TRADE, FINANCE, SPECIALIZATION, AND SYNCHRONIZATION

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Abstract—I investigate the determinants of business cycle synchronization across regions. The linkages between trade in goods, financial openness, specialization, and business cycle synchronization are evaluated in the context of a system of simultaneous equations. The main results are as follows. (i) Specialization patterns have a sizable effect on business cycles. Most of this effect is independent of trade or financial policy, but directly reflects differences in GDP per capita. (ii) A variety of measures of financial integration suggest that economic regions with strong financial links are significantly more synchronized, even though they also tend to be more specialized. (iii) The estimated role of trade is closer to that implied by existing models once intra-industry trade is held constant. The results obtain in a variety of data sets, measurement strategies, and specifications. They relate to a recent strand of international business cycle models with incomplete markets and transport costs and, on the empirical side, point to an important omission in the list of criteria defining an optimal currency area, namely, specialization patterns.

I. Introduction

THE interactions between trade openness, financial integration, specialization, and business cycle synchronization are complex. In theory, trade both in goods and in financial assets may affect the cross-country synchronization of business cycles. Intense bilateral trade will tend to accompany highly correlated business cycles in a wide range of theoretical models, ranging from multisector international models with intermediate-good trade, to one-sector versions with either technology or monetary shocks.[1] The impact of financial integration on cycle synchronization, in turn, is not unambiguous. On the one hand, limited ability to borrow and lend internationally hampers the transfer of resources across countries and can increase GDP correlations. But on the other hand, if investors have imperfect information or face liquidity constraints, limiting capital flows can actually decrease GDP correlations, as investors herd, or withdraw capital from many destinations simultaneously.[2] Specialization, finally, is likely to affect the international synchronization of business cycles directly. This will naturally occur in the presence of sector-specific shocks, as two economies producing the same types of goods will then be subjected to similar stochastic developments. But it may also happen if sectors differ in their response to monetary shocks—for instance, because of different market structures or labor market arrangements. Then countries with similar production patterns will be synchronized even though shocks are purely aggregate.[3]

Theory also points to potentially important indirect interactions. It is for instance well known that openness to goods trade results in specialization.[4] Similarly, financial liberalizations may induce specialization, as access to an increasing range of state-contingent securities unhinges domestic consumption patterns from domestic production, which then becomes free to specialize according to comparative advantage, for instance.[5] Thus, as figure 1 illustrates, both goods and assets trade may have direct as well as indirect effects on business cycle synchronization, with ambiguous overall effect. Classic Ricardian or Heckscher-Ohlin specialization may mitigate the direct effect of openness to goods trade, whereas financial integration may decrease (or increase) synchronization, but will also unambiguously induce specialization. The theoretical possibility for direct as well as indirect channels calls for a simultaneous-equation methodology, that is, a unified framework in which to study the magnitude of all these linkages. This is this paper's purpose.

The main results are as follows. (i) Specialization patterns have a sizable effect on business cycles. The bulk of this effect is independent of trade and financial policy, but directly reflects levels of GDP per capita. (ii) A variety of alternative measures of financial integration suggest that economic regions with strong financial links are significantly more synchronized, even though they are also more specialized. The positive direct effect of finance on synchronization dominates the negative indirect one working via higher specialization. (iii) The simultaneous-equation approach makes it possible to disentangle the importance of inter- and intra-industry trade, as illustrated by the upward dashed arrow in figure 1. The estimated effect of trade on cycle synchronization is closer to—and in some cases consistent with—that implied by existing models, once intra-industry trade is held constant.

Results implied by international and intranational data are strikingly similar. This is important for three reasons:

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[1] A nonexhaustive list includes Ambler, Cardia, and Zimmermann (2002), Canova and Dellas (1993), Baxter (1995), Kollman (2001), and Kose and Yi (2002). See Imbs (2001) for details.

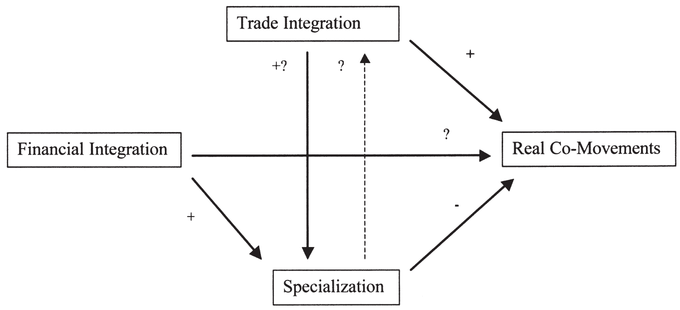
[2] For the first line of argument, see Heathcote and Perri (2002a, 2002b). For the second one, see Calvo and Mendoza (2000) or Mendoza (2001). These latter models were written with the purpose of explaining sudden reversal of capital flows to emerging markets, but there is no reason why their logic could not apply more generally. Both view financial integration as exogenous to cycle synchronization. According to portfolio theory, however, financial flows should be most prevalent between economic regions that are out of phase.

[3] For a recent theoretical development of this possibility, see Kraay and Ventura (2001).

[4] Most classical trade models make this prediction. For instance, falling transport costs in Dornbusch, Fischer, and Samuelson (1977) result in a narrowing nontraded sector, as it becomes cheaper to import goods rather than produce them domestically. Thus resources are freed up and used more intensely in fewer activities.

[5] For early models of this mechanism, see Helpman and Razin (1978), Grossman and Razin (1985), and Saint-Paul (1992).

FIGURE 1.—DIRECT AND INDIRECT CHANNELS: EQUATION-BY-EQUATION EVIDENCE IN THE LITERATURE



This figure summarizes the views in the literature on the interactions between goods trade, financial integration, specialization, and cycle synchronization. A + or − sign reflects a theoretical prediction that was confirmed empirically. A sign with a question mark means theory has not directly been tested empirically. A bare question mark means theoretical predictions are ambiguous.

(i) The fact that financial integration appears to result in correlated business cycles is not an artifact of an international convergence of policymaking, in particular monetary. Similar estimates obtain across U.S. states and across countries with substantially different monetary policies.
(ii) The importance of specialization patterns in affecting cycles is not due to the arbitrary choice of a time period or geographic coverage. In particular, the results cannot stem from the prevalence of one given type of shock in a given sample.[6]
(iii) Trade treatment is constitutionally homogenized across the states of the union. This legitimizes focusing on bilateral trade flows, since third-party treatment is the same for all pairs of states.[7]

The results in this paper suggest that theories of the international business cycle should build on the following ingredients: some sectoral heterogeneity (for example, in the responses of different sectors to a given macroeconomic shock), trade both within and between industries, and some herding in international capital flows (for example, through liquidity constraints or imperfect information).[8] The rest of the paper proceeds as follows. Section II provides a brief review of the relevant literature. Section III describes the data, main econometric issues, and general methodology. Section IV presents the main results, and section V concludes.

II. Literature

Most, but not all, of the linkages in this paper have been investigated empirically, but never simultaneously.[9] Most famously, the direct effect of trade on synchronization is documented by Frankel and Rose (1998), who estimate a strong and robust positive relationship between trade and cycle synchronization. They interpret their single-equation estimate as indicative that trade-induced specialization has but a small effect on business cycles, and is dominated by the direct positive link. Given the large evidence of the specialization effects of goods trade, it is of independent interest to quantify precisely the magnitude of this indirect effect of trade on business cycle correlations.[10] This is a first justification for the simultaneous estimation method implemented in this paper.[11]

The impact of financial integration on specialization is well documented too. For instance, Kalemli-Ozcan, Sorensen, and Yosha (2003) show there is a significantly positive relationship between specialization and risk-sharing. Thus, financial integration should affect (negatively) cycles synchronization, indirectly via its effect on specialization. The evidence on a *direct* link between finance and the extent of cofluctuations is, as suggested by theory, equivocal. Heathcote and Perri (2002b) argue the U.S. business cycle has become increasingly idiosyncratic over the past 30 years, and relate this to the increasing share of international assets held in the United States.[12] However, a considerable amount of empirical work lends support to the claim that capital flows are correlated internationally, and that financial integration tends to synchronize business cycles.[13] This is the second justification for the simultaneous approach in this paper, as financial integration could in theory affect synchronization both directly and indirectly. As illustrated in figure 1, the link between finance and cycle correlations is ambiguous for two reasons: first, the sign of the direct link is unclear in theory, and second, the indirect

[6] Furthermore, this paper uses altogether three different sources of sectoral data, measured at three different levels of aggregation (one-, two-, and three-digit levels). The specialization variable is always significant, regardless of the coarseness of the data. This makes it hard to ascribe the results to sampling.

[7] An extensive sensitivity analysis, including estimates corresponding to intranational U.S. state data, is available upon request, and posted on the author's Web site at faculty.london.edu/jimbs.

[8] These results are based on a measure of business cycle synchronization that is simultaneous. Thus, channels with a lag of more than a year (the lowest frequency of the data used) are not the focus here. This centers the analysis onto relatively fast transmission channels. This is also done for the sake of comparison with a large existing literature, indeed concerned with the determinants of the *contemporaneous* correlations between business cycles.

[9] Otto, Voss, and Willard (2001) estimate a reduced-form equation where GDP correlations are regressed on bilateral trade, financial openness, and an indicator of monetary policy. They also control for specialization. There are several differences between their approach and that of this paper: (i) they do not estimate a system, as they do not propose to identify specific channels; (ii) they do not allow for the endogeneity of specialization patterns; (iii) they do not allow for the possibly complex variance-covariance structure of the residuals, which is done here using GMM. Their results are on the whole consistent with those presented here.

[10] For instance, Harrigan (2001) and Harrigan and Zakrajsec (2000) show trade-induced specialization patterns to be significant, and consistent with theory.

[11] As will become clearer, the procedure also estimates the proportion of the overall effect of trade that is due to intra-industry trade. This is akin to Gruben, Koo, and Millis (2002) and to Shin and Wang (2003) for East Asian countries, although using different data and methodology.

[12] This, in turn, is endogenously caused by a stronger diversification motive, as shocks are argued to have become less correlated since the 1970s.

[13] See for instance Claessens, Dornbusch, and Park (2001), Calvo and Reinhart (1996), or Cashin, Kumar, and McDermott (1995). Admittedly, most of this evidence concerns pathological cases experienced by emerging economies, but there is no a priori reason to dismiss similar, if milder, arguments between developed economies.

specialization effect could either mitigate or reinforce the direct link.[14]

It is important to note that these channels correspond to the view that financial integration is exogenous to specialization patterns and business cycle synchronization. That is not necessarily so. Specialization in production could affect financial flows, if for instance specialization patterns were a low-frequency phenomenon, largely exogenous to policy changes, that would produce more or less of a need for financial integration. This would create a positive bias in the estimates of the effects of finance on specialization. Similarly, diversification motives imply that capital should flow between economies at different stages of their cycle, that is, economies that tend to be out of phase. This would create a negative bias in the estimates of the effects of finance on synchronization. This paper follows Kalemli-Ozcan et al. (2003) in using the instruments for financial development introduced by La Porta et al. (1998) to allow for both of these potential endogeneity biases.[15] The results are consistent with the presence of both biases, and instrumental variables estimations confirm results due to Kalemli-Ozcan et al. (2003) that financially integrated regions are more specialized, and less correlated as a result. Results also point to a significant *direct* and *positive* effect of finance on synchronization, which is new to the literature.

Finally, the direct effect of sectoral specialization on business cycle synchronization, although intuitive, is perhaps the least researched empirical question amongst those addressed in this paper. Otto et al. (2001), Kalemli-Ozcan, Sorensen, and Yosha (2001), and Imbs (2001) all find a significantly positive role for an index of similarity in production structures. Clark and van Wincoop (2001) use a similar index to account for higher business cycle correlations within than between countries.[16] But although they all point to a sizable direct effect of specialization on business cycles, none of these papers include the possibility that specialization could be an indirect manifestation of trade or financial integration, and amend the estimated effects of trade, finance, and specialization accordingly. This is the third justification for a simultaneous approach, which appears to be implicit in most of the existing empirical work.[17]

[14] This paper thus asks the question of how financial integration affects the real side of the economy. Other papers in the Symposium focus instead on how real variables affect financial returns. For instance, Brooks and Del Negro (2003) uncover an increasing role for the sectoral component of returns, perhaps a manifestation that countries themselves are becoming increasingly specialized.

[15] To be precise, La Porta et al. (1998) construct international data on a number of institutional determinants of financial development, in three broad categories: indices of shareholder rights, of creditor rights, and of enforcement laws.

[16] Their reasoning is based on the premise that regions within countries have more similar production structures than regions in different countries.

[17] In a similar exercise applied to financial markets, Chinn and Forbes (2003) assess the relative magnitudes of trade, banking, and FDI linkages in explaining international correlations of financial returns.

III. Methodology and Econometric Issues

This section introduces the system of equations estimated in the paper, and relates it to the relevant literature. It then briefly describes the variables involved, their measurement, and data sources, and then closes with an account of the specific heteroskedasticity problem in a cross section of bilateral correlations.

A. *The System*

This paper estimates the following system of equations simultaneously:

$$\rho_{i,j} = \alpha_0 + \alpha_1 T_{i,j} + \alpha_2 S_{i,j} + \alpha_3 F_{i,j} + \alpha_4 I_{1,i,j} + \varepsilon_{1,i,j}, \tag{1}$$

$$T_{i,j} = \beta_0 + \beta_1 S_{i,j} + \beta_2 I_{2,i,j} + \varepsilon_{2,i,j}, \tag{2}$$

$$S_{i,j} = \gamma_0 + \gamma_1 T_{i,j} + \gamma_2 F_{i,j} + \gamma_3 I_{3,i,j} + \varepsilon_{3,i,j}, \tag{3}$$

$$F_{i,j} = \delta_0 + \delta_1 I_{4,i,j}, \tag{4}$$

where i, j index country pairs, ρ is bilateral business cycle correlation, T is bilateral trade intensity, F is bilateral financial integration, and S is a specialization index capturing how different the sectorial allocations of resources are between countries i and j. Business cycle correlations, bilateral trade, financial integration, and specialization all are endogenous variables, and I_1, I_2, I_3, and I_4 contain the vectors of their exogenous determinants, respectively. Identification of the system requires differences between at least I_2 and I_3, as well as instruments for F. Fortunately, a substantial literature exists to provide guidance on these issues. I next turn to this question.

The dependent variable in equation (1) is prominent in the list of optimal currency area criteria.[18] It is therefore of interest in its own right, and indeed, its determinants have been the object of intense scrutiny. Frankel and Rose (1998) focus on α_1, reasoning that if currency unions affect trade and trade in turn boosts cycle correlations, then currency areas can endogenously become optimal.[19] Imbs (2001) focuses on α_2, arguing that measured bilateral trade may partly be a manifestation of differences in the degrees of specialization between the trading countries, which could affect ρ independently, as well as the estimates of α_1.[20] In a

[18] It is for instance one of the five tests, set by Gordon Brown, that the U.K. economy has to pass to enter the EMU.

[19] Alesina, Barro, and Tenreyro (2002) use a slightly different methodology to answer a similar question. They investigate the effect of currency unions on both trade and comovements. They instrument the advent of currency unions with gravity variables involving a third (anchor) country, rather than the bilateral characteristics used to explain bilateral trade intensity.

[20] Kalemli-Ozcan et al. (2001) estimate a variant of equation (1), but without a trade term. Then, Kalemli-Ozcan et al. (2003) estimate a variant of equation (3), and let specialization depend on financial integration. Implicitly, therefore, their two papers seek to document *one* of the channels in this paper, although not using simultaneous techniques. Their

single-equation framework, however, it is impossible to identify the reasons why α_2 is significant. Specialization can be the result of trade or financial integration, or can have its own dynamics, reflecting low-frequency changes in the economy.

Theoretical models are typically unable to replicate (single-equation) estimates of α_1, a point developed by Kose and Yi (2002). Baxter (1995) reviews the theories that imply a positive α_1, and Canova and Dellas (1993), Schmitt-Grohé (1998), Crosby (2002), and Kose and Yi (2002) have all used different methods to document the inability of existing models to reproduce the magnitude of standard estimates for α_1.[21] A plausible explanation is that we are not using the appropriate modeling strategy when attempting to reproduce the observed effects of trade. In particular, α_1 represents the effects of both inter- and intra-industry trade, two dimensions that the models typically do not share. The simultaneous approach makes it possible to decompose the two effects, as β_1 in equation (2) captures the extent to which bilateral trade can be accounted for by the similarities in the two countries' economic structures, that is, on the basis of intra-industry trade, denoted by an upward dashed arrow in figure 1. Thus, the total effect of trade in the simultaneous estimation equals $\alpha_1\beta_1 + \alpha_1\beta_2$, where the first term captures the importance of intra-industry trade. One-sector models should seek to reproduce $\alpha_1\beta_2$ only.[22]

This leaves open the question of what additional regressors to include in equation (1). Since this paper is concerned with the (direct and indirect) effects of financial integration, an important variable is one capturing the extent of impediments to capital flows between each pair of countries, as measured for instance by capital account restrictions or an estimate of international risk sharing.[23] Thus, the determinants of $\rho_{i,j}$ in equation (1) are trade intensity, specialization, and financial integration, which is already more than what is usually included in single-equation estimations in the literature, which are typically focused on one of these variables only.[24] An important omission pertains to convergence in policy. A companion sensitivity analysis available on the author's Web site shows that the results are not changed when controls for monetary policy are included, nor if the same estimations are implemented using data on U.S. states, even though they are subjected to a single monetary policy.[25,26]

The specification of equation (2) is more straightforward, although also a subject of debate. The empirical performance of so-called gravity variables in accounting for trade flows goes back at least to Tinbergen (1962), and has subsequently been used extensively.[27] The gravity variables in I_2 customarily include measures of both countries' GDP (or sometimes their populations), the geographic distance between their capitals, and binary variables capturing the presence of a common border and linguistic similarities between them. The list is usually argued to contain clearly exogenous variables with high predictive power for trade flows, thus supplying an exceptional instrument set.[28] This paper uses similar insights to isolate the exogenous effects of trade on cycle synchronization and specialization.[29]

The exogenous determinants of specialization, summarized in I_3, are less established empirically. Two sets of variables do however spring to mind. First, access to financial markets will influence specialization patterns and how similar they are between countries. Thus, the vector I_3

result that financially integrated regions specialize, and are less correlated as a result, obtains here as well.

[21] Canova and Dellas and Schmitt-Grohé, for instance, use structural VAR techniques. Crosby focuses on cycle synchronizations within the Asia-Pacific region. Kose and Yi simulate a three-country model and argue that standard models with technology shocks predict that $\alpha_1 < 0$, at least with complete markets. Kollman (2001) argues that nominal rigidities and demand shocks are crucial in reproducing international output correlations.

[22] A similar point is developed by Gruben et al. (2002), who include a measure of both inter- and intra-industry trade on the right-hand side of equation (1). The coefficients on the two components are found to be significantly different.

[23] All variables, and their instruments, are described in detail in the next section.

[24] For instance, Frankel and Rose (1998) only include trade, and Heathcote and Perri (2002b) focus on finance. Kalemli-Ozcan et al. (2001) focus on measures of specialization, while controlling for population, the GDP shares of agriculture and mining, GDP per capita, and human capital, but not for trade. In a companion paper available on the author's Web site, extensive sensitivity analysis shows that the inclusion of country size, GDP per capita, and human capital in equation (1) does not alter the results.

[25] The possibility that international economic fluctuations are caused by common shocks is also a prominent explanation of cycle correlations, and the object of a burgeoning literature. Leading candidates for global shocks are the sudden swings in the price of crude oil, witnessed throughout the 1970s and some of the 1980s. Loayza, Lopez, and Ubide (2001) perform a decomposition of output fluctuations in the developing world into global, country, and sectoral components, and find a dominant role for sectoral interdependences. In a paper at this symposium, Kose, Otrok, and Whiteman (2003) perform a similar decomposition using Bayesian techniques. The sample analyzed in this paper excludes the time periods commonly thought to correspond to global shocks, which suggests they are not particularly prevalent in the data used. Furthermore, once again, the results are almost identical across countries and across U.S. states.

[26] Stockman (1988) documents the prevalence of country-specific shocks in European countries. Although S could be an important determinant of ρ even in the absence of sectoral shocks, if S turns out to capture country-specific developments better than other variables, the international results can be interpreted differently. There is no particular reason to expect S to capture country-specific shocks. Furthermore, the analog in the U.S. state data would entail the prevalence of state-specific shocks, a highly improbable assumption in view of the constitutional restrictions on state-level fiscal policy and the absence of independent state central banks.

[27] See, among many others, Frankel and Rose (1998, 2002), Frankel and Romer (1999), or Rose (2000).

[28] Used for instance to identify causality between trade and growth by Frankel and Romer (1999), to control for other determinants of trade by Rose (2000), and to do both by Frankel and Rose (2002).

[29] The list could be extended to include endowment variables, as implied by Ricardian trade theory. This would however make it difficult to compare the results with existing single-equation estimates, such as those in Frankel and Rose (1998). The only role for the gravity variables in the present context is to isolate the exogenous component of T in the system, another reason why no proxies for factor endowment are included within I_2. Thus the methodology does not fall victim to the criticism in Rodrik, Subramanian, and Trebbi (2002) that gravity variables merely capture good institutions, in turn conducive to high growth, nor to Persson's (2001) critique that geography is inherently more conducive to trade within currency unions. The influence of currency unions on trade and that of trade on growth are not central to this paper.

includes the measures of financial integration already used in I_1. Second, Imbs and Wacziarg (2003) show economies go through two stages of specialization as income per capita grows: they initially diversify but respecialize once a (relatively high) level of income per capita is reached. This empirical fact suggests two additional components for I_3: GDP per capita levels in both economies, but also, because of the nonmonotonicity, the gap between them.[30]

In summary, since the intersection between I_2 and I_3 is empty, the system can be identified through a choice of instruments that is largely warranted by an existing literature. The main contribution of the present exercise is simultaneity. The main assumption is the exogeneity of GDP per capita and relative GDP per capita to S.[31] I now turn to a detailed description of the data and measurement of the variables included in the vectors I_i.

B. *Data and Measurement*

Bilateral correlations in business cycles are computed on the basis of the cyclical component of quarterly GDP, isolated using the bandpass filter introduced by Baxter and King (1999).[32] The quarterly data cover the 1980s and 1990s in 24 countries, and come from the International Financial Statistics issued by the IMF.[33] This gives rise to a cross section of 276 bilateral correlations. A concern with this cross section is that it builds on 20 years of data, over which the United States, for instance, has experienced two, or at most three, business cycles. Can one safely draw inferences on the medium- to long-run characteristics of the cross-section of business cycle synchronization on such a reduced sample? This is of course a concern that pervades any such exercise, and in particular *all* the papers mentioned earlier.[34] The concern is assuaged using two alternative data sources. First, bilateral correlations in business cycles are computed on the basis of yearly data taken from version 6.1 of the Penn-World Tables, which covers the period 1960–2000. Second, (annual) series on gross state product from 1977 to 2001 are taken from the Bureau of Economic Analysis. Results pertaining to these alternative data are reported in the sensitivity analysis. Similar conclusions obtain in all cases.[35]

Bilateral trade intensity is computed in two ways. The first one is standard, used by Frankel and Rose (1998) among others, and writes

$$T^1_{i,j} = \frac{1}{T}\sum_t \frac{X_{i,j,t} + M_{i,j,t}}{Y_{i,t} + Y_{j,t}},$$

where $X_{i,j,t}$ denotes total merchandise exports from country i to j in year t, $M_{i,j,t}$ denotes imports to i from j, and Y_i denotes nominal GDP in country i. Bilateral trade data are from the IMF's Direction of Trade Statistics. I use this standard measure for the benchmark case. Clark and van Wincoop (2001) use an alternative measure of trade intensity, independent of size, based on the model in Deardorff (1998), which can be constructed as

$$T^2_{i,j} = \frac{1}{2}\frac{1}{T}\sum_t \frac{(X_{i,j,t} + M_{i,j,t})Y^W_t}{Y_{i,t}Y_{j,t}},$$

where Y^W_t is world GDP. T^2 differs from T^1 in that it depends only on trade barriers, and not on country size. In particular, Deardorff shows that $T^2 = 1$ if preferences are homothetic and there are no trade barriers. Estimations based on T^2 are reported in the sensitivity analysis.

There are no standard measures of similarity in industry specialization. Krugman (1991) and Clark and van Wincoop (2001) favor a variable akin to a Herfindahl index of concentration, whereas Imbs (2001) uses a correlation coefficient between sectoral shares in aggregate output or employment. Here, sectoral real value added is used to compute

$$S_{i,j} = \frac{1}{T}\sum_t \sum_n^N |s_{n,i} - s_{n,j}|,$$

where $s_{n,i}$ denotes the GDP share of industry n in country i. In words, $S_{i,j}$ is the time average of the discrepancies in the economic structures of countries i and j.[36] Thus, S reaches its maximal value for two countries with no sector in common: one should therefore expect $\alpha_2 < 0$. The sectoral shares s are computed using two alternative data sources: two-digit manufacturing value-added data issued by the UNIDO, and, for robustness, data from the United Nations Statistical Yearbook (UNYB), which provides sectoral value

[30] Kim (1995) investigates the dynamics of geographic specialization in U.S. states over the long run. His findings confirm the importance of development variables in affecting patterns of specialization.

[31] The determinants of S are poorly known beyond the role of trade and finance. I interpret the evidence in Imbs and Wacziarg (2003) as suggestive that specialization is a manifestation of growth, rather than the opposite.

[32] The parameters are set according to Baxter and King's recommendations. In particular, the filter is set to preserve the components of the data with period between 6 and 32 quarters for quarterly data, and between 2 and 8 years for annual data. In the sensitivity analysis, the Hodrick-Prescott filter is also applied to the data, with $\lambda = 1600$ or 100 in quarterly or annual data, respectively. Initial and final observations are discarded, following Baxter and King's recommendations.

[33] This is the maximal uninterrupted coverage afforded by the November 2002 IFS CD-ROM where sectoral data are available. The countries included are Australia, Austria, Belgium, Canada, Chile, Finland, France, Germany, Israel, Italy, Japan, Korea, Mexico, the Netherlands, Norway, Peru, the Philippines, Portugal, South Africa, Spain, Sweden, Switzerland, the United Kingdom, and the United States.

[34] And resorting to longer (monthly) time series on industrial production will not solve the problem either, given the shrinking share of economic activity this represents.

[35] Alesina et al. (2002) use an alternative measure of cycle synchronization. They fit an AR(2) to relative per capita GDP. A measure of (lack of) comovements is given by the root-mean-square error of the residual. For comparability with most existing single-equation estimates, this paper focuses on a measure of comovements based on correlation coefficients.

[36] Both the trade and specialization measures are based on time averages. Results do not change if the initial value is used instead.

added at the one-digit level for all sectors in the economy, with (incomplete) coverage from 1960 to 1998. The UNIDO data cover only manufacturing sectors, and thus a shrinking share of most economies in the sample, but country coverage is more patchy in the UNYB data.[37] For the intranational estimation, whose results are relegated in the sensitivity analysis, I use real sectorial state value-added series issued by the Bureau of Economic Analysis. These cover all economic activities, at the three-digit aggregation level.[38]

Bilateral financial integration is notoriously difficult to measure. I therefore maximize the number of alternative data sources and measurement strategies. The proxies can be separated into two distinct groups: indices capturing *restrictions* on capital flows, and variables meant to reflect *effective* financial flows. The two measures are fundamentally different in that the absence of any impediments to capital flows does not necessarily result in deep financial integration. Included in the first group are the IMF's binary index of capital account restrictions, and the extension thereof reported in the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions* (*AREAER*).[39] These indices are summed pairwise, and thus report the average number of countries with restrictions on financial flows, over all country pairs.[40,41]

Measuring *effective* bilateral capital flows is an even taller order, but an important one given the limitations inherent in indices of capital account restrictions. Here two proxies are proposed. Using U.S. state data, Asdrubali, Sorensen, and Yosha (1996) and Kalemli-Ozcan et al. (2003) compute a state-specific index of risk sharing by estimating

$$\ln gsp_t - \ln dy_t = \alpha + \beta \ln gsp_t + \varepsilon_t, \tag{5}$$

where gsp is the gross state product per capita and dy is state disposable income per capita. Both papers interpret β as an index of risk sharing. If interstate risk sharing (or income insurance) is perfect, then $\beta = 1$ as disposable income is unrelated to GSP per capita, and equation (5) is simply a regression of gsp on itself. Conversely, if there is no interstate risk sharing, then $\beta = 0$, because the dependent variable becomes essentially noise. A measure of cross-region financial integration is then given by pairwise sums of the region-specific estimates of β. I use data on gross domestic product and private consumption to extend these estimates to the international data.[42]

An alternative uses the recent data set constructed by Lane and Milesi-Ferretti (2001), reporting cumulated external positions for a large sample of developing and industrial economies. Pairs of countries with intense capital flows should display different (or even opposite) external positions. Two countries with massively positive (negative) net foreign assets holdings will both tend to be issuers (recipients) of capital flows, and should experience less bilateral flow than two countries where one is structurally in surplus and the other in deficit. Thus, the Lane–Milesi-Ferretti data can be used to construct $LMF = |(NFA/\text{GDP})_i - (NFA/\text{GDP})_j|$, where NFA denotes the net foreign asset position in country i. LMF will take high values for pairs of countries with diverging external positions, more likely to lend and borrow from each other than pairs of countries with similar external positions. The NFA position is computed using both accumulated current accounts and the sum of net positions in foreign direct investment, equities, and debt.[43]

The interpretation of α_3 and γ_2 changes with the measure used for F: with indices of restrictions to the capital account, a prospecialization effect of finance implies $\gamma_2 < 0$. On the other hand, with risk-sharing indices or LMF, it implies $\gamma_2 > 0$: more risk sharing and different net foreign positions are associated with higher specialization and higher S. The same is true for the interpretation of α_3.

The gravity variables in I_2 are standard, and include the following: a measure of the (log mile) distance between the countries' capitals, the (log) products of each country's GDP, and binary variables indicating the presence of a common border and whether they share the same language. The vector I_3 is different from I_2, and includes the proposed measure of financial integration, already in I_1; the (log) product of each country's GDP per capita; and the (log) GDP disparity, defined as $\max[(Y_i/Y_j), (Y_j/Y_i)]$.[44] Finally, the vector I_4 contains instruments taken from La Porta et al. (1998), including measures of shareholder rights (with variables capturing the percentage of capital necessary to call an extraordinary shareholders' meeting, whether one share carries one vote, whether the distribution of dividends is mandatory, and whether proxy vote by mail is allowed), an aggregate index of creditor rights, and an assessment of accounting standards and the rule of law.

[37] This is the main data limitation for the present exercise. The intersection between quarterly GDP IFS data and sectoral UNIDO data is what is used in this paper. UNYB data is used in the sensitivity analysis, available in a companion paper. One-digit sectors are: 1. Agriculture, Hunting, Forestry, Fishing; 2. Mining & Quarrying; 3. Manufacturing; 4. Electricity, Gas & Water; 5. Construction; 6. Wholesale, Retail Trade, Restaurants & Hotels; 7. Transport & Communications; 8. Finance, Insurance, Real Estate & Business Services; 9. Community, Social & Personal Services. There are 28 two-digit manufacturing sectors.

[38] There are 61 sectors in each state.

[39] The latter includes four binary variables reflective of: (i) multiple exchange rates, (ii) current account restrictions, (iii) capital account restrictions, and (iv) surrender of export proceeds.

[40] The composite index from AREAER is averaged each year, and thus can take values 0.25, 0.5, 0.75, or 1. It is then summed pairwise. I use initial values of both indices, in 1970, but taking an average over the whole period makes no difference.

[41] A few alternative measures of capital account restrictions are available, some within this very symposium, but mostly for developing economies. For instance, the data sets of Kaminsky and Schmukler (2002) and of Edison and Warnock (2001) cover almost exclusively developing economies.

[42] A time trend is included in equation (5) to isolate the idiosyncratic component. This follows Kalemli-Ozcan et al. (2003).

[43] LMF is computed at the beginning of the sample, in 1980. Averaging over the whole sample does not change any results.

[44] The gravity variables are taken from Andrew Rose's Web site at http://faculty.haas.berkely.edu/arose.

TABLE 1.—SUMMARY STATISTICS

Statistic	Mean	Min	Max	Std. Dev.	N.Obs
ρ_Q	0.193	−0.586	0.887	0.310	276
ρ_Y	0.201	−0.357	0.790	0.245	276
T^1	0.004	7.22×10^{-5}	0.069	0.007	276
T^2	0.285	0.013	5.347	0.489	276
S_{YB}	0.380	0.143	0.747	0.125	171
S_{Mfg}	0.540	0.150	0.976	0.159	276

Table 1 reports summary statistics for the three endogenous variables in the system. Table 2 reports the corresponding unconditional correlations. As usual, and particularly in the present context of simultaneity, unconditional correlations are informative only superficially. Several points are however worth noting. First, the cross sections of cycle synchronizations appear to be very similar, irrespective whether the cycles are computed on the basis of quarterly or annual data. ρ_Q and ρ_Y have similar moments and extreme values, and are highly correlated. Second, average specialization in manufactures is higher than across all economic activities, and the correlation between the two cross sections is only 0.47, which warrants some sensitivity analysis. In what follows, manufacturing data are privileged in computing indices of specialization because of their wider coverage, even though it represents a shrinking portion of the economy in most of the sample. The sensitivity analysis shows however that the results obtain as well in the UNYB data. Third, both measures of trade intensity are very similar, with a correlation of 0.85. They both correlate positively with ρ, although T^1 does so more significantly. Fourth, specialization is correlated negatively with cycle synchronization, no matter the combination of measures used. The evidence for a trade-induced specialization is weak a priori, because the correlation between specialization and trade measures is only positive when using T^2, and then very weakly so. Unsurprisingly, this calls for appropriate conditioning, as well as a simultaneous approach.

Table 3 reports a few extreme values for the four main variables. It is remarkable that the only country pair displaying both high business cycle correlation and high trade linkages is the United States with Canada, precisely where Schmitt-Grohé (1998) showed the theoretical inability of trade per se to account for cycle synchronization. Thus, even between the two countries where the case for trade would be the strongest, the variable alone appears insufficient to account for observed synchronization. The data suggest trade may be important, but probably insufficient to explain the whole cross section of cycle correlations.

TABLE 2.—UNCONDITIONAL CORRELATIONS

	ρ_Q	ρ_Y	T^1	T^2	S_{YB}	S_{Mfg}
ρ_Q	1.000					
ρ_Y	0.743	1.000				
T^1	0.324	0.437	1.000			
T^2	0.163	0.262	0.848	1.000		
S_{YB}	−0.414	−0.469	−0.187	0.032	1.000	
S_{Mfg}	−0.250	−0.410	−0.228	−0.027	0.476	1.000

NOTES: ρ_Q (ρ_Y) denotes bilateral correlations in GDP on the basis of bandpass-filtered quarterly (yearly) data. S_{YB} (S_{Mfg}) denotes the measure of sectoral similarities implied by the UNYB (UNIDO) data. T^1 and T^2 are the alternative measures of bilateral trade described in the text.

TABLE 3.—SELECTED MINIMA AND MAXIMA

Correlation		Trade	
−0.586	Norway–Portugal	7.22×10^{-5}	Australia–Peru
−0.522	Philippines–S. Africa	7.40×10^{-5}	Norway–Philippines
−0.471	Israel–U.S.	8.43×10^{-5}	Norway–Peru
0.839	Switzerland–Nlds	0.029	France–Germany
0.862	Belgium–Italy	0.031	Germany–Nlds
0.867	U.S.–Canada	0.074	Belgium–Nlds
Specialization		**Financial Integration**	
0.143	Australia–Canada	0	Mexico–U.S.
0.155	Netherlands–Canada	0	Canada–U.S.
0.157	U.K.–France	0	Mexico–Canada
0.653	Israel–Mexico	2	Italy–S. Africa
0.692	Norway–Philippines	2	U.K.–S. Africa
0.747	Israel–Philippines	2	Netherlands–Japan

NOTES: Correlations are based on bandpass-filtered quarterly data. Trade is based on T^1, specialization is based on UNIDO data, and financial integration is given by a binary variable capturing capital account restrictions across countries.

Turning to measures of specialization, the three most similar country pairs (on the basis of manufacturing data) are Australia–Canada, the U.K.–France, and the Netherlands–Canada. The importance of sectoral specialization is therefore of special interest to inform the debate on Sterling's entry into the EMU. Finally, financial integration as measured by capital account restrictions is highest between the North American countries and lowest when such countries as South Africa are included.[45]

C. Three-Stage Least Squares and Heteroskedasticity

The estimation strategy must combine the features of simultaneous equations procedures, and allow for the possible endogeneity of some dependent variables, since for instance specialization can be endogenous to trade, and vice versa. Three-stage least squares (3SLS) does exactly that, using simultaneous estimation and instrumentation to isolate the different components of the endogenous variables. The estimator combines insights from instrumental variable and generalized least squares methods, achieving consistency through instrumentation, and efficiency through appropriate weighting in the variance-covariance matrix. The procedure consists of the following two steps: (i) estimate the system equation by equation using two-stage least squares, and retrieve the covariance matrix of the equations disturbances; then (ii) perform a type of generalized least squares estimation on the stacked system, using the covariance matrix from the first step.[46]

[45] The pair Belgium–the Netherlands has very high values for both T^1 and T^2, and could thus be considered an outlier. The subsequent results are however invariant to omitting this country pair. The estimations that follow do include the "outlier," and are run in logarithms. Estimations in levels yield virtually identical results, with lower fits.

[46] See for instance Tavares and Wacziarg (2001) for an application of three-stage least squares to the effects of democracy on growth.

TABLE 4.—SIMULTANEITY

	OLS		3SLS	
	(i)	(ii)	(iii)	(iv)
(1) Correlation ρ				
T	0.077 (6.00)	0.064 (4.25)	0.074 (3.32)	0.079 (3.63)
S		−0.123 (1.92)	−0.298 (2.67)	−0.304 (2.43)
F		0.019 (0.57)	0.011 (0.35)	−0.112 (2.08)
R^2	0.116	0.124	0.083	0.140
(2) Trade T				
Distance	−0.951 (18.69)	−0.970 (17.43)	−0.814 (14.49)	−0.685 (11.79)
Border	0.007 (0.03)	0.119 (0.48)	0.152 (0.69)	0.028 (0.14)
Language	0.300 (2.07)	0.347 (2.02)	0.401 (2.58)	0.206 (1.40)
GDP Product	0.225 (7.37)	0.231 (7.26)	0.216 (5.66)	0.141 (2.61)
S	−0.424 (2.49)	−0.401 (2.22)	−1.470 (4.54)	−2.682 (5.18)
R^2	0.708	0.697	0.640	0.468
(3) Specialization S				
GDP per capita Product	−0.139 (6.02)	−0.185 (7.74)	−0.144 (6.11)	−0.079 (3.34)
GDP Gap	0.613 (2.21)	0.783 (2.81)	1.098 (4.26)	0.814 (2.93)
T	−0.051 (3.64)	−0.046 (3.37)	−0.079 (4.68)	−0.094 (5.64)
F		−0.101 (3.39)	−0.074 (2.62)	−0.064 (1.43)
R^2	0.274	0.347	0.322	0.202

NOTES: T is the logarithm of T_1; S is the (logarithm) measure of sectoral similarities based on UNIDO data, covering manufacturing activities. Both variables are averaged over time. F is an index of capital account restrictions in 1970. Intercepts are not reported. t-statistics between parentheses. (i) and (ii) are single-equation OLS estimates. (iii) and (iv) both perform 3SLS, but the latter instruments F with the institutional variables in La Porta et al. (1998). In particular, the instruments reflect shareholder rights (with variables capturing whether one share carries one vote, whether the distribution of dividends is mandatory, whether proxy vote by mail is allowed, and the percentage of capital necessary to call an extraordinary shareholders' meeting), creditor rights, and an assessment of accounting standards and the rule of law.

In addition, the correlations in business cycles, ρ, are measured with error, in a way that is likely to create a specific type of heteroskedasticity. In particular, following Clark and van Wincoop (2001), let $\hat{\rho} = \rho + \upsilon$ denote the estimated correlation coefficients, with υ the sampling error. It is then possible that υ is correlated across observations in $\hat{\rho}$, since many correlation coefficients involve the GDP series for the same country or state. This will create a kind of heteroskedasticity in the residuals of equation (1) that standard White corrections cannot take into account.

Ignoring this heteroskedasticity is likely to result in understated standard errors for the estimates in equation (1) and, since the estimation is simultaneous, in the rest of the system as well. It is however possible to take account of this possibility, at the cost of fairly mild assumptions. If the true vector of bilateral correlations, ρ, is assumed to be deterministic, equation (1) becomes

$$\rho_{i,j} = \alpha_0 + \alpha_1 T_{i,j} + \alpha_2 S_{i,j} + \alpha_3 I_{i,j,1} + \varepsilon_{i,j,1} + \upsilon. \tag{6}$$

The variance-covariance matrix of this equation involves $\hat{\Sigma}_\upsilon = E(\upsilon\upsilon)$, which requires using a GMM estimator. GMM results are presented in the sensitivity analysis, without any notable changes in the results.

IV. Trade, Finance, Specialization, and Synchronization

This section reports the paper's main results. Least squares results are compared with existing (single-equation) evidence, and 3SLS estimates are presented to evaluate the effects of simultaneity and endogeneity. Some robustness analysis is presented, evaluating sensitivity to alternative measures of financial integration. A final subsection discusses the relative magnitudes of the direct and indirect channels between trade, finance, specialization, and synchronization.

A. *Benchmark Results*

The first two specifications in table 4 report equation-by-equation estimates of the system, and confirm established single-equation results. The large and significant effect of trade in accounting for ρ is evident in column (i). The point estimate means that doubling trade results in a correlation higher by 0.048, which is close to the estimates in Frankel and Rose (1998), Clark and van Wincoop (2001), and Kose and Yi (2002).

The second specification adds the two variables specific to the simultaneous approach, namely specialization S and financial integration F. Three results are of interest:

(a) Financial integration has the predicted specialization effect, documented in Kalemli-Ozcan et al. (2003). Capital account restrictions (a high F) are associated with low S, that is, financially integrated economies tend to specialize in different sectoral patterns. However, the estimate of γ_2 potentially suffers from an endogeneity bias away from zero, as specialized countries could have an acute need for income insurance.

(b) Finance has no direct effect on synchronization, but estimates of α_3 potentially suffer from an attenuating bias, as capital should tend to flow between economies that are out of phase.

(c) Country pairs with low S have significantly higher ρ: similarities in economic structure result in correlated business cycles, and the R^2 suggest this is a quantitatively important variable.

Estimates for equation (2) in the first two columns are unsurprising. The gravity variables all have the expected signs, as is now completely standard in any empirical work concerned with explaining the geography of trade. The same is true of equation (3), with estimates of the γ's in line with theory. In particular, pairs of rich countries tend to have significantly lower values of S, that is, their economic structures are more similar, as would happen if growth were accompanied by diversification. On the other hand, pairs of countries at different stages of development, as measured by the gap between their GDPs, have significantly higher S, that is, tend to display different economic structures. Finally, the effect of trade on S is significant, but with the "wrong" sign, that is, more bilateral trade results in lower S, or in more similar economies. However, this result turns out to disappear when trade is measured using T^2.

Column (iii) implements 3SLS on the system. This tends to magnify the estimate of α_1 relative to column (ii): as in Frankel and Romer (1999), instrumenting trade with gravity variables results in a higher point estimates, as it controls for an attenuating endogeneity bias, because nonsynchronized economies tend to trade more. 3SLS also magnifies the estimate of β_1 relative to OLS. This makes sense, in that the endogeneity of S to T would if anything tend to bias β_1 upward, as trading partners specialize and thus have high S. A negative β_1 can be interpreted as meaning that countries with similar economic structures trade more, a quantification of the extent of intra-industry trade. Estimates of β_1 are therefore key to disentangling the effects of inter- versus intra-industry trade in simultaneous estimations. Finally, 3SLS magnifies the estimate of α_2, now that specialization is allowed to depend on both trade and financial integration.

The last specification in table 4 instruments F using the institutional variables introduced by La Porta et al. (1998).[47] Consistent with the theoretical reasons for the endogeneity of F, financial integration becomes a significantly positive determinant of ρ, when the tendency for capital to flow between nonsynchronized economies is accounted for. Furthermore, consistent with the conjecture that finance is acutely needed between specialized economies, estimates of γ_2 are weakened somewhat once F is instrumented.[48] Thus, this paper confirms the prospecialization effects of financial integration, and thus a negative *indirect* impact of finance on ρ. But the data also point to a *direct* effect of finance on ρ, found to be significantly positive. The next section verifies the robustness of these results to alternative measures of financial integration.

B. *Measures of Effective Financial Integration*

The previous section built on a particular measure of financial integration based on capital account restrictions, a mere quantification of potential capital flows between economies. Table 5 reports 3SLS estimates for the system (1)–(4), where the composite index of capital account restrictions is used in column (i), the index of risk sharing introduced by Asdrubali et al. (1996) is used in column (ii), and two measures of LMF are used in columns (iii) and (iv). In all cases, F is instrumented using the same set of institutional variables taken from La Porta et al. (1998). The positive direct effect of F on ρ, offset by a negative indirect one working via S, is confirmed in almost all cases. The effect of finance on cycles synchronization remains twofold in most instances.

Specification (i) uses the composite index of capital account restrictions issued by AREAER, with an interpretation of the signs of α_3 and γ_2 similar to table 4. Its estimates are virtually identical. When financial integration is measured by an index of (effective) risk sharing, the results are weaker. There are no significant direct effects of finance, and γ_2 has the "wrong" sign, suggesting that countries engaging in risk sharing (high F) tend to resemble each other. These weaker results are however the exception rather than the rule, as the rest of table 5, and the intranational results on U.S. states, both confirm the benchmark results. Economies with differing net foreign positions display significantly more correlated cycles, even though they also tend to specialize, and are less correlated as a result.[49]

Overall, the picture painted in table 5 is that of a substantial role for financial integration in affecting business cycle synchronization, whether it be measured using indices of capital account restrictions or net foreign positions. Consistent with theory and recent empirical results, finance affects specialization patterns, and thus indirectly business cycles. More intriguing from a theoretical standpoint is the result that financial integration directly increases business cycle correlations. The next subsection presents actual quantifications of these direct and indirect channels, pertaining to finance, but also to trade integration.

C. *Channels*

Panel A of table 6 relates the direct and indirect channels illustrated in figure 1 to the parameters obtained from the

[47] In particular, the instruments reflect: shareholder rights (one-share one-vote, mandatory dividend distribution, proxy by mail allowed, percentage of capital necessary to call an extraordinary shareholders' meeting), an aggregate index of creditor rights, and indices of accounting standards and the rule of law. Several permutations of instruments were experimented with, without sizable differences in the results. The instrument were chosen to maximize the fit of first-stage estimations, whose R^2's never fall below 0.33, no matter what the measure of financial integration used.

[48] But remain significant for alternative measures of F, as the next section documents.

[49] F is always instrumented in table 5, so that it is possible to draw causal inferences from the estimates.

TABLE 5.—DIFFERENT MEASURES OF FINANCIAL INTEGRATION

	(i) Composite	(ii) Risk	(iii) CA	(iv) CAPA
		(1) Correlation ρ		
T	0.067 (2.92)	0.079 (3.40)	0.116 (5.67)	0.057 (2.16)
S	−0.257 (2.23)	−0.352 (2.41)	−0.275 (2.04)	−0.429 (3.21)
F	−0.212 (3.70)	−0.149 (1.70)	1.177 (3.95)	0.135 (2.28)
R^2	0.186	0.121	0.151	0.106
		(2) Trade T		
Distance	−0.796 (14.90)	−0.714 (12.84)	−0.769 (14.44)	−0.839 (15.91)
Border	0.048 (0.24)	0.034 (0.19)	0.101 (0.51)	0.035 (0.17)
Language	0.308 (2.09)	0.222 (1.61)	0.312 (2.13)	0.316 (2.12)
GDP product	0.214 (4.94)	0.190 (4.17)	0.196 (4.93)	0.226 (5.26)
S	−1.358 (3.38)	−2.153 (5.10)	−1.623 (4.58)	−0.958 (2.41)
R^2	0.648	0.547	0.622	0.679
		(3) Specialization S		
GDP/capita product	−0.148 (5.43)	−0.059 (2.89)	−0.075 (3.31)	−0.122 (4.99)
GDP gap	1.291 (4.55)	0.912 (3.76)	1.108 (4.30)	0.983 (3.73)
T	−0.093 (5.61)	−0.099 (6.26)	−0.066 (3.97)	−0.100 (5.72)
F	−0.235 (4.32)	−0.251 (4.33)	1.153 (5.55)	0.196 (4.63)
R^2	0.272	0.252	0.322	0.285

NOTES: T is the logarithm of T^1; S is the (logarithm) measure of sectoral similarities based on UNIDO data, covering manufacturing activities. Both variables are averaged over time. The table presents results for different measures of F: (i) uses the composite index for capital account restrictions from AREAER, (ii) uses the index of risk sharing introduced in Asdrubali et al. (1996), (iii) uses the bilateral discrepancy in cumulated current accounts, and (iv) uses the equivalent measure on the capital account side. Intercepts are not reported. t-statistics between parentheses. All specifications perform 3SLS, with F instrumented using the institutional variables in La Porta et al. (1998). In particular, the instruments reflect shareholder rights (with variables capturing whether one share carries one vote, whether the distribution of dividends is mandatory, whether proxy vote by mail is allowed, and the percentage of capital necessary to call an extraordinary shareholders' meeting), creditor rights, and an assessment of accounting standards and the rule of law.

TABLE 6.—CHANNELS TO BUSINESS CYCLE SYNCHRONIZATION

	Direct	Indirect
	A. Formulas	
(1) Trade		
Intra-industry trade	$\alpha_1\beta_1$	
Geographic trade	$\alpha_1\beta_2$	
Via specialization		$\alpha_2\gamma_1$
(2) Specialization		
Trade-induced	$\alpha_2\gamma_1$	
Finance-induced	$\alpha_2\gamma_2$	
Stages of diversification	$\alpha_2\gamma_3$	
Via trade		$\alpha_1\beta_1$
(3) Finance		
Financial integration	α_3	
Via specialization		$\alpha_2\gamma_2$
	B. Estimates	
(1) Trade		
Intra-industry trade	−0.188 (0.000)	
Geographic trade	0.057 (0.044)	
Via specialization		0.018 (0.069)
(2) Specialization		
Trade-induced	0.018 (0.069)	
Finance-induced (CAPA)	0.317 (0.057)	
Stages of diversification	−0.627 (0.000)	
Via trade		−0.188 (0.000)
(3) Finance		
Current account	1.177 (0.000)	
Via specialization		0.317 (0.057)

NOTES: The values in panel B are computed on the basis of the estimates in specification (iii) of table 5. In particular, financial integration is measured using the bilateral discrepancy between cumulated current accounts. Two stages are necessary to estimate the geographic component of bilateral trade: First, estimate a gravity model for T^1. Second, use the fitted value of bilateral trade in the ρ equation of a system involving equations (1) and (3), to control for S's endogeneity. Similarly, stages of diversification are isolated in a regression of S on GDP product and GDP gap, and a fitted value for S is included in the ρ equation of a system involving equations (1) and (2), to control for T's endogeneity. P-values for the joint significance of each product are reported between parentheses.

simultaneous estimation. The coefficients α in equation (1) can be decomposed into contributions from trade, specialization, and financial integration. The direct effects of trade, captured in α_1, are either a reflection of exchanges within industry ($\alpha_1\beta_1$) or standard Ricardian trade ($\alpha_1\beta_2$). Separate estimates of β_1 and β_2 make it possible to identify the proportion of α_1, or the single-equation estimate of the effect of trade on cycle synchronization, that comes from intra-industry trade. Indirect effects of trade, in turn, come from the possibility that economies open to goods specialize, and have a higher value of S as a result ($\alpha_2\gamma_1$).

The direct effects of specialization are captured in α_2, and can originate in trade ($\alpha_2\gamma_1$), in financial integration ($\alpha_2\gamma_2$), or in the exogenous stages of development reached by both economies ($\alpha_2\gamma_3$). Specialization will have indirect effects too, working via trade, or in particular via intra-industry trade ($\alpha_1\beta_1$). Finally, the direct effects of financial integration are captured by α_3, and its indirect effects work via specialization ($\alpha_2\gamma_2$).

Panel B of table 6 reports the values for these channels as implied by 3SLS estimates in table 5.[50] A number of results are of interest. First, a large proportion of single-equation estimates of α_1 correspond to intra-industry trade.[51] Estimates in Table 6B suggest about three-quarters of α_1 can be ascribed to intra-industry trade, and only the remaining quarter to interindustry trade. In other words, models with

[50] Estimates from specification (iii) were used. The main conclusions remain if alternative measures of financial integration are used instead. Estimates for β_2 and γ_3 were obtained from the 3SLS fitted values for T on gravity variables and S on GDP variables, respectively.

[51] These results are in line with Fidrmuc (2002), who documents an important channel going from intra-industry trade to output correlations.

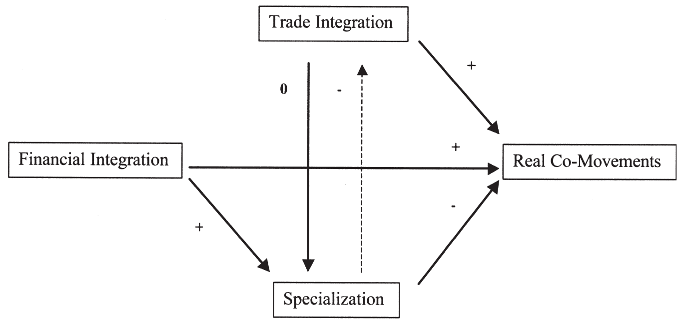
interindustry trade only should seek to reproduce the much smaller value of $\alpha_1 = 0.029$.

One word of comment is in order here. A number of authors have puzzled over the inability of standard one-sector models to reproduce the large effect trade has on cycle correlations. Recently, Kose and Yi (2002) have calibrated and simulated a three-country business cycle model with transport costs and technology shocks. Depending on the parametrization, their model yields simulated values for α_1 ranging from 0.0007 to 0.036.[52] On the basis of estimates in table 6, the share of α_1 originating from interindustry trade appears to be within this range, although still at its upper end. Thus, once focused on the link between interindustry trade and business cycle correlations, the data are not inconsistent with elasticities predicted by one-sector models.[53]

Most of the direct effect of S on ρ works through the determinants of specialization based on GDP per capita levels, labeled "Stages of diversification" in Table 6B, rather than policy changes. Financial integration is estimated to result in higher S, an effect that is quantitatively important although it is only significant at the 5.7% confidence level. But the bulk of specialization patterns originates in stages of diversification, from both an economic and a statistical standpoint. In other words, S is largely beyond the reach of short-term policymaking. Inasmuch as the international correlation of business cycles is an important constraint on policy, this puts into perspective the significantly positive estimates for the α's arising from single-equation estimations. Though it may be possible to manipulate T through trade policy, there is no immediate equivalent for S.

Finally, most of the effect of financial integration is direct, as countries with different net foreign positions tend to be more synchronized, even though they are also more specialized, and less correlated as a result.[54]

FIGURE 2.—DIRECT AND INDIRECT CHANNELS: SIMULTANEOUS EVIDENCE



This figure summarizes the paper's estimation results on the interactions between goods trade, financial integration, specialization, and cycle synchronization. Zero means the coefficient is estimated to be insignificant.

V. Conclusion

This paper estimates a system of simultaneous equations to disentangle the complex interactions between trade, finance, sectoral specialization, and business cycle synchronization. A large theoretical and empirical literature is referred to in choosing the sets of instruments necessary to achieve identification. Simultaneity, implicit in most theories, is also revealing empirically, as summarized on figure 2. The overall effect of trade on business cycle synchronization is confirmed to be strong, but a sizable portion is found to actually work through intra-industry trade. Estimates of the link between interindustry trade and cycle correlations are smaller in magnitude, and not inconsistent with existing models.

Patterns of specialization have a sizable direct effect on business cycle correlation, as two economies with a similar economic structure are significantly more correlated ceteris paribus. This is shown to happen mostly because economies grow through evolving stages of diversification. Finally, ceteris paribus, business cycles in financially integrated regions are significantly more synchronized. This remains true even though financial integration tends to result in more specialized economies, and less synchronized cycles as a consequence.

The results obtain across countries and U.S. states. They hold for a variety of sectoral data sets, collected at different aggregation levels, for various measures of financial integration and trade linkages, and for various filtering methods. They suggest an additional item on the list of criteria characterizing optimal currency areas, namely, the economic structure of the putative member countries. They also provide some guidance on relevant strategies to model international business cycles, namely, due allowances for intra-industry trade, and international capital flows that are coordinated internationally.

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[52] The latter upper bound is obtained in simulations where both bilateral and world trade are assumed to increase.

[53] Kollman (2001) suggests that models with nominal rigidities and demand shocks might be more appropriate from the standpoint of reproducing the high observed correlation between T and ρ. Kollmann describes three diffusion channels for demand shocks: (i) a substitution effect, whereby agents substitute (depreciated) domestic for foreign goods in response to a domestic monetary shock, (ii) a quantity effect, whereby foreign aggregate demand increases because part of domestic demand falls on foreign goods, and (iii) a price effect, whereby the foreign price index decreases as it embeds prices of some domestic goods, now relatively cheaper. Only the quantity effect will be present in models with technology shocks. The evidence in this paper suggests that the measured effects of interindustry trade on ρ can be compatible even with models based on technology shocks only.

[54] A somewhat strange result in Table 6B is that trade is estimated to lower specialization, and thus indirectly increase business cycle correlations. This turns out to disappear once T^2 is used to measure bilateral trade, or a different filter is applied to the data. The lack of evidence supporting an indirect effect of trade on ρ is consistent with the conjecture in Frankel and Rose (1998).

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