jt}	0.149[a]	0.135[a]	0.135[a]
	(2.70)	(2.53)	(2.53)
FTA_{ijt}	0.032	0.037	0.037
	(0.47)	(0.56)	(0.56)
dis_{ij}	0.007	0.009	0.009
	(0.19)	(0.25)	(0.25)
$COMLAN_{ij}$	0.177	0.176[b]	0.176[b]
	(1.64)	(1.65)	(1.65)
COL_{ij}	0.100	0.099	0.099
	(0.65)	(0.65)	(0.65)
$LANG_{ij}$	0.141[a]	0.142[a]	0.142[a]
	(2.13)	(2.10)	(2.10)
$RXRVOL_{ijt}$	-0.172	-0.216	-0.216
	(-0.27)	(-0.34)	(-0.34)
$CU\text{-}CU_{ijt}$	0.221[a]	0.265[a]	0.265[a]
	(3.03)	(3.49)	(3.49)
Diagnostics from the short-run model			
Wald test: all regressors~χ^2(df)	2686.87[a] (30)	2658.19[a] (30)	2658.19[a] (30)
Wald test: time dummies~χ^2(df)	62.01[a] (22)	64.13[a] (22)	64.13[a] (22)
Wald test: exchange rate dummy~χ^2(df)	9.15[a] (1)	13.10[a] (1)	13.10[a] (1)
J-test~χ^2(df)	290.91 (321)	290.10 (321)	290.10 (321)
AR(1) test~N(0,1)	-3.53[a]	-3.54[a]	-3.54[a]
AR(2) test~N(0,1)	-0.34	-0.35	-0.35
Number of observations	7,270	7,270	7,270
R^2	0.692	0.693	0.693

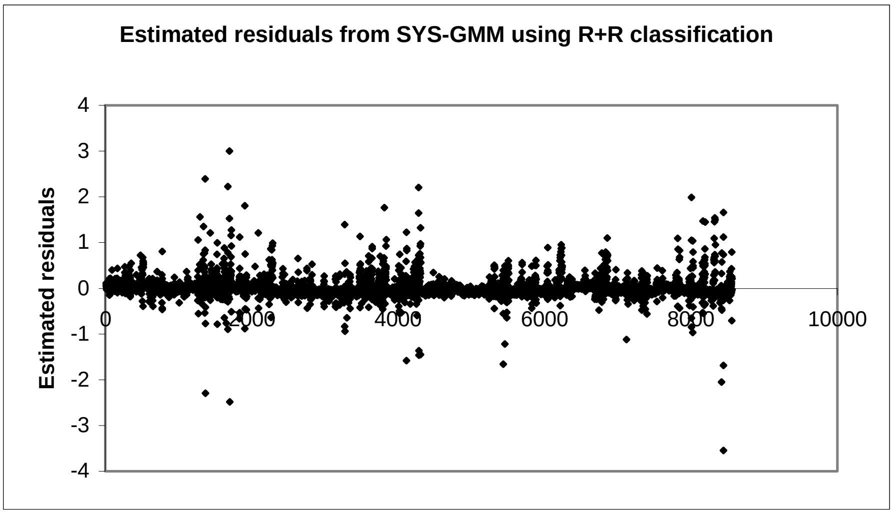
Notes: A short-run dynamic model was estimated using a two-step system GMM procedure (time dummies also included but not reported), from which the above long-run coefficients are derived. Robust t-ratios (reported in parentheses) are approximated using the delta method. The Wald tests are for exclusion restrictions on the coefficients, the number of restrictions is reported in parentheses. J-test denotes the Hansen J-test for over-identifying restrictions, the null hypothesis being instrument validity. AR(1) and AR(2) are tests for 1st and 2nd order serial correlation. "a" denotes significance at the 5% level and "b" at the 10% level. The CU-CU coefficient compares the level of total FDI when both countries are part of EMU vis-à-vis a default benchmark that aggregates all other regimes.

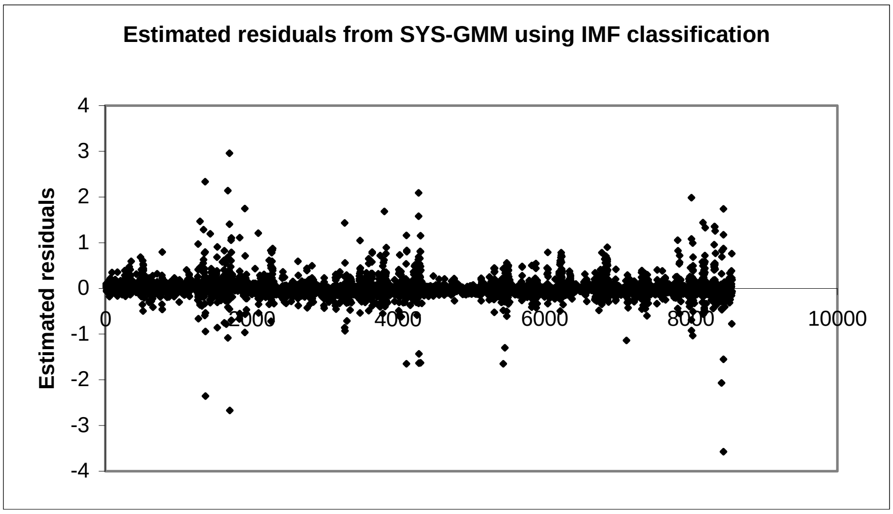
Table 3: Full menu of exchange rate regime dummies *vs.* 'double float' as the benchmark

Variable	R+R	IMF	Shambaugh
constant	5.470[a]	4.260[a]	4.231[a]
	(4.07)	(3.97)	(3.61)
tbt_{ijt}	0.044[a]	0.060[a]	0.167[a]
	(2.04)	(3.95)	(2.06)
y_{it}	0.235[a]	0.147[a]	0.166[a]
	(2.40)	(3.32)	(3.37)
y_{jt}	0.016	0.220[a]	0.207[a]
	(0.28)	(3.27)	(3.01)
FTA_{ijt}	0.111[b]	-0.009	0.0003
	(1.82)	(-0.19)	(0.01)
dis_{ij}	-0.005	0.004	0.0055
	(-0.18)	(0.14)	(0.18)
$COMLAN_{ij}$	0.190[a]	0.115	0.141
	(2.05)	(1.20)	(1.45)
COL_{ij}	0.119	0.135	0.159
	(0.78)	(0.98)	(1.09)
$LANG_{ij}$	0.136[a]	0.120[a]	0.124[a]
	(2.11)	(2.02)	(2.17)
$RXRVOL_{ijt}$	0.273	0.215	0.590
	(0.60)	(0.30)	(0.88)
$CU\text{-}CU_{ijt}$	0.187[a]	0.408[a]	0.443[a]
	(2.09)	(4.16)	(4.84)
$FIX\text{-}FIX_{ijt}$	-0.073	0.150[a]	0.125[b]
	(-0.88)	(2.11)	(1.89)
$CU\text{-}FLT_{ijt}$	0.161[a]	0.219[a]	0.229[a]
	(2.08)	(2.92)	(3.71)
$CU\text{-}FIX_{ijt}$	0.040	-0.005	0.048
	(0.52)	(-0.09)	(0.69)
$FIX\text{-}FLT_{ijt}$	-0.012	-0.031	0.043
	(-0.17)	(-0.64)	(1.14)
$DFIX_{ijt}$	0.064	-0.123[a]	-0.034
	(0.89)	(-2.49)	(-0.62)
Diagnostics from the short-run model			
Wald test: all regressors~χ^2(df)	1315.18[a] (36)	3137.73[a] (37)	3408.71[a] (36)
Wald test: time dummies~χ^2(df)	47.43[a] (22)	78.38[a] (22)	65.96[a] (22)
Wald test: exchange rate dummies~χ^2(df)	26.61[a] (6)	34.18[a] (6)	32.30[a] (6)
J-test~χ^2(df)	255.87 (274)	332.69 (549)	323.35 (548)
AR(1) test~N(0,1)	-3.56[a]	-3.55[a]	-3.55[a]
AR(2) test~N(0,1)	-0.01	-0.31	-0.29
Number of observations	7,270	7,270	7,270
R^2	0.679	0.695	0.693

Notes: A short-run dynamic model was estimated using a two-step system GMM procedure (time dummies also included but not reported), from which the above long-run coefficients are derived. Robust t-ratios (reported in parentheses) are approximated using the delta method. The Wald tests are for exclusion restrictions on the coefficients, the number of restrictions is reported in parentheses. J-test denotes the Hansen J-test for over-identifying restrictions, the null hypothesis being instrument validity. AR(1) and AR(2) are tests for 1st order and 2nd order serial correlation. "a" denotes significance at the 5% level and "b" at the 10% level.

Figure 2: Plots of estimated residuals from SYS-GMM





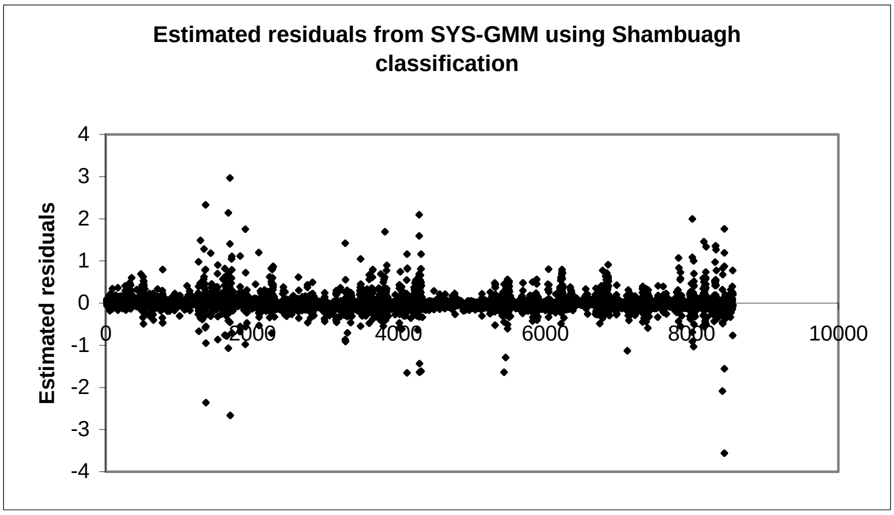


Table 4: Matrix of Correlation Coefficients (using IMF classification for exchange rate dummies)

	tbt_{ijt}	y_{it}	y_{jt}	FTA_{ijt}	dis_{ij}	$COMLAN_{ij}$	COL_{ij}	$LANG_{ij}$	$RXRVOL_{ijt}$	$CU\text{-}CU_{ijt}$	$FIX\text{-}FIX_{ijt}$	$CU\text{-}FLT_{ijt}$	$CU\text{-}FIX_{ijt}$	$FIX\text{-}FLT_{ijt}$	$DFIX_{ij}$
tbt_{ijt}	1.000														
y_{it}	-0.001	1.000													
y_{jt}	0.036	0.058	1.000												
FTA_{ijt}	0.040	0.138	0.205	1.000											
dis_{ij}	-0.134	-0.126	-0.224	-0.610	1.000										
$COMLAN_{ij}$	0.083	0.057	0.174	0.185	-0.465	1.000									
COL_{ij}	0.093	0.004	0.010	-0.054	0.013	0.0576	1.000								
$LANG_{ij}$	-0.058	-0.004	0.060	-0.016	0.013	0.287	0.323	1.000							
$RXRVOL_{ijt}$	-0.088	-0.123	-0.146	-0.334	0.463	-0.235	0.053	-0.053	1.000						
$CU\text{-}CU_{ijt}$	0.027	0.093	0.047	0.238	-0.202	0.101	-0.038	-0.017	-0.246	1.000					
$FIX\text{-}FIX_{ijt}$	0.089	0.084	0.008	0.045	-0.356	0.202	-0.064	-0.007	-0.327	-0.053	1.000				
$CU\text{-}FLT_{ijt}$	0.0010	0.096	0.172	0.008	0.030	-0.005	-0.004	-0.014	-0.025	-0.046	-0.081	1.000			
$CU\text{-}FIX_{ijt}$	0.0023	0.058	-0.028	0.052	0.022	-0.046	-0.027	-0.028	0.044	-0.033	-0.059	-0.050	1.000		
$FIX\text{-}FLT_{ijt}$	-0.005	-0.077	-0.075	-0.127	0.145	-0.083	-0.015	-0.073	0.161	-0.148	-0.263	-0.225	-0.163	1.000	
$DFIX_{ijt}$	-0.088	0.082	-0.027	0.199	-0.111	-0.002	-0.023	-0.075	0.052	-0.072	-0.126	-0.110	-0.008	-0.356	1.000

Table 5: Mean FDI flows and shares of World FDI flows and GDP by country

Country	Mean Total FDI Flows	Proportion of World FDI flows (%)	Proportion of World GDP
United States	6,304.03	22.15%	29.80%
United Kingdom	4,598.98	8.01%	4.66%
Belgium-Luxembourg	2,528.65	7.28%	0.81%
Germany	2,519.92	5.36%	6.31%
Netherlands	2,474.61	4.09%	1.16%
France	2,212.81	5.41%	4.39%
Japan	1,483.85	0.68%	16.05%
Switzerland	1,083.64	1.38%	0.86%
Canada	955.78	3.14%	3.32%
Sweden	744.18	2.19%	0.79%
Spain	738.59	3.37%	1.83%
Italy	612.30	1.39%	3.68%
Australia	553.49	1.76%	1.19%
Ireland	454.41	1.54%	0.23%
Denmark	359.71	1.12%	0.53%
Finland	297.87	0.60%	0.39%
Singapore	228.67	1.76%	0.22%
Norway	213.35	0.47%	0.52%
Hong Kong	162.86	1.75%	0.48%
Austria	148.97	0.56%	0.63%
Portugal	136.60	0.51%	0.33%
Korea	117.71	0.57%	1.32%
New Zealand	114.01	0.45%	0.17%
Greece	40.94	0.21%	0.38%
Israel	32.70	0.32%	0.32%
Iceland	5.86	0.02%	0.03%
United Arab Emirates	3.36	n/a	0.21%
Average	1,078.81	2.72%	2.78%
Sum	8,904,785	76.1%	80.72%

Notes: The first column is calculated for each country using an average of the FDI values against all of the country-pairs used in our sample, US$ million, constant 2000 prices. Data Source: OECD International Direct Investment Statistics. In the second column the contribution to World FDI flows is calculated using data on the net FDI flow of each country and the World over the period 1980-2003. Data source: World Bank World Development Indicators. The third column is the percentage of World output accounted for by each country's real GDP over the period 1980 to 2003. Data source: World Bank World Development Indicators.

Table 6: Summary of mean FDI flows by exchange rate regime, US$ million, constant prices (2000=100)

Regime Type / Year	R+R				IMF				Shambaugh			
	Currency Union countries	Floating currency countries	Floating currency countries (ex UK & USA)	Fixed Exchange Rate Agreement	Currency Union countries	Floating currency countries	Floating currency countries (ex UK & USA)	Fixed Exchange Rate Agreement	Currency Union countries	Floating currency countries	Floating currency countries (ex UK & USA)	Fixed Exchange Rate Agreement
1980	141.23	1,271.98	32.35	165.67	141.23	420.70	38.82	278.34	141.23	221.61	29.13	279.43
1981	126.71	1,206.80	46.78	88.32	126.71	298.78	52.24	250.00	126.71	193.33	58.00	148.92
1982	120.02	1,061.84	59.30	116.28	120.02	323.50	56.11	221.23	120.02	178.79	57.72	103.00
1983	119.85	1,077.15	106.31	144.09	119.85	438.99	56.83	216.66	119.85	215.78	60.85	140.12
1984	132.61	1,506.57	123.65	199.80	132.61	570.61	55.97	245.57	132.61	329.71	51.91	207.79
1985	109.57	1,579.61	152.18	104.73	109.57	503.02	43.79	190.33	109.57	318.26	45.46	124.98
1986	239.82	2,718.57	234.24	222.30	239.82	811.54	85.16	424.27	239.82	424.78	79.03	482.06
1987	303.65	4,098.42	467.46	353.02	303.65	1,251.85	160.67	459.89	303.65	709.07	109.75	379.83
1988	527.53	6,587.18	1,165.47	384.76	527.53	1,524.19	283.94	946.93	527.53	763.00	154.58	683.27
1989	781.28	8,305.63	1,396.07	530.60	781.28	2,155.53	350.81	1,324.87	781.28	1,003.85	208.09	1,286.76
1990	1,188.26	6,599.16	1,920.02	782.96	1,188.26	1,244.79	362.48	2,410.45	1,188.26	973.89	263.68	1,518.84
1991	996.22	4,320.48	1,140.18	732.64	996.22	846.91	242.07	1,713.96	996.22	692.08	164.72	1,209.27
1992	1,153.66	2,711.66	795.19	761.88	1153.66	673.71	140.63	1,683.35	1153.66	603.06	156.92	1,875.81
1993	802.12	3,637.37	439.35	617.83	802.12	1,019.86	173.65	927.76	802.12	878.89	235.29	1,396.41
1994	771.80	2,862.61	337.68	722.13	771.80	941.19	186.33	992.03	771.80	831.57	190.44	1,514.21
1995	1,230.93	4,679.36	719.39	1017.83	1,230.93	1,154.83	207.21	1,564.13	1,230.93	1,144.75	316.56	2,147.56
1996	1,140.29	4,867.16	584.63	989.52	1,140.29	1,499.02	190.13	1,187.90	1,140.29	1,312.31	229.36	1,568.22
1997	1,241.70	4,372.07	565.68	1,117.17	1,241.70	1,713.98	219.96	1,265.04	1,241.70	1,689.11	334.00	1,396.95
1998	1,919.70	8,214.38	755.26	1,670.88	1,919.70	5,280.79	424.30	1,722.10	1,919.70	3,551.76	229.36	2,050.06
1999	3,481.22	7,953.40	521.80	2,492.23	3,481.22	6,918.53	611.82	788.05	3,481.22	5,063.71	324.00	742.04
2000	6,433.85	5,695.841	817.59	4,287.67	6,433.85	5,648.35	843.19	1,791.55	6,433.85	4,124.89	484.07	1,815.40
2001	3,815.71	1,977.694	264.76	2,546.62	3,815.71	3,785.62	287.99	1,240.10	3,815.71	2,776.79	145.86	1,266.79
2002	2,886.00	1,764.84	264.33	1,826.40	2,886.00	2,569.09	324.34	592.15	2,886.00	1,911.86	226.91	575.68
2003	2,616.10	2,303.269	461.18	1,933.85	2,616.10	3,075.63	516.28	824.65	2,616.10	2,411.41	273.66	764.34

Table 7: Re-estimation excluding USA and UK

Variable	R+R	IMF	Shambaugh
constant	8.246[a]	7.476[a]	7.674[a]
	(19.70)	(17.41)	(16.71)
tbt_{ijt}	0.00009	0.008	0.009[b]
	(0.02)	(1.41)	(1.69)
y_{it}	0.075[a]	0.082[a]	0.090[a]
	(2.01)	(2.80)	(2.96)
y_{jt}	-0.016	0.013	-0.008
	(-0.91)	(0.77)	(-0.43)
FTA_{ijt}	-0.024	0.012	-0.004
	(-0.89)	(0.41)	(-0.14)
dis_{ij}	-0.041[a]	-0.011	-0.021
	(-2.45)	(-0.62)	(-1.27)
$COMLAN_{ij}$	0.238[a]	0.220[a]	0.223[a]
	(3.27)	(3.50)	(3.44)
COL_{ij}	-0.005	0.030	0.006
	(-0.07)	(0.65)	(0.10)
$LANG_{ij}$	0.042	0.049[b]	0.052
	(1.40)	(1.65)	(1.61)
$RXRVOL_{ijt}$	-0.058	0.086	0.203
	(-0.19)	(0.28)	(0.62)
$CU\text{-}CU_{ijt}$	0.220[a]	0.356[a]	0.370[a]
	(3.48)	(4.69)	(5.02)
$FIX\text{-}FIX_{ijt}$	-0.057	0.140[a]	0.116[a]
	(-1.55)	(3.12)	(2.64)
$CU\text{-}FLT_{ijt}$	0.009	0.114[a]	0.124[a]
	(0.25)	(2.82)	(3.93)
$CU\text{-}FIX_{ijt}$	-0.089[a]	-0.017	0.008
	(-2.40)	(-0.72)	(0.25)
$FIX\text{-}FLT_{ijt}$	-0.063	-0.024	0.011
	(-1.54)	(-1.17)	(0.72)
$DFIX_{ijt}$	-0.092[a]	-0.068[a]	-0.030[b]
	(-2.30)	(-3.02)	(-1.90)
Diagnostics from the short-run model			
Wald test: all regressors~χ^2(df)	1780.58[a] (35)	1669.99[a] (35)	1608.59[a] (35)
Wald test: time dummies~χ^2(df)	71.11[a] (22)	76.48[a] (22)	68.97[a] (22)
Wald test: exchange rate dummies~χ^2(df)	30.62[a] (6)	44.42[a] (6)	35.46[a] (6)
J-test~χ^2(df)	281.35 (550)	284.24 (551)	281.87 (551)
AR(1) test~N(0,1)	-4.09[a]	-4.07[a]	-4.08[a]
AR(2) test~N(0,1)	-0.17	-0.17	-0.13
Number of observations	6,177	6,177	6,177
R^2	0.579	0.580	0.578

Notes: Estimates derived from GMM estimates of a short-run model (time dummies also included). Robust t-ratios (reported in parentheses) are approximated using the delta method. The Wald tests are for exclusion restrictions on the coefficients, the number of restrictions is reported in parentheses. J-test denotes the Hansen test for over-identifying restrictions (the null hypothesis is instrument validity). AR(1) and AR(2) are tests for 1st and 2nd order serial correlation. "a" denotes significance at the 5% level and "b" at the 10% level.

Table 8: Extended model (using GDP)

Variable	R+R	IMF	Shambaugh
constant	4.568^a	3.621^a	3.386^a
	(5.70)	(3.73)	(3.57)
tbt_{ijt}	0.030^a	0.032^a	0.030^a
	(3.32)	(3.64)	(3.55)
y_{it}	0.058^a	0.064^a	0.064^a
	(3.75)	(3.90)	(4.16)
y_{jt}	0.072^a	0.075^a	0.080^a
	(5.03)	(4.63)	(4.57)
FTA_{ijt}	0.054	0.022	0.048
	(1.14)	(0.57)	(1.10)
dis_{ij}	-0.036	-0.026	-0.017
	(-1.23)	(-1.10)	(-0.71)
$COMLAN_{ij}$	0.113	0.108	0.103
	(1.38)	(1.39)	(1.34)
COL_{ij}	0.054	0.070	0.085
	(0.48)	(0.69)	(0.78)
$LANG_{ij}$	0.106^a	0.080^b	0.091^a
	(2.48)	(1.87)	(2.04)
$RXRVOL_{ijt}$	0.057	1.080^b	1.545^a
	(0.10)	(1.73)	(2.54)
$LRRXRVOL_{ijt}$	0.349	1.435	1.076
	(0.34)	(1.40)	(0.94)
$(infi \times inf_j)_t$	0.064^a	0.090^a	0.093^a
	(3.50)	(3.94)	(4.08)
$CU\text{-}CU_{ijt}$	-0.092	0.239^a	0.300^a
	(-0.74)	(3.07)	(4.13)
$FIX\text{-}FIX_{ijt}$	-0.369^a	0.027	0.094^b
	(-3.01)	(0.48)	(1.93)
$CU\text{-}FLT_{ijt}$	-0.116	0.105^b	0.151^a
	(-1.09)	(1.79)	(3.18)
$CU\text{-}FIX_{ijt}$	-0.294^a	-0.021	0.042
	(-3.04)	(-0.41)	(0.73)
$FIX\text{-}FLT_{ijt}$	-0.347^a	-0.077^a	0.035
	(-2.66)	(-2.19)	(1.25)
$DFIX_{ijt}$	-0.259^a	-0.062^b	0.039
	(-2.44)	(-1.67)	(0.90)
Diagnostics from the short-run modeb			
Wald test: all regressors~χ^2(df)	2535.30^a (37)	2770.39^a (39)	2381.71^a (37)
Wald test: time dummies~χ^2(df)	76.00^a (22)	70.36^a (22)	70.72^a (22)
Wald test: exchange rate dummies~χ^2(df)	26.58^a (6)	24.94^a (6)	27.13^a (6)
J-test~χ^2(df)	329.72 (643)	322.76 (641)	324.04 (640)
AR(1) test~N(0,1)	-3.55^a	-3.55^a	-3.54^a
AR(2) test~N(0,1)	-0.25	-0.20	-0.16
Number of observations	7,270	7,270	7,270
R^2	0.701	0.699	0.700

Notes: A short-run dynamic model was estimated using a two-step system GMM procedure (time dummies also included but not reported), from which the above long-run coefficients are derived. Robust t-ratios (reported in parentheses) are approximated using the delta method. The Wald tests are for exclusion restrictions on the coefficients, the number of restrictions is reported in parentheses. J-test denotes the Hansen J-test for over-identifying restrictions, the null hypothesis being instrument validity. AR(1) and AR(2) are tests for 1st and 2nd order serial correlation. "a" denotes significance at the 5% level and "b" at the 10% level.

Data Appendix

Variable	Description	Source
fdi	Log of total real bi-lateral FDI flows. Total bi-lateral flows are the sum of inward and outward FDI flows. All series are expressed in US dollars and converted to 2000 prices using the US GDP deflator (Source: UN common database). Converting to logs required adding a positive constant to the series, given that FDI values can be negative, reflecting disinvestment.	OECD International Direct Investment Statistics*
tbt	Log of total real bi-lateral trade. Total bi-lateral trade is the sum of exports plus imports. Series expressed in US dollars and in constant prices.	IMF Direction of Trade Statistics*
y_i/y_j	Log of real per capita GDP or real GDP, expressed in US dollars.	UN common database*
RXRVOL	Real exchange rate volatility, defined as the annual standard deviation of the monthly percentage changes in the real bi-lateral exchange rate.	IMF International Financial Statistics/ OECD Main Economic Indicators*
LRRXRVOL	Measure of long-run exchange rate volatility: the standard deviation of the log monthly changes in the real exchange rates using observations from the preceding five years.	
FTA	Dummy variable that takes the value of 1 if both countries share a free trade agreement, 0 otherwise.	World Trade Organisation.
$inf_i \times inf_j$	Product of informational flows, proxied by the number of fixed telephone lines per 1000 people.	World Bank, World Development Indicators*
dis	Log of geographic distance between the two countries. For each country pair, we use weighted distances between the biggest cities of the two countries, *"inter-city distances being weighted by the share of the city in the overall country's population"* (Clair et al., 2004, pp. 4-5)[29].	Centre d'Etudes Prospectives et d'Informations Internationales http://www.cepii.fr/
LANG	Dummy variable that takes the value of 1 if both countries share the same official language, 0 otherwise.	
COL	Dummy variable that takes the value of 1 if both countries have ever had a colonial relationship, 0 otherwise.	
COMLAN	Dummy variable that takes the value of 1 if both countries have a common land border, 0 otherwise.	
CU-CU	Dummy variables that takes the value of 1 if both countries share the same currency.	Calculated from classifications produced by Reinhart and Rogoff (2004), Shambaugh (2004) and various issues of the IMF's *Annual Report on Exchange Rate Arrangements and Restrictions*
FIX-FIX	Dummy variables that takes the value of 1 if both countries have a fixed exchange rate either directly or indirectly.	
CU-FLT	Dummy variables that takes the value of 1 if one country is a member of a currency union and the other has a floating exchange rate.	
CU-FIX	Dummy variables that takes the value of 1 if one country is a member of a currency union and the other has a fixed exchange rate.	
FIX-FLT	Dummy variables that takes the value of 1 if one country has a fixed exchange rate and the other is floating.	
DFIX	Dummy variables that takes the value of 1 if both countries follow a fixed exchange rate policy but they peg to different anchor currencies.	

Notes: * indicates that the series was taken from http://www.esds.ac.uk.

[29] Clair, G., Gaulier, G. Mayer, T. and Zignago, S. (2005), 'Notes on CEPII's distance measures'. *Source*: http://www.cepii.fr/.