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CURRENCY UNIONS AND INTERNATIONAL INTEGRATION

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ABSTRACT

This paper characterizes the integration patterns of international currency unions (such as the CFA Franc zone and the East Caribbean Currency Area). We empirically explore different features of currency unions, and compare them both to countries with sovereign monies, and to regions within nations. We ask: are countries within international currency unions as integrated as regions within political unions? We do this by examining the criteria for Mundell's concept of an optimum currency area. We find that members of currency unions are more integrated than countries with their own currencies, but less integrated than regions within a country. For instance, we find that currency union members have more trade and less volatile real exchange rates than countries with their own monies, but less trade and more volatile exchange rates than regions within individual countries. Similarly, business cycles are more highly synchronized across currency union countries than across countries with sovereign monies, but not as synchronized as regions of a single country. Finally, currency union membership is not associated with significantly greater risk sharing, though risk sharing is widespread within countries.

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1. Introduction: Currency Unions and "Home Bias"

Is "dollarization" associated with enhanced international economic integration?[1] We examine the behavior of countries that are or have been members of international currency unions, and ask whether existing currency unions replicate the desirable features of optimal currency areas as set out by Mundell (1961). Specifically, we ask whether the countries and political units that constitute currency unions are as integrated economically as regions within nations. We find that while a common currency enhances economic integration, the degree of integration is far smaller than within nations

A number of studies have shown that national borders inhibit economic integration. Internal trade is disproportionately large compared to international trade; relative prices are more stable inside countries than across national boundaries; domestic assets tend to be held disproportionately, and so forth. Perhaps the large size of this "border effect" is mostly the result of exchange rate volatility or, more generally, the consequence of having different national moneys. The objective of this paper is to investigate this hypothesis.

This paper is empirical. Our strategy is to exploit data on the many existing currency unions. We differentiate between intranational *political unions* (i.e., sovereign states with a single currency but also common laws, political environments, cultures, and so forth), and international *currency unions* (i.e., sovereign countries that have delegated monetary policy to some international or foreign authority but retain sovereignty in other domains). The United States, France, and the United Kingdom are examples of political unions. Behavior of regions within these countries is the focus of the emerging literature on intranational economics (Hess and van Wincoop (2000), Bachetta, Rose and van Wincoop, 2001). The CFA Franc Zone, and the East Caribbean Currency Area are examples of currency unions.

Our approach is to ask whether currency unions exhibit the type of economic integration that Mundell (1961) argues is desirable for an "optimum currency area". We measure a number of economic characteristics for international monetary unions, intranational political unions and other countries. Mundell's framework implies that the gains from a common currency are proportional to the size of international transactions. Using disaggregated international trade data, we find that currency unions are more open and more specialized than non-currency union countries of comparable size. More directly, we examine international trade patterns. Using a gravity equation, we find that trade between members of a currency union (e.g., Brunei and Singapore) is indeed much higher than trade between comparable countries with their own currencies, by a factor of over three. However, even this sizable effect is small in comparison with the "home market bias" which shows that intranational trade is higher than international trade by a factor of almost twenty, even for units of comparable economic size. That is, our estimates show that a hypothetical country which is as large (in terms of population, GDP, geographic area and so forth) as Brunei and Singapore combined would engage in much more intranational trade than Brunei and Singapore do in reality.

We examine real exchange rates and deviations from purchasing power parity.[2] The volatility of real exchange rates is lower for members of currency unions than for countries with independent currencies. But much of this effect stems from the fact that no currency union has experienced a hyperinflation; low inflation countries with sovereign currencies have real exchange rate volatility that is only modestly higher than that of currency union members. Currency union members do not have detectably different rates of mean-reversion in their real exchange rates. Compared to the benchmark of exchange rates between cities in comparably sized countries, currency unions exhibit slightly more integrated prices.

We also investigate other characteristics of currency unions. We find that business cycles are systematically more highly correlated between members of currency unions than between countries with sovereign currencies, but not as much as regions of a single country. Finally, we examine risk sharing between members of currency unions and countries with independent currencies, by examining consumption and income, and find only a small impact of currency union on risk sharing.[3]

We conclude that members of a common currency area are more economically integrated than non-currency union members, but not nearly as much as those that are fully politically integrated. That is, dollarized countries are more likely to satisfy Mundell's criteria for being members of an optimum currency area, but not nearly as much as regions within a single country.

International trade entails foreign exchange transactions, unless it occurs between members of common currency areas. While we ordinarily think of such costs as being small (at least for OECD countries facing deep liquid foreign exchange markets), avoiding it seems to have non-trivial consequences.[4] So, currency unions may encourage integration. We are concerned with the association between integration and currency unions. We do not consider whether causality flows from integration to currency union (integrated countries are more likely to join and remain in currency unions), in the reverse direction (currency union induces integration), or both.[5]

In section 2 below, we provide a gross characterization of currency union members, taking special note of their openness and specialization. We analyze the impact of currency union membership on international trade in section 3, and the impact on prices in the section that follows. Section 5 examines the international synchronization of business cycles, while section 6 looks at risk sharing. The paper concludes with a brief summary and conclusion.

2. Characterizing Currency Union Members

We begin our analysis of common currency areas by providing an aggregate description of their members.

2.1 Openness

Our first (macroeconomic) data set consists of annual observations for 210 “countries” between 1960 and 1996 extracted from the 1998 World Bank *World Development Indicators* (WDI) CD-ROM. The list of countries is tabulated in the appendix Table A2. This data set includes all countries, territories, colonies and other entities covered by WDI (all are referred to as “countries” for simplicity), and is extremely comprehensive.[6] The data set has been checked and corrected for mistakes.

In this data set, some 1891 (country-year) observations (24% of the sample) were members of a common currency area; the list of countries is tabulated in Table A1. We include: members of common currency areas (such as Benin, a member of the CFA franc zone); countries which operated without a sovereign currency (such as Panama which uses the US dollar); long-term 1:1 fixers where there is substantial currency substitution and essentially no probability of a move from parity (such as the Bahamas); and colonies, dependencies, overseas territories/departments/collectivities (such as Guadeloupe). Anchor countries (such as the US and France), whose currencies are used by others, are tabulated solely for reference (i.e., they are not included as currency-union members in our empirical analysis).[7]

Table 1 shows some descriptive statistics for both the whole sample of available observations, and for (periphery) currency union members. The number of available observations is tabulated along with the means and standard deviation. There is also a *p*-value for a *t*-test of equality of means for currency union members and non-members.

Table 1 indicates that members of currency unions tended to be poorer and smaller than non-currency union members. Currency unions are associated with lower and more stable inflation. However, they have lower ratios of M2 to GDP (a standard measure of financial depth), which may be because they tend to be poor. A better indicator of their financial markets may be the fact that the spread of the domestic loan rate above LIBOR tends to be lower (even after one has excluded high inflation observations). The country-specific standard deviation of the output growth rate, a crude measure of output volatility, seems to be similar for currency union members and non-members. Finally, there is little indication that currency unions are associated with either more or less fiscal discipline.

What of openness? Currency unions are more open than countries with their own currencies. Both exports and imports are larger as percentages of GDP to a degree that is both statistically significant and economically important. Interestingly, while export duties are lower, import duties are higher for currency union members, as is the importance of trade taxes. This is probably because most currency union members have poorly developed income and value added tax bases. Currency union members run current accounts that are larger (in absolute value) as a percentage of GDP, and also more variable. Currency unions are also more open to private capital flows, and to foreign direct investment. That is, both the intertemporal and the intratemporal evidence indicate that currency union members are more open to capital than non-members.

Succinctly, members of currency unions seem to be more open to international flows of goods, services, and capital than countries with their own currencies. But one can overstate the importance of these differences. Currency union members tend to be small countries, which are well known to be more open than larger countries. Accordingly, in section 3 we control for size and income in determining whether membership in a common currency area is systematically associated with more intense trade.

2.2 Specialization

Given that members of currency unions are more open to international influences than countries with their own currencies, it is natural to ask if members of common currency areas are also more specialized and therefore potentially more vulnerable to asymmetric industry shocks. Kenen (1969) first discussed specialization in this context.

One way to examine this question would be to compare *production* structures and see if currency union members are more specialized in production. However the data set necessary to examine this question does not exist. Nevertheless, it is possible to examine the patterns of specialization exhibited by countries engaging in *international trade*. To examine specialization patterns manifest in international trade, we exploit the "World Trade Data Base" (WTDB), the second data set that we exploit extensively in this paper.

The WTDB is a consistent recompilation of United Nations trade data, discussed in Feenstra, Lipsey and Bowen (1997).[8] The WTDB is estimated to cover at least 98% of all trade. Annual observations of nominal trade values (recorded in thousands of American dollars) are available in the WTDB for some 166 countries from 1970 through 1995; the countries in the WTDB data set are tabulated in Appendix Table A3.[9] These observations are available at the four-digit ("sub-group") Standard International Trade Classification (SITC) level (revision 2). There are a total of 897,939 observations in this three-dimensional panel (goods x countries x years). A typical observation is the exports (totalling $740,000) from South Africa of SITC good 11 in 1970.[10]

For each country-year observation, we compute the Herfindahl index, a measure of specialization. The Herfindahl index is the sum of squared shares of the individual goods, defined as:

$$H_{it} \equiv \sum_j (x_{ijt} / X_{it})^2 \qquad j = 1, \ldots, J$$

where x_{ijt} denotes the exports for country i of SITC subgroup j in year t, X_{it} denotes total exports for i in year t, and the summation is taken over all SITC subgroups. H is bounded by (0,1]; a high value of H indicates that the country is specialized in the production of a few goods.

We have some 3,045 country-year observations of the Herfindahl index for the WTDB. Of these, 388 (some 13%) are for countries that are members of currency unions. As Table 2a shows, Herfindahl indices for countries with their own currencies are systematically lower (averaging .23) than those for members of currency unions (which average .31). That is, members of common currency areas tend to be more specialized. The difference is not only of economic importance; it is also statistically significant (the t-test for a difference in means is 5.7). Currency union members also export (122) fewer sub-goods on average than countries with their own currencies, consistent with the hypothesis of greater specialization (again, the difference is statistically significant with a t-statistic of 17.7).

It might be objected that currency union members are smaller and poorer than other countries, so that more specialization is to be expected. We control for these other factors by regressing the Herfindahl index on the Penn World Table (mark 5.6) measure of real GDP per capita, population, and a dummy variable that is unity if the country-year observation is for a currency union member. The results are tabulated in Table 2b. They show that our conclusions are insensitive to the addition of controls for real GDP per capita, country size, and either country- or time-specific fixed effects. Currency union members consistently have higher Herfindahl indices and export smaller numbers of goods.[11]

To summarize, members of currency unions are more open than countries with their own currencies. They are also more specialized.[12]

3. Trade Integration

In this section of the paper, we show that members of currency unions systematically engage in more international trade. This question is of obvious interest since the benefits from using a single money in terms of saved transactions costs depend on the amount of trade between two regions, as recognized since at least Mundell (1961). We follow Rose (2000) in using a "gravity" model of international trade as our framework. In particular, we ask whether bilateral trade between two countries is higher if they both use the same currency, holding constant a variety of other determinants of international trade.

The large literature which employs the gravity model of international trade points to distance, income levels and country size as being the most critical drivers of bilateral trade flows, a result which we corroborate here. The precise model we employ is completely standard and can be written:

$$\ln(X_{ij}) = gCU_{ij} + b_0 + b_1 \ln(D_{ij}) + b_2 \ln(Y_iY_j / Pop_iPop_j) + b_3 \ln(Y_iY_j) + d \bullet Z_{ij} + e_{ij}$$

where X_{ij} denotes the value of bilateral trade between countries i and j, CU is a binary dummy variable which is unity if i and j use the same currency and zero otherwise, D_{ij} denotes the distance between countries i and j, Y denotes real GDP, Pop denotes Population, Z denotes a vector of other controls, the b and d coefficients are nuisance coefficients, and e denotes the residual impact of all other factors driving trade . The coefficient of interest to us is g, which measures the impact of a common currency on international trade. A positive coefficient indicates that two countries that use a common currency also tend to trade more.

We begin by estimating this equation using 1995 data from the WTDB, augmented by data from the UN *International Trade Statistics Yearbook*. Over 150 countries, dependencies, territories, overseas departments, colonies, and so forth (referred to simply as "countries" below) for which the United Nations Statistical Office collects international trade data are included in the data set. Country location (used to calculate Great Circle distance) is taken from the CIA's web site, which also provides observation for other variables of interest such as: contiguity, official language, colonial background, area, and so forth.[13] Real GDP and population are taken from the 1998 World Bank *World Development Indicators* CD-ROM.[14]

Estimation results are contained in Table 3. OLS is used, and robust standard errors are recorded parenthetically. At the extreme left of the table, the simplest gravity model is employed; that is, no auxiliary Z's are included. The b coefficients indicate that the gravity model works well, in two senses. First, the coefficient estimates are sensible and strong. Greater distance between two countries lowers trade, while greater economic "mass" (proxied by real GDP and GDP per capita) increases trade. These intuitive and plausible effects are in line with the estimates of the literature; they are also of enormous statistical significance with t-statistics exceeding 20 (in absolute value). Second, the equation fits the data well, explaining a high proportion of the cross-sectional variation in trade patterns.

While it is reassuring that the gravity model performs well, its role is strictly one of auxiliary conditioning. We are interested in understanding the relationship between currency union membership and trade flows after accounting for gravity effects. Even after taking out the effects of output, size, and distance, there is a large effect of a common currency on trade. The point estimates indicate that two countries that share a common currency trade together by a factor of $\exp(1.88) \cong 6.5$! This effect is not only economically large, but also statistically significant at traditional confidence levels (the t-statistic is 3.3).

One can think of a number of reasons for this strong result. At the top of the list would be model mis-specification, implying that the currency union variable is picking up the effect of some other omitted variable(s). But this hunch is mistaken; the results are robust. Four different perturbations of the gravity model are included in Table 3; they augment the basic results with extra (Z) controls. These extra effects are usually statistically significant and economically sensible, though they add little to the overall explanatory power of the model. Being partners in a regional trade agreement, sharing a common language, having the same (post-1945) colonizer, being part of the same nation (as e.g., France and an overseas department like French Guiana), and having had a colonizer-colony relationship all increase trade by economically and statistically significant amounts. Landlocked and large countries tend to trade less; islands trade more. But inclusion of these extra controls does not destroy the finding of an economically large and statistically significant positive g. While the coefficient falls somewhat with extra controls, the lowest estimate of g in Table 3 indicates that trade is some 285% higher for members of a common currency than for countries with sovereign currencies.

Rose (2000) estimated a number of gravity equations with a comparable data set spanning 1970 through 1990, and found similar results; his point estimate of g was 1.2. He also showed his results to be robust to: the exact measurement of CU, the exact measure of distance, the inclusion of extra controls, sub-sampling, and different estimation techniques.

To summarize: members of a currency union trade more, *ceteris paribus*. A reasonable estimate is that trade is three times as intense for members of a common currency area as for countries with their own currencies. While this estimate seems provocatively high, it is actually quite low compared with the well-documented size of "home bias" in international trade. McCallum (1995) and Helliwell (1998) find home bias in goods markets to be on the order of 12x to 20x, far greater than our estimates here. While membership in a common currency area does intensify trade, it does not intensify it nearly enough for common currency areas to resemble countries.

4. Price Integration

Obstfeld and Rogoff (1996, p. 633) mention two of the main benefits of currency union as:

- Reduced accounting costs and greater predictability of relative prices for firms doing business in both countries and
- Insulation from monetary disturbances and speculative bubbles that might otherwise lead to temporary unnecessary fluctuations in real exchange rates (given sticky nominal prices)

In this section, we explore whether real exchange converge in currency unions are more stable in the sense of converging more quickly and having lower short-run volatility. To answer the first question, we estimate the equation

$$qroot_{ij} = a + bCU_{ij} + d \bullet Z_{ij} + e_{ij}\,.$$

Here, $qroot_{ij}$ is the estimated autoregressive coefficient in an AR1 regression for the (log of the) real exchange rate of country i relative to country j. A large value of $qroot_{ij}$ indicates slow adjustment of the real exchange rate. CU_{ij} is a dummy variable that takes the value of one if countries i and j were in a currency union for the entire post-1960 period, and a zero otherwise. Z_{ij} is a vector of auxiliary conditioning variables (such as the distance between countries i and j, the volatility of the nominal exchange rate, etc.) that are included in the regression as controls, but that are not directly of interest to us. e_{ij} is a random error that contains factors that affect the speed of adjustment of real exchange rates that are not included in our regression.

We hypothesize that b_{ij} is negative: that the persistence of real exchange rates is lower for currency union countries. If currency unions are successful in their objective of reducing real exchange rate volatility, one measure of success is the speed at which real exchange rates converge to equilibrium.

Our real exchange rate data is based on annual consumer price indices and exchange rates from our World Bank macroeconomic data set. For each country in the data set, we first estimate an AR1 regression (with intercept, given that the price data is in index form) for (log) real exchange rates from 1960-1996.[15] We use the slope coefficient in these time-series regressions as the regressand in the cross-section regression defined above.

The results reported in Table 4 indicate no support for the hypothesis that real exchange rates adjust more quickly in currency unions. The first column of the table reports results for the basic regression. In addition to the currency union dummy variable, the regression contains the log of distance (in miles) between countries *i* and *j*; a dummy variable for whether *i* and *j* are divisions of the same country (e.g., metropolitan France and Guadeloupe); the standard deviation of the first difference of the log of the nominal exchange rate; and a constant. The currency union dummy variable has a positive sign, but is not statistically significant at conventional levels.

The other variables in the regression are not directly of interest to us, but we note that two variables are highly significant in this and each of our other specifications: the same-country dummy, and the nominal exchange rate volatility. As we expect, the coefficient on the same-country dummy is negative, indicating that real exchange rates adjust more quickly for these pairs. Also unsurprisingly, the speed of adjustment is significantly faster when nominal exchange rate volatility is higher. Transitory real exchange rate volatility is closely associated with volatile nominal exchange rates. When shocks to nominal exchange rates are very large and lead to large misalignments of real exchange rates, there is rapid adjustment.

The other specifications in Table 4 introduce other control variables (not reported in the table.) The second column introduces average inflation rates in countries *i* and *j*; their presence does not appreciably alter the effect of the other regressors. The third column includes all of the control variables as the second column, but also includes a dummy variable for each country. In this specification, the currency union dummy variable is significant, but with a positive sign. That is, real exchange rates are more persistent in currency-union countries. The fourth and fifth regressions reported in Table 4 control for high inflation in alternative manners. The regression in the fourth column includes the maximum annual inflation rate of each country, while the regression of the fifth column is identical to the base specification reported in column 1 but excludes all countries that have experienced high inflations. (High inflation is defined here as average inflation that exceeds 100 per cent.) We find the coefficient on the currency union dummy is not changed under these specifications. The bottom line from Table 4 is that being a member of a currency union does not increase the speed of adjustment of real exchange rates.

There is a logical inconsistency in the approach taken in Table 4. We want to allow for differing speeds of adjustment of real exchange rates. But if we model the real exchange rate of countries i and j as an AR1, and the real exchange rates of countries i and k as an AR1 with a different speed of adjustment, then the real exchange rate of countries j and k cannot follow an AR1. More generally, we would want to model the adjustment of the real exchange rate of two countries i and j as depending not only on its own lags, but on the lags of real exchange rates of countries that are economically integrated.

To handle this problem, Table 5 reports results from first-order VARs of real exchange rates for groups of countries. The real exchange rates for members of currency unions are grouped together. Countries that are not members of currency unions are grouped by continent. The statistic reported in Table 5 for each group of countries is the largest (in absolute value) eigenvalue of the matrix of coefficients from the first-order VAR. The largest eigenvalue ultimately determines the persistence in the vector of real exchange rates. The larger is this eigenvalue, the more slowly the group adjusts.

There are several advantages to using this statistic to measure the speed of adjustment of real exchange rates. First, as we have alluded to, the VAR specification does not suffer from the inconsistency that modeling all real exchange rates as AR1s does. Second, as is well known, OLS estimation is efficient even with errors that are correlated across real exchange rates. Third, while the real exchange rates for each group of countries are all calculated relative to a base country, the eigenvalues are independent of the choice of base country.

If the real exchange rate system is stationary, the asymptotic distribution of the largest eigenvalue is standard. But we cannot be certain of stationarity. Even if the system is stationary, we cannot be certain that the asymptotic distribution is reliable in small samples. So we undertake Monte Carlo and bootstrap exercises. The most straightforward null hypothesis to test for each VAR is that all real exchange rates in the group are simple random walks.[16]

The test statistics are reported in Table 5a. First, notably, these tests do not have enough power to reject the unit root hypothesis for the vast majority of country groups. Only for the European group and the ECCA currency union can we reject the unit-root null at the 95 percent level of confidence, and then only with the bootstrap test. We never reject the unit-root null with the Monte Carlo test. Moreover, Table 5 reveals little difference in the persistence of real exchange rates among currency-union and non-currency-union groups. There is no clear difference in the persistence (as measured by the largest eigenvalue), or in the p-values of the test statistics for the unit root null.

Table 5b reports similar statistics for groups of cities within each of seven countries. This data set is monthly, in contrast to the country-level data that are annual. So, the measures of the speed of adjustment are not comparable.[17] But, even with the city data, there is only one country for which we can clearly reject the unit-root null: Canada.

To sum up, Tables 4 and 5 suggest that the speed of adjustment of real exchange rates is not clearly related to monetary union, or even political union. This result is perhaps not surprising. The literature has found mixed results concerning the speed of adjustment of prices within countries and across borders. Parsley and Wei (1996) find that prices converge rapidly between cities in the U.S. The speed of convergence is much greater than is typically found for real exchange rates between countries (see Rogoff (1996).) But, their data is for prices of very narrowly defined goods (as opposed to the aggregate price indexes used in international comparisons), and they have no comparable data for countries other than the U.S. In contrast, Rogers and Jenkins (1995) and Engel, Hendrickson and Rogers (1997) find no significant difference between intranational and international speeds of convergence of aggregate real exchange rates.

In contrast, there is a well-known "border" effect for short-term volatility of real exchange rates. For example, Engel and Rogers (1996) find that U.S.-Canadian relative prices are far more volatile than relative prices between cities within each country, even taking into account distance between cities. We ask here whether currency unions have a similar effect in reducing real exchange rate volatility. In Table 6 we report results from regressions of the form:

$$qvol_{ij} = a + bCU_{ij} + d \bullet Z_{ij} + e_{ij} .$$

Here, $qvol_{ij}$ is a measure of the volatility of the real exchange rate of countries i and j. We use as our measure the standard deviation of the residual from the AR1 regressions discussed above. This measures the volatility of shocks to real exchange rates, as distinct from variance arising from slow adjustment. As before, CU_{ij} is a dummy variable that takes the value of one if countries i and j were in a currency union. Z_{ij} is a vector of other variables that are included in the regression as controls, and e_{ij} is a random error.

The regression specifications across the five columns of Table 6 are identical to those of Table 4, except that the regressand is the *volatility* of the real exchange rate rather than its *persistence*. In all specifications, the currency union dummy variable is negative and is highly significant in all but the last. The specification that appears most plausible here is the third specification, which contains dummy variables for each country. In this regression, the log of distance has a positive and significant sign, indicating that more distant countries have greater real exchange rate volatility. The variance of the change in the (log) nominal exchange rate is a highly significant variable in this regression (and all others.) Our interest is focussed on the currency union dummy, which is very statistically significant: being a member of a currency union reduces the standard deviation of annual real exchange rates by 6 percentage points.

We conclude that real exchange rates have much lower short-term volatility among currency-union countries, even holding constant the volatility of the nominal exchange rate. That is, the reduction in real exchange rate variance is not solely attributable to fixed exchange rates; currency-union membership appears to stabilize real exchange rates through other channels as well. But, real exchange rate volatility of currency union members is still higher on average than for cities within countries. The average annual standard deviation of real exchange rates among currency union countries in our sample is 3.6 percent, compared to 1.1 percent for city pairs within the seven countries listed in the lower panel of Table 5.

5. Business Cycle Synchronization

We now examine whether countries that use the same currency tend to have more highly synchronized business cycles. This has been a natural question to ask since Mundell (1961);

countries with highly synchronized business cycles forego little monetary independence if they share a common currency. Thus countries with highly synchronized business cycles have a higher propensity to adopt a common currency. Of course, since a common monetary policy also eliminates idiosyncratic monetary policy, causality flows in the reverse direction. That is, members of a common currency union should experience more synchronized business cycles since they do not experience national monetary policy shocks. Rather than try to determine either part of the relationship structurally, we are simply interested here in seeing whether members of a common currency area in fact experience more synchronized business cycles. It is especially interesting to us since we have already found that currency union members are quite specialized in international trade, making them potentially subject to asymmetric shocks.

The regressions we estimate take the form:

$$Corr(s)_{ij} = a + bCU_{ij} + d \bullet Z_{ij} + e_{ij}$$

where: $Corr(s)_{ij}$ denotes the estimated correlation between real GDP for country i and real GDP for country j de-trended with method s, CU is a binary dummy variable which is unity if countries i and j are members of the same currency union, a and d are nuisance coefficients, Z is a vector of controls, and e denotes omitted residual factors. The coefficient of interest to us is b; a positive b indicates that two countries with a common currency tend to have more tightly correlated business cycles. Since our analysis is reduced-form in nature, we are not able to tell whether countries with more tightly synchronized business cycles tend to belong to common currency areas, or whether membership in a currency union tends to synchronize business cycles (or both).

In forming the regressand, we take advantage of our macroeconomic data set (the list of potential countries is tabulated in Table A1). In particular for each pair of countries in the sample, we estimate the bivariate correlation between de-trended annual real GDP for countries *i* and *j* over the sample period 1960-1996 (or the maximum available span of data).[18] We use two different time-series models to de-trend the data: (country-specific) first-differences of natural logarithms; and a log-linear time trend model. After (the natural logarithm of) each country's real GDP has been de-trended, we then estimate simple bivariate correlations between the de-trended GDP series.[19]

Results are tabulated in Tables 7a and 7b. Table 7a contains results where the regressand is constructed from GDP series de-trending via growth rates; Table 7b is the analogue with linear de-trending.

The extreme left column of each of the tables presents a simple OLS regression of business cycle synchronization on the currency union dummy variable. We find a positive b coefficient, indicating that business cycles are more highly synchronized for countries that trade more. The size and statistical significance of the estimate depends on the de-trending method employed.

Six perturbations of the basic model are also displayed in Tables 7a and 7b to check the sensitivity of the analysis. The first five perturbations (all estimated with OLS) simply add extra control regressors to the right hand side of the equation (i.e., extra Z's). We choose the five different sets of regressors used in Table 3, (this encompasses the controls used by Clark and van Wincoop (2000); other controls sets, including country fixed effects, deliver similar results). Robust t-statistics are displayed in parentheses.

The estimates in the tables indicate that business cycles are in fact more tightly synchronized for members of a currency union. The exact point estimate depends on both the de-trending method and the exact set of auxiliary regressors. But the coefficient is consistently positive and almost always statistically significant at conventional levels. Being a member of a common currency area increases international business cycle correlations by perhaps .1, an economically significant amount.[20]

In the extreme right column, the natural log of bilateral trade between countries i and j is used as the sole control regressor, following Frankel and Rose (1998). This is an important test of the model, since Clark and van Wincoop find that inclusion of trade as a control destroys the border effect. When trade is included, its coefficient is estimated with IV, using the first nine regressors of the gravity equation as instrumental variables.[21] Trade appears to have a strong positive effect on business cycle synchronization. This result twins well with the literature. For instance, Frankel and Rose (1998) found that increased international trade induces more tightly synchronized business cycles, using data for the OECD; our result is consistent with theirs. However, controlling for trade does not destroy the significance of b.

To summarize, countries that are members of a common currency union tend to have more highly synchronized business cycles; the correlation is perhaps .1 higher on average for currency union members than for non-members. While economically and statistically significant, the size of this effect is small in an absolute sense. Most recently, Clark and van Wincoop (2000) compare the coherences of business cycles within countries and across countries, using annual data for both employment and real GDP. They show that intranational business cycle correlations are approximately .7 for regions within countries, but in the range of (.2,.4) for comparable regions drawn across countries. That is, the effect of international borders on business cycle synchronization ranges between .3 and .5. Thus, only a small part of the "border effect" is explained by membership in a common currency area.

6. Risk sharing

In this section, we turn to international risk sharing. It is well known that the apparent degree of international risk sharing is low. In a classic contribution, Feldstein and Horioka (1980) found that national saving and investment rates are highly correlated, apparently inconsistent with international risk sharing. Alternatively, if risk-sharing opportunities were widespread, there should be little country-specific idiosyncratic consumption risk. As Backus, Kehoe and Kydland (1992) noted, consumption should be more highly correlated across countries than output in the presence of risk sharing. In fact, the data show the opposite. Furthermore, as French and Poterba (1991) and others have reported, there is strong home bias in asset holdings. There seems to be very little international diversification of portfolios.

Obstfeld and Rogoff (2000) have argued that international risk sharing might be diminished in the presence of transactions costs. Specifically, they cite costs of trading goods (rather than assets) as an impediment to risk sharing. They also note that these costs might conceivably be related to the need to make foreign exchange transactions in order to buy and sell goods internationally. In other words, countries that are members of currency unions might do more risk sharing.

We run a cross-section regression of the form:

$$ccorr_{ij} = a + bCU_{ij} + d \bullet Z_{ij} + e_{ij}.$$

where, $ccorr_{ij}$ is calculated as the correlation of the first difference in the log of consumption per capita for country i with the analogue for country j. The right-hand-side of the regression is of the same generic form as the regressions of the previous two sections. Thus, CU_{ij} is a dummy variable which is unity if countries i and j were in a currency union; Z_{ij} is a vector of control variables; and e_{ij} is a random error. The consumption data in this section is taken from the Penn World Tables, and is adjusted for purchasing power parity. The data are annual, and the maximum data span available is 1960-1992.[22]

Table 8 reports the regression results. If risk sharing is greater among currency unions, we expect a positive coefficient on the currency union dummy. If more distant countries find it more difficult to share risks, we also expect a negative coefficient on the log of distance. We report results from six regressions. All regressions include the currency union dummy and log distance as explanatory variables. The first regression (reported in the first column) uses a single intercept. The second regression uses a comprehensive set of country-specific fixed effects, so that both the dummies for *i* and *j* take on a value of one when the regressand is $ccorr_{ij}$. The third regression is identical to the first regression, but is estimated with weighted least squares.[23] The second set of three regressions repeats the analysis, but augments the regression with the bivariate correlation between the growth rates of output (that is, the correlation of the first difference in the log of output for country *i* with the analogue for country *j*, the analogue to the regressand).

The results are weak. The log of distance always enters significantly with the correct sign. The currency union dummy always enters with the correct sign. However, it is not significant in the first specification; it is only of marginal significance in the second; and it is highly significant only in the third. In all three estimates, the economic size of the effect of currency unions is small. For instance, the currency union effect is to increase the consumption correlation by .04 percentage points with weighted least squares. Since the intercept term in the regression is 0.31, then ignoring the effect of distance (that is, for two countries whose log distance is zero), being in a currency union raises the consumption correlation from 0.31 to 0.35.

Even these modest results may overstate the risk sharing opportunities within currency unions. A high correlation of consumption for a pair of countries may not actually reflect greater risk sharing opportunities between those two countries. It may simply reflect less idiosyncratic risk. That is, the consumption of two countries may be correlated simply because their output is correlated. Thus, even in the absence of avenues for risk sharing, there may be a high consumption correlation that should not be interpreted as indicating substantial international risk sharing.

This concern is particularly relevant since in the previous section we found that business cycles are more highly correlated for currency union countries. So controlling for the degree of output correlation is a potentially important robustness check. We pursue this by adding the actual correlation of (detrended) GDP per capita as a control in the right-hand columns of Table 8. As it turns out, the output correlation coefficient is always statistically and economically significant as a control variable, but its presence has little effect on our estimate of β.

To summarize, we have found little statistically and economically significant evidence that international risk sharing is enhanced by membership in a currency union. This is perhaps unsurprising, given the absence of substantive international fiscal transfer arrangements and the shallow private financial markets of most currency union members.

7. Conclusion

This paper contributes to the dollarization dialogue by quantifying some of the features associated with common currencies, *using actual data*. Using the historical record, we have found that the extra degree of integration associated with a common currency is substantial but finite. Members of international currency unions tend to experience more trade, less volatile exchange rates, and more synchronized business cycles than do countries with their own currencies. Of course, since well-integrated countries are more likely to adopt a common currency, some of these integration "effects" of currency union may be illusory. That is, the causality may flow from integration to currency union rather than the reverse. In any case, while members of international currency unions are more integrated than countries with their own monies, they remain far from integrated compared with the intranational benchmark of regions within a country.

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Table 1: Descriptive Macroeconomic Statistics and Measures of Openness

	---- Whole Sample ----			--- Currency Unions ---			
	Obs.	**Mean**	**St.Dev.**	**Obs.**	**Mean**	**St.Dev.**	**Test of Equality (p-val.)**
Real GDP per capita ($)	2454	5285	5262	416	3615	4474	.00
Population (millions)	5102	23.6	9.3	1052	1.8	2.7	.00
Inflation (%)	4152	40.3	499	672	7.8	9.0	.00
M2/GDP (%)	3197	38.0	23.9	510	30.4	16.7	.00
Loan Rate – LIBOR (%)	2131	72.7	2643	412	5.2	6.9	.24
Loan Rate – LIBOR (%) (inflation<100%)	1858	7.6	13.3	348	5.4	7.2	.00
Output Growth Rate volatility (std dev, %)	211	6.1	5.5	51	5.9	3.1	.17
Budget Deficit (% GDP)	2289	-3.6	5.8	268	-3.7	6.1	.84
Exports (% GDP)	4732	32.3	23.7	783	39.8	23.5	.00
Imports (% GDP)	4729	37.8	25.4	783	53.2	27.1	.00
Export Duties (% exports)	1621	3.4	6.1	237	2.6	3.8	.00
Import Duties (% imports)	2226	12.3	9.6	241	18.0	8.4	.00
Trade Taxes (% Revenue)	2252	19.5	17.1	300	31.9	20.1	.00
Current Account (% GDP)	2942	-4.5	11.5	477	-8.3	13.3	.00
|Current Account| (% GDP)	2942	7.3	10.0	477	10.8	11.4	.00
Gross FDI (% GDP)	2058	1.5	2.6	339	2.0	3.4	.00
Private Capital Flows (% GDP)	2067	12.0	31.6	352	22.4	67.6	.00

Table 2a: Measures of Specialization

		- Herfindahl Index -		- Number Exports -	
	Obs.	**Mean**	**Std. Dev.**	**Mean**	**Std. Dev.**
Non-Currency Union Members	2657	.23	.24	254	132
Currency Union Members	388	.31	.19	132	89

Table 2b: Regression-Based tests of Specialization

	--------------- Regressors		-----------------	
Regressand:	**Real GDP per capita**	**Population**	**Currency Union**	**Controls**
Herfindahl Index	-.10 (6.8)	-2.8 (20.2)	.06 (4.4)	
Herfindahl Index	.05 (2.4)	-2.8 (3.9)	.12 (4.1)	**Country Controls**
Herfindahl Index	.10 (6.8)	-2.7 (18.8)	.05 (4.4)	**Time Controls**
Number of Exports	.02 (23.9)	.0003 (24.3)	-67.2 (11.9)	
Number of Exports	.0002 (0.4)	-.00006 (2.2)	-28.5 (1.8)	**Country Controls**
Number of Exports	.018 (25.4)	.0003 (26.4)	-60.9 (11.4)	**Time Controls**

Absolute values of robust t-statistics recorded in parentheses. Intercepts not reported.
Sample size = 2,806 throughout.
* Coefficients for real GDP per capita (population) multiplied by 10^4 (10^7) for convenience.

Table 3: Gravity Models of International Trade for 1995

Currency Union	1.88 (.46)	1.37 (.42)	1.06 (.42)	1.19 (.37)	1.37 (.38)
(Log) Distance	-1.38 (.04)	-1.24 (.04)	-1.23 (.04)	-1.18 (.04)	-1.19 (.04)
(Log Product) Real GDP per capita	.76 (.02)	.73 (.02)	.74 (.02)	.61 (.02)	.51 (.02)
(Log Product) Real GDP	.87 (.01)	.89 (.01)	.91 (.01)	.99 (.02)	1.08 (.02)
Regional Trade Agreement		1.08 (.16)	1.04 (.15)	.88 (.15)	1.00 (.15)
Common Language		.82 (.07)	.56 (.07)	.72 (.07)	.64 (.07)
Common Land Border		-.19 (.18)	-.08 (.19)	.17 (.19)	.19 (.19)
Common Colonizer			.81 (.13)	.52 (.13)	.47 (.13)
Same Nation			.79 (.66)	.79 (.65)	.64 (.66)
Colonial Relationship			1.68 (.14)	1.43 (.14)	1.42 (.15)
Number of Landlocked Countries				-.62 (.06)	
(Log of) Sum of Land Area				-.25 (.02)	
(Log of) Product of Land Area					-.18 (.01)
Number of Island Countries					.14 (.05)
R^2	.71	.72	.72	.74	.74
RMSE	1.757	1.724	1.703	1.663	1.656

OLS estimation. Robust standard errors recorded in parentheses. Intercepts not recorded.
Sample size = 4493. Regressand is log of bilateral trade.

Table 4: Real Exchange Rate Persistence and Currency Unions

Currency Union	.03 (1.0)	.01 (0.5)	.10 (3.9)	.01 (0.3)	-.00 (0.1)
(Log) Distance	-.00 (0.5)	.00 (0.0)	.02 (0.5)	.01 (0.2)	-.00 (0.4)
Same Nation	-.12 (3.3)	-.11 (3.9)	-.06 (3.3)	-.11 (4.2)	-.10 (4.5)
Nominal Exchange Rate Volatility	-.13 (18.0)	-.22 (11.4)	-.16 (3.3)	-.26 (21.2)	-.28 (13.2)
Intercept	.90 (34.4)	.89 (34.3)		.90 (34.6)	.92 (34.4)
Number of observations	3647	3647	3647	3647	3236
Controls		Inflation Controls	Country Dummies, Inflation	Max. Inflation	Without High Inflation Countries

Absolute values of robust t-statistics recorded in parentheses.
Regressand is estimated root from autoregression of log real exchange rate.

Table 5a: Real Exchange Rate VARs and Currency Unions

Country grouping	Currency union?	Principal Root	p-value (Monte Carlo)	p-value (bootstrap)
Africa	No	0.998	0.82	0.71
Asia	No	1.005	0.79	0.66
Europe	No	0.947	0.07	0.01
North America	No	1.031	0.90	0.86
South America	No	0.940	0.25	0.13
Oceania	No	0.950	0.69	0.62
Belgium-Lux.	Yes	0.944	0.77	0.30
Britain-Ireland	Yes	0.876	0.73	0.56
Bhutan-India	Yes	0.648	0.37	0.34
France	Yes	0.915	0.67	0.21
South Africa	Yes	0.882	0.55	0.40
ECCA	Yes	0.753	0.10	0.01
USA 1	Yes	1.036	0.97	0.94
USA 2	Yes	1.166	0.99	0.99
CFA 1	Yes	1.071	0.99	0.99
CFA 2	Yes	0.990	0.75	0.63

Table 5b: Real Exchange Rate Convergence between Cities within Countries

Country	Principal Root	p-value (Monte Carlo)	p-value (bootstrap)
USA	0.977	0.16	0.10
Canada	0.980	0.01	0.01
Mexico	0.986	0.09	0.05
Germany	0.985	0.14	0.06
Italy	0.993	0.43	0.26
Spain	0.993	0.39	0.34
Switzerland	0.976	0.09	0.07

Table 6: Real Exchange Rate Volatility and Currency Unions

Currency Union	-.04 (5.9)	-.02 (3.4)	-.06 (7.9)	-.02 (3.3)	-.01 (0.8)
(Log) Distance	-.005 (2.2)	-005 (2.8)	.005 (6.1)	-.006 (3.5)	-.000 (0.1)
Same Nation	.05 (1.5)	.04 (1.7)	.00 (0.4)	.04 (1.8)	.02 (1.5)
Exchange Rate Volatility	.28 (27.5)	.40 (24.4)	.11 (4.5)	.41 (31.2)	.48 (39.6)
Intercept	.12 (7.2)	.11 (6.9)		.11 (7.8)	.05 (5.0)
Number of observations	3647	3647	3647	3647	3236
		Inflation Controls	Country Dummies, Inflation	Max. Inflation	Without High Inflation Countries

Absolute values of robust t-statistics recorded in parentheses.

Table 7a: Business Cycle Synchronization and Currency Unions
Real GDP de-trended via growth rates

Currency Union	.05 (1.4)	.10 (1.8)	.07 (1.3)	.11 (2.0)	.11 (2.0)	.10 (1.8)	.11 (2.0)
(Log) Distance		-.04 (8.1)	-.02 (4.3)	-.02 (4.4)	-.02 (4.5)	-.02 (4.2)	
(Log Product) Real GDP per capita		.04 (14.2)	.04 (13.0)	.03 (12.5)	.03 (11.5)	.04 (12.5)	
(Log Product) Real GDP		.00 (2.5)	.00 (2.5)	.00 (1.6)	.00 (0.9)	-.00 (1.0)	
Regional Trade Agreement			.13 (5.7)	.14 (6.2)	.14 (6.2)	.14 (6.5)	
Common Language			.01 (1.5)	.03 (2.9)	.03 (2.8)	.03 (2.8)	
Land Border			.06 (1.9)	.05 (1.8)	.05 (1.7)	.04 (1.5)	
Common Colonizer				-.08 (5.5)	-.08 (5.3)	-.06 (4.5)	
Same Nation				.12 (1.2)	.12 (1.2)	.13 (1.3)	
Colonial Relationship				-.05 (1.8)	-.05 (1.8)	-.04 (1.4)	
Number of Landlocked Countries					.00 (0.0)		
(Log of) Sum of Land Area					.00 (0.6)		
(Log of) Product of Land Area						.00 (2.2)	
Number of Island Countries						-.02 (2.9)	
(Log of) Bilateral Trade							.02 (12.5)
RMSE	.262	.235	.234	.233	.233	.233	.241

Regressand is bilateral correlation of real GDPs (1960-1996), de-trended by first-difference of natural logs.
OLS estimation, except for last column (IV with first 10 regressors as instrumental variables).
Absolute robust t-statistics recorded in parentheses. Intercepts not recorded.
Sample size = 4419, except for bivariate regression where sample size = 5913.
Regressand is bivariate correlation of real GDPs 1960-1996, de-trended via growth rates.

Table 7b: Business Cycle Synchronization and Currency Unions
Real GDP de-trended via linear time trend

Currency Union	.14 (2.5)	.13 (2.1)	.08 (1.3)	.15 (2.4)	.15 (2.3)	.10 (1.6)	.15 (2.2)
(Log) Distance		-.04 (4.4)	-.02 (1.9)	-.02 (.01)	-.02 (2.4)	-.02 (2.3)	
(Log Product) Real GDP per capita		.09 (17.1)	.08 (16.2)	.08 (15.8)	.09 (15.2)	.12 (18.8)	
(Log Product) Real GDP		-.01 (2.9)	-.01 (2.6)	-.01 (3.5)	-.02 (4.4)	-.04 (9.2)	
Regional Trade Agreement			.10 (2.5)	.11 (3.1)	.12 (3.2)	.13 (3.5)	
Common Language			.06 (2.9)	.09 (4.5)	.08 (3.9)	.08 (3.9)	
Land Border			.10 (2.0)	.09 (1.8)	.07 (1.6)	.04 (0.9)	
Common Colonizer				-.16 (5.5)	-.14 (5.0)	-.10 (3.4)	
Same Nation				-.19 (1.1)	-.19 (1.1)	-.15 (0.9)	
Colonial Relationship				-.09 (1.6)	-.07 (1.3)	-.04 (0.7)	
Number of Landlocked Countries					-.01 (0.5)		
(Log of) Sum of Land Area					.02 (2.9)		
(Log of) Product of Land Area						.03 (8.4)	
Number of Island Countries						-.04 (3.4)	
(Log of) Bilateral Trade							.02 (9.4)
RMSE	.447	.449	.448	.447	.446	.442	.464

Regressand is bilateral correlation of real GDPs (1960-1996), de-trended by time trend.
OLS estimation, except for last column (IV with first 10 regressors as instrumental variables).
Absolute robust t-statistics recorded in parentheses. Intercepts not recorded.
Sample size = 4419, except for bivariate regression where sample size = 5913.
Regressand is bivariate correlation of real GDPs 1960-1996, de-trended via time trend.

Table 8: Risk Sharing and Currency Unions: Consumption Correlations

Currency union	.05 (0.9)	.10 (1.8)	.04 (4.13)	.07 (1.2))	.11 (1.9)	.03 (3.9)
Log of Distance	-.03 (6.3)	-.04 (7.9)	-.03 (39.9)	-.02 (3.4)	-.03 (5.9)	-.02 (22.9)
Constant	.29 (7.8)		.31 (49.1)	.15 (4.3)		.39 (166.2)
Output Correlation				.28 (19.4)	.19 (12.3)	.16 (26.3)
	OLS	Country Dummies	Weighted Least Squares	OLS	Country Dummies	Weighted Least Squares

Absolute value of robust t-statistics reported in parentheses

Table A1: Members of Monetary Unions with WDI Data
(* denotes country treated as anchor in multilateral currency unions)

CFA Franc Zone
Benin
Burkina Faso*
Cameroon
Central African Republic
Chad
Comoros
Congo Rep.
Cote d'Ivoire
Equatorial Guinea
Gabon
Guinea-Bissau
Mali
Niger
Senegal
Togo

USA
American Samoa
The Bahamas
Bermuda
Guam
Liberia
Marshall Islands
Micronesia Fed. Sts.
Northern Mariana Islands
Palau
Panama
Puerto Rico
Virgin Islands (U.S.)

France
French Guiana
Guadeloupe
Martinique
Mayotte
Monaco
New Caledonia
Reunion

ECCA
Antigua and Barbuda
Dominica
Grenada
St. Kitts and Nevis
St. Lucia*
St. Vincent and the Grenadines

South Africa
Lesotho
Namibia
Swaziland

UK
Channel Islands
Ireland
Isle of Man

Australia
Kiribati
Tonga

West Africa
Kenya*
Tanzania
Uganda

France* and Spain
Andorra

India
Bhutan

Singapore
Brunei

Norway
Faeroe Islands

Denmark
Greenland

Switzerland
Liechtenstein

Belgium
Luxembourg

Israel
West Bank and Gaza

Table A2: Countries in Macroeconomic Data Set

Afghanistan
Albania
Algeria
American Samoa
Andorra
Angola
Antigua and Barbuda
Argentina
Armenia
Aruba
Australia
Austria
Azerbaijan
The Bahamas
Bahrain
Bangladesh
Barbados
Belarus
Belgium
Belize
Benin
Bermuda
Bhutan
Bolivia
Bosnia and Herzegovina
Botswana
Brazil
Brunei
Bulgaria
Burkina Faso
Burundi
Cambodia
Cameroon
Canada
Cape Verde
Cayman Islands
Central African Republic
Chad
Channel Islands
Chile
China
Colombia
Comoros
Congo Dem. Rep.
Congo Rep.
Costa Rica
Cote d'Ivoire
Croatia
Cuba
Cyprus
Czech Republic
Denmark
Djibouti
Dominica
Dominican Republic
Ecuador
Egypt Arab Rep.
El Salvador
Equatorial Guinea
Eritrea
Estonia
Ethiopia
Faeroe Islands
Fiji
Finland
France
French Guiana
French Polynesia
Gabon
The Gambia
Georgia
Germany
Ghana
Greece
Greenland
Grenada
Guadeloupe
Guam
Guatemala
Guinea
Guinea-Bissau
Guyana
Haiti
Honduras
Hong Kong China
Hungary
Iceland
India
Indonesia
Iran Islamic Rep.
Iraq
Ireland
Isle of Man
Israel
Italy
Jamaica
Japan
Jordan
Kazakhstan
Kenya
Kiribati
Korea Dem. Rep.
Korea Rep.
Kuwait
Kyrgyz Republic
Lao PDR
Latvia
Lebanon
Lesotho
Liberia
Libya
Liechtenstein
Lithuania
Luxembourg
Macao
Macedonia FYR
Madagascar
Malawi
Malaysia
Maldives
Mali
Malta
Marshall Islands
Martinique
Mauritania
Mauritius
Mayotte
Mexico
Micronesia Fed. Sts.
Moldova
Monaco
Mongolia
Morocco
Mozambique
Myanmar
Namibia
Nepal
Netherlands
Netherlands Antilles
New Caledonia
New Zealand
Nicaragua
Niger
Nigeria
Northern Mariana Islands
Norway
Oman
Pakistan
Palau
Panama
Papua New Guinea
Paraguay
Peru
Philippines
Poland
Portugal
Puerto Rico
Qatar
Reunion
Romania
Russian Federation
Rwanda
Samoa
Sao Tome and Principe
Saudi Arabia
Senegal
Seychelles
Sierra Leone
Singapore
Slovak Republic
Slovenia
Solomon Islands
Somalia
South Africa
Spain
Sri Lanka
St. Kitts and Nevis
St. Lucia
St. Vincent and the Grenadines
Sudan
Suriname
Swaziland
Sweden
Switzerland
Syrian Arab Republic
Tajikistan
Tanzania
Thailand
Togo
Tonga
Trinidad and Tobago
Tunisia
Turkey
Turkmenistan
Uganda
Ukraine
United Arab Emirates
United Kingdom
United States
Uruguay
Uzbekistan
Vanuatu
Venezuela
Vietnam
Virgin Islands (U.S.)
West Bank and Gaza
Yemen Rep.
Yugoslavia FR (Serbia/Montene
Zambia
Zimbabwe

Table A3: Countries in World Trade Data Bank

Afghanistan
Albania
Algeria
Angola
Argentina
Australia
Austria
Bahamas
Bahrain
Bangladesh
Barbados
Belize
Benin
Bermuda
Bhutan
Bolivia
Brazil
Brunei
Bulgaria
Burkina Faso
Burundi
Cambodia
Cameroon
Canada
Cayman Islands
Central African Rep.
Chad
Chile
China
Colombia
Comoros
Congo
Costa Rica
Cote D'Ivoire
Cuba
Cyprus
Denmark
Djibouti
Dominican Rep
Ecuador
Egypt
El Salvador
Eq. Guinea
Ethiopia
Faeroe Islands
Fiji
Finland
France
French Guiana
Gabon
Gambia
Germany West
Ghana
Greece
Greenland
Grenada
Guadeloupe
Guatemala
Guinea
Guinea Bissau
Guyana
Haiti
Honduras
Hong Kong
Hungary
Iceland
India
Indonesia
Iran
Iraq
Ireland
Israel
Italy
Jamaica
Japan
Jordan
Kenya
Kiribati
Korea
Korea North
Kuwait
Laos
Lebanon
Liberia
Libya
Madagascar
Malawi
Malaysia
Maldives
Mali
Malta
Martinique
Mauritania
Mauritius
Mexico
Mongolia
Morocco
Mozambique
Myanmar (Burma)
Nepal
Netherlands
Netherlands Antilles
New Caledonia
New Zealand
Nicaragua
Niger
Nigeria
Norway
Oman
Pakistan
Panama
Papua New Guinea
Paraguay
Peru
Philippines
Poland
Portugal
Qatar
Reunion
Romania
Rwanda
Saudi Arabia
Senegal
Seychelles
Sierra Leone
Singapore
Solomon Islands
Somalia
South Africa
Spain
Sri Lanka
St. Kitts & Nevis
St. Lucia
St. Vincent & Grenadines
States
Sudan
Surinam
Sweden
Switzerland
Syria
Taiwan
Tanzania
Thailand
Togo
Trinidad & Tobago
Tunisia
Turkey
Uganda
UK
United States
United Arab Emirates
Uruguay
Venezuela
Vietnam
Western Samoa
Yemen North
Yugoslavia
Zaire
Zambia
Zimbabwe

Endnotes

[1] We define "dollarization" as a situation where a country does not have its own sovereign money; the currency it uses need not be a dollar (US or other).
[2] McKinnon (1963) has argued that in practice real exchange rate behavior does not appreciably depend on the choice of monetary regime, and the desire to influence real exchange rate behavior is not a justification for having an independent currency.
[3] We disregard labor mobility since it is so difficult to construct an appropriate data set, and since monetary policy can only be used to offset transitory nominal shocks where labor movement is probably inappropriate. We also ignore asset and financial market integration.
[4] Our investigation is in the spirit of Obstfeld and Rogoff (2000) who urge the profession to examine the consequences of (presumably small) costs of international trade. Frankel and Rose (1998) raise the possibility that the degree of integration among economies (and hence their suitability for membership in a currency union) might increase upon the formation of a common currency area.
[5] It is difficult to examine the direction of causality since currency unions are long-lived. Rose (2000) provides more analysis which supports the idea that currency union tends to promote trade integration rather than the reverse.
[6] There are however many missing observations for variables of interest.
[7] In the case of multilateral currency unions, there is no clear anchor
[8] This has been augmented with data from the UN's *International Trade Statistics Yearbook*.
[9] The specialization data set includes usable observations for the following countries: Algeria, Angola, Argentina, Australia, Austria, Bahamas, Bahrain, Bangladesh, Barbados, Belgium, Belize, Benin, Bhutan, Bolivia, Brazil, Bulgaria, Burkina Faso, Burundi, C.A.R., Cameroon, Canada, Chad, Chile, China, Colombia, Comoros, Congo, Costa Rica, Cyprus, Czechoslovakia, Denmark, Djibouti, Dominican Rep., Ecuador, Egypt, El Salvador, Ethiopia, Fiji, Finland, France, Gabon, Gambia, Germany East, Germany West, Ghana, Greece, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, Hong Kong, Hungary, Iceland, India, Indonesia, Iran, Iraq, Ireland, Israel, Italy, Ivory Coast, Jamaica, Japan, Jordan, Kenya, Korea, Kuwait, Laos, Liberia, Madagascar, Malawi, Malaysia, Mali, Malta, Mauritania, Mauritius, Mexico, Mongolia, Morocco, Mozambique, Myanmar, Nepal, Netherlands, New Zealand, Nicaragua, Niger, Nigeria, Norway, Oman, Pakistan, Panama, Papua N. Guinea, Paraguay, Peru, Philippines, Poland, Portugal, Qatar, Reunion, Romania, Rwanda, Saudi Arabia, Senegal, Seychelles, Sierra Leone, Singapore, Solomon Is., Somalia, South Africa, Spain, Sri Lanka, St. Kitts & Nevis, Sudan, Suriname, Sweden, Switzerland, Syria, Taiwan, Tanzania, Thailand, Togo, Trinidad & Tobago, Tunisia, Turkey, U.A.E., U.K., U.S.A., U.S.S.R., Uganda, Uruguay, Venezuela, Yemen, Yugoslavia, Zaire, Zambia, and Zimbabwe.
[10] SITC Code 11 denotes "Animals of the Bovine Species, incl. Buffaloes, live." Other examples of 4-digit sub-groups include: "Tyres, pneumat. new, of a kind used on buses, lorries" (SITC code 6252), and "Int. combustion piston engines for marine propuls." (SITC code 7133).
[11] Our findings are not affected by the inclusion of quadratic terms for income as in Imbs and Wacziarg (2000).
[12] This specialization makes them more vulnerable to industry-specific shocks, and might be expected to increase the idiosyncratic nature of their business cycles.
[13] The 1998 *World Factbook* available at http://www.odci.gov/cia/publications/factbook/index.html.
[14] We sometimes include a control for common membership in a regional free trade agreement. We include a number of such agreements, including: the EU; the Canada-US FTA; EFTA; the Australia/New Zealand closer economic relationship; the Israeli/US FTA; ASEAN; CACM; PATCRA; CARICOM; SPARTECA; and the Cartagena Agreement, all taken from the WTO's web site (http://www.wto.org/wto/develop/webrtas.htm).
[15] We only estimate the AR1 if there are at least fifteen observations for each country.
[16] Under the Monte Carlo experiment, we assume that the errors have a Normal distribution, with a covariance matrix equal to the sample covariance of the first-differences of the (logs) of the real exchange rates. We measure the frequency with which the largest eigenvalue is smaller than the estimated largest eigenvalue for each group, based on 5000 replications with sample sizes equal to the sample size of our data. Under the bootstrap experiment, we use the first differences of the actual (log) real exchange rates to construct our bootstrap sample. We sample, with replacement, the vector of real exchange rate changes at each date, thus maintaining the structure of correlation across real exchange rates within each group. As with the Monte Carlo statistics, we measure the frequency with which the largest eigenvalue is smaller than the estimated largest eigenvalue for each group, based on 5000 replications with sample sizes equal to the sample size of our data.

[17] It might be natural to compare the eigenvalues by raising the city-level eigenvalues to the twelfth power, but that would only be a rough approximation given that the annual CPI data is average for the year, not end-of-period.
[18] We only estimate the bilateral correlation if we have at least five matching GDP observations for each country.
[19] Thus, we first separately de-trend Afghani and Australian real GDP with linear time trend models. Then we estimate the correlation between the two de-trended real GDPs over time (the actual correlation is -.002). We then repeat this procedure for all possible country pairs, resulting in a vector of correlations. De-trending via taking deviations of growth rates (first-differences of natural logarithms) from the average (country-specific) growth rate yields another measure of the regressand. For regressors, we use the same set of regressors used in the gravity model of trade. That is, we model business cycle synchronization as being a function of the distance between the countries, the product of their real GDPs, the product of their real GDP per capitas, and so forth.
[20] As a robustness check, we have substituted the correlation between labor forces for the correlation between GDPs (employment, unemployment, and industrial production data are simply not available for many countries even at the annual frequency). This regressand also delivers a consistently positive, statistically significant effect of currency union on business cycle coherence.
[21] This is necessary because while trade may effect business cycle synchronization, it is equally plausible that causality flows in the reverse direction, as pointed out by Frankel and Rose (1998).
[22] Again, we only estimate the bilateral correlation if we have at least fifteen matching observations for each country.
[23] Specifically, we give proportionately greater weight to observations in which the correlation is based on more data. That is, when we can base a correlation on thirty-two years of data, that correlation in the cross-section regression receives double the weight of a correlation based on only sixteen years of data.