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**ABSTRACT**

In this paper, we document evidence on the synchronization of business cycles across 16 countries over the past century and a quarter, demarcated into four exchange rate regimes. We find using three different methodologies that there is a secular trend towards increased synchronization for much of the twentieth century and that it occurs across diverse exchange rate regimes. This finding is in marked contrast to much of the recent literature, which has focused primarily on the evidence for the past 20 or 30 years and which has produced mixed results. We then considered a number of possible explanations for the observed pattern of increased synchronization. We first ascertained the role of shocks demarcated into country-specific (idiosyncratic) and global (common). Our key finding here is that global (common) shocks are the dominant influence across all regimes. The increasing importance of global shocks we posit reflects the forces of globalization, especially the integration of goods and services through international trade and the integration of financial markets. Our evidence shows a modest role for increasing bilateral trade in explaining synchronization, with stronger evidence for regional integration in Europe and North America but the evidence for the role of financial integration proxied by the removal of capital controls is inconclusive.



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## I. Introduction

The increasing worldwide integration of goods, capital and financial markets is widely believed to have led to more interdependence between national business cycles. In the recent encyclopedia *Business Cycles and Depressions,* Dore (1997) considers the synchronization of international cycles a stylized fact and argues that “[i]nstitutional changes such as free capital mobility, floating exchange rates, and the increase in international arbitrage and speculative activities have increased interdependence among the major capitalist nations, which is likely to lead to further synchronization of cycles.” From a theoretical perspective, however, the correlation between business cycle synchronization and integration is not necessarily positive. Krugman (1993) noted that stronger trade integration may lead to greater regional specialization, which can lead to less output synchronization with industry-specific shocks. Relatedly, Heathcote and Perri (2002) showed how increased financial integration may be an endogenous reaction to the regionalization of real sector linkages, as the latter allow for gains from the global diversification of asset portfolios.

The broadly synchronized recent downturn in the industrial countries has generally reinforced the notion of international business cycle synchronization. Paradoxically, however, the empirical evidence for the past three decades is so mixed that it remains difficult to make a strong case for the notion of increased or increasing business cycle linkages among industrial countries. Depending on the sample period, output correlations have even decreased in recent decades, largely on account of a remarkable cycle de-synchronization among the major industrial countries in the late 1980s and early 1990s (Helbling and Bayoumi, 2003). The empirical case against increased synchronization is made forcefully by Doyle and Faust (2002), who analyze changes in the comovements among the growth rates of G-7 countries since 1971 and found no evidence for significant increases in output correlations over time, even for Canada and the United States or for the euro area member countries. On the other hand, studies using dynamic factor models find evidence of increased international business cycle linkages. For example, Kose, Otrok and Whiteman (2002) report that a global factor explains larger shares of output variances in the G-7 countries during 1986-2001 than it does during the Bretton Woods period (1960-1972).

In this paper, we study international business cycle synchronization over 120 years, using annual data that covers four distinct eras with different international monetary regimes.[1] The four eras covered are 1880-1913 when much of the world adhered to the classical Gold Standard, the interwar period (1920-1938), the Bretton Woods regime of fixed but adjustable exchange rates (1948-1972), and the modern period of managed floating among the major currency areas (1973 to 2001). The annual data for 16 industrial countries that we use in this paper come from Mitchell (1998a, 1998b, and 1998c) and other sources. They were used by Bergman, Bordo, and Jonung (1998) and Bordo and Jonung (2001).[2] The IMF's *International Financial Statistics* was used to update the dataset to 2001.

[1] This approach follows Bergman, Bordo and Jonung (1998) and Backus and Kehoe (1992). Unlike these papers, we focus on international synchronization and devote less attention to the comparison of national business cycle properties such as output volatility or output-consumption comovements.

[2] These references also provide more details on the data. The countries included in our data set are Australia, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States. In some of the tables, we list the countries by their geographical proximity rather than in alphabetical order.

We believe that using a much longer data sample provides a much-needed broader and complementary perspective on business cycle synchronization. In the short term, much of the business cycle dynamics depends on the shock dynamics, which tends to overshadow the effects of integration. Changes in the latter, as a recent essay in the IMF's *World Economic Outlook* (2001) has pointed out, are often minor over short periods of time. Across the four eras that we examine, the variation in cross-border integration in the markets for goods, capital and financial assets has been considerably larger than over the last 20 years while the influence of particular shock realizations appears arguably to have been somewhat less important.

The paper is organized as follows. Section II discusses conceptual issues regarding international business cycle synchronization and provides the basic stylized facts. The subsequent section looks at the role that changes in this structure of shocks may have played in the observed changes in the synchronization of cycles. Section IV analyzes how changes in trade integration and the exchange rate regime have contributed to the observed increased international synchronization of business cycles.

## II. CROSS-COUNTRY BUSINESS CYCLE SYNCHRONIZATION OVER TIME

The notion of business cycles becoming increasingly synchronized across countries captures the observation that the timing and magnitudes of major changes in economic activity appear increasingly similar. For example, in the most recent slowdown, output growth started to weaken at about the same time in the major advanced economies (e.g., Helbling and Bayoumi, 2003). Despite its frequent use, however, synchronization as a formal concept has only recently been formally introduced into the business cycle literature, where comovements among cyclical time series have been the dominant object of analysis for many decades. Harding and Pagan (2002b) proposed a definition of cross-country cycle synchronization that is an offspring of the traditional concepts developed by the National Bureau of Economic Research (NBER) starting in the early 1920s. Specifically, they consider national business cycles to be synchronized if turning points in the corresponding reference cycles occur roughly at the same points in time. On this basis, they have derived a statistical measure, the concordance correlation, that allows one to test whether national cycles are significantly synchronized or not. This approach to measuring synchronization boils down to national business cycles being in the same phase—expansions and recessions—at about the same time.

Using NBER-based business cycle concepts has distinct advantages, including the fact that there is no need to apply detrending procedures and a rich collection of empirical studies dating back to the 1920s and 1930s. There are also distinct disadvantages. First, the use of non-linear filters and judgments in the derivation of turning points in the reference cycle often precludes standard statistical analysis based on the linear models widely used in economics.[3] Second, since Koopman's (1947) review of *Measuring Business Cycles* by

[3] However, recent advances in econometrics and information technology have facilitated the use of nonlinear models in empirical economic research. Diebold and Rudebusch (1999) address business cycle issues with (nonlinear) regime switching approaches.

Burns and Mitchell (1946), the mostly a-theoretical framework of the traditional NBER business cycle measures has been recognized as problematic by many. In particular, it remains difficult to link the phase-based NBER measures with structural models of macroeconomic fluctuations, including with the impulse-propagation approach to modeling business cycles that has increasingly dominated the academic literature. Finally, NBER reference cycle dates do not exist for all the countries in the sample from 1880-2001.

For an analytical paper, the costs associated with these disadvantages weigh heavily and we will adopt a more eclectic approach to measuring business cycle synchronization. In particular, we make extensive use of standard correlations and factors-based measures of synchronization, both measures which have frequently been used in the academic literature. We will, however, follow Harding and Pagan (2002, 2003), Stock and Watson (1999), and others by using real gross domestic product—or output in short—as the measure of aggregate economic activity or the business cycle rather than the synthetic reference cycle series used by the NBER. In addition, we will also examine the synchronization in industrial production. While we would not associate national business cycles with industrial production (IP), studying the latter is nevertheless interesting for two reasons. First, fluctuations in IP are more cyclical than fluctuations in some other sectors, especially services or government, which helps to understand the extent to which the cyclical sectors in an economy are interrelated with those abroad. Second, IP approximates fluctuations in the tradables sectors.

### A. Output Synchronization

#### 1. Concordance Correlations

Harding and Pagan (2002) proposed to measure the cross-country synchronization of business cycles in countries *i* and *j* by the correlation between the binary cycle indicator variables $S_{it}$ and $S_{jt.}$ The latter are 1-0 variables that take on the value 1 in period *t* if an economy is in an expansion and the value 0 if it is in a recession (determined on the basis of output developments, as discussed).[4] As the correlations are related to the so-called concordance statistic (the sum of the products of the binary cycle indicators), which determines the number of periods during which national cycles are in the same phase (as a fraction of the total number of periods in the sample), we will refer to them as concordance correlations.[5] If two cycles are perfectly synchronized, in the sense of being in the same state in every period, the concordance correlation coefficient $\tilde{n}_S$ is 1. If the two cycles are exactly in the opposite state in every period ($S_{it}$=1-$S_{jt}$, $t$=1...$T$), the correlation is -1. Finally, if the two cycles are uncorrelated, the correlation is 0.

We have calculated the binary business cycle state indicator variable $S_{it}$ for four eras for a panel of 16 industrial countries, for which GDP data are available from 1880. (As noted

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[4] See Chapter 3 of the April 2002 *World Economic Outlook* for an overview of business cycle concepts and methods.

[5] Specifically, Harding and Pagan (2002) demonstrate that the bilateral concordance correlation coefficient and its statistical significance can be estimated by regressing the binary cycle indicator of one country, adjusted by its standard deviations, on that of the other country.

above, real GDP is our measure of aggregate economic activity). With annual data, the determination of expansions and recessions is straightforward. A recession is defined as one or more consecutive years of negative real GDP growth, while an expansion consists of a year or more of positive growth. The resulting business cycle turning points broadly match the dates in the National Bureau of Economic Research (NBER) chronologies for the United States, the United Kingdom, France, and Germany (available in Glasner, 1997). The differences reflect the use by the NBER of higher frequency (monthly) data and a broader variety of indicators, such as employment and department store sales. On this basis, correlation coefficients between all country pairs and their significance were estimated.

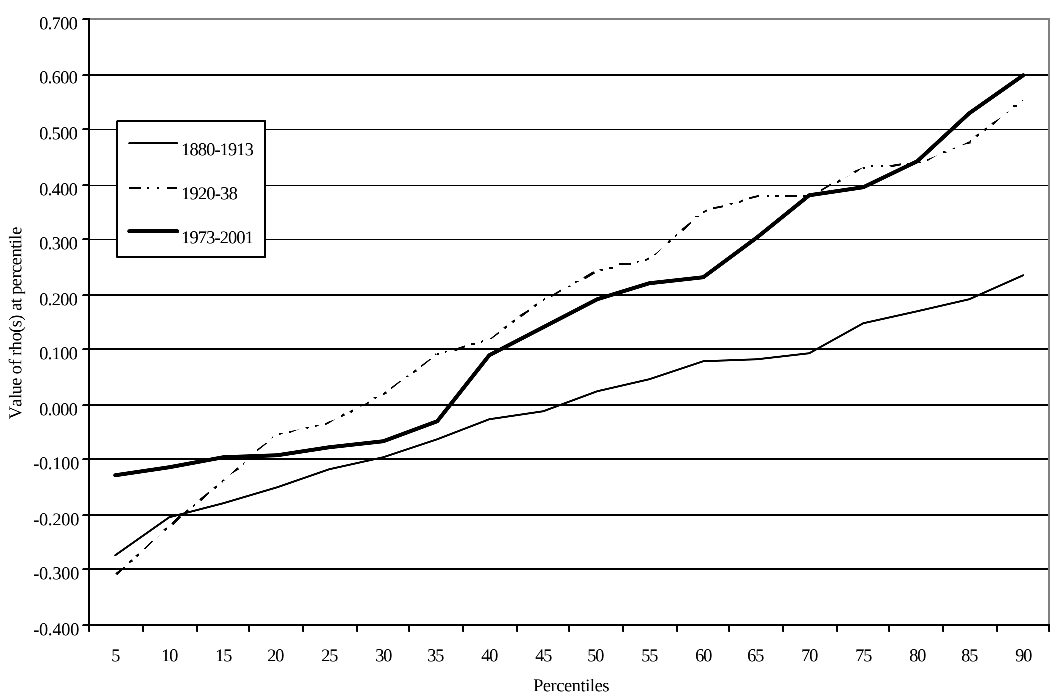
Figure 1 shows the concordance correlation coefficients by percentile for three of the four eras. The reason for leaving out the Bretton Woods era is that most countries did not experience any classical recession during this period. Hence, the business cycle indicator $S_{it}$ takes on the value 1 in every period. Naturally, the regression-based correlation coefficients can not be derived in the circumstances, as cycles (cycle states) become perfectly correlated without variation in the cycle phases. Table 1 shows the average concordance correlations by era, including for a number of country groups.

Comparing the distribution of the correlation coefficients by era suggests the following.

- In the Gold Standard era, the average of the correlation coefficients was just about zero, as about half of all the pairs of national business cycles were negatively related to each other while the other half was positively related to each other. Hence, on average, business cycles were not synchronized according to this measure.

- In the interwar and the post-Bretton Wood periods, more than half of all pairs of national business cycles were positively related to each other. On average, the correlation is about 0.2, suggesting generally synchronized business cycles during these eras. Nevertheless, more than one fourth of all cycles remained negatively related, which explains why the average "synchronization" remains relatively weak. Comparing the distributions of the two periods suggests no significant difference (which is confirmed by a Wilcoxon Rank sum test).

**Figure 1. Bilateral Output Concordance Correlations By Percentiles**



Overall, we find that synchronization was most pronounced in the Bretton Woods era, when many economies enjoyed uninterrupted expansions. As shown below, the finding is misleading because output comovements were not stronger in that era than in either the preceding or the subsequent era. The relevant key finding is that synchronization patterns appear to have systematically changed once the world went off the classical Gold Standard. During the latter era, cycles were, on average, uncorrelated while, beginning with the interwar period, they started to become synchronized on average.

## 2. Output Correlations

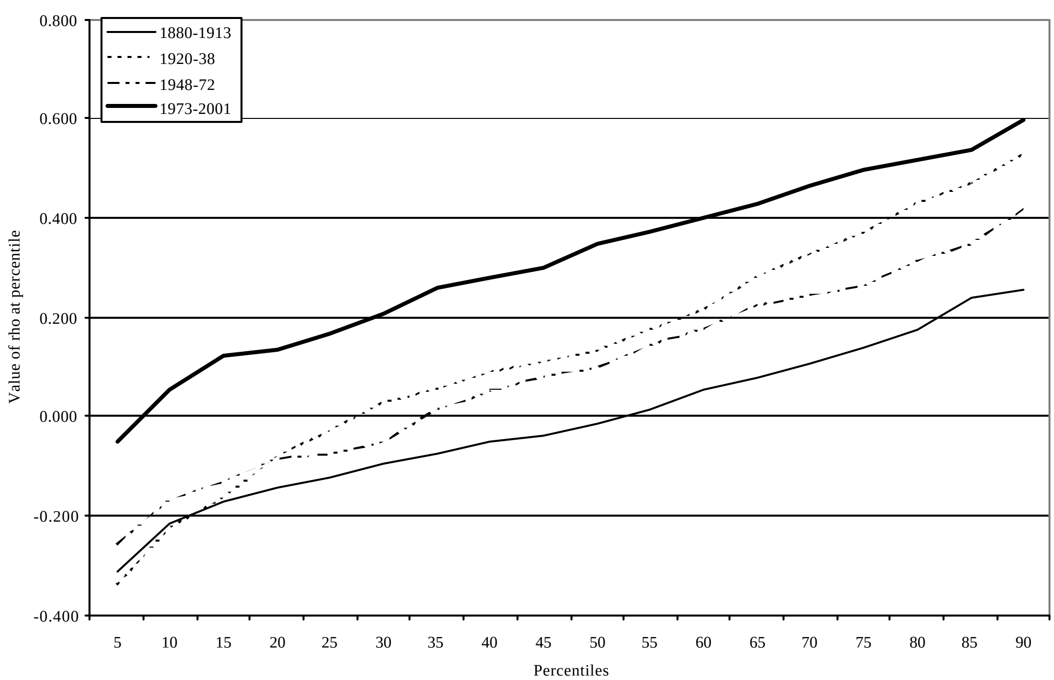
Output correlations have been the perhaps most frequently used measures of business cycle synchronization. According to this measure, national cycles are synchronized if they are positively and significantly correlated with each other. The higher are the positive correlations, the more synchronized are the cycles. Compared with concordance correlations, measuring synchronization with standard contemporaneous correlations is more stringent, as the latter require similarities in both the direction and magnitudes of output changes.[6]

Figure 2 shows the correlation coefficients for log output growth by percentile for the four eras. The distribution of the correlation coefficients differs substantially from era to era. In

[6] As we use annual data, it is not surprising that there is little evidence of strong lead-lag relationships among the log output growth rates in our dataset.

particular, there has been a tendency toward higher, positive output correlations, not just a one-time level increase in synchronization, as observed in the previous subsection based on concordance correlations. During the Gold Standard, about one half of all country pairs were characterized by negative output correlations and the average output correlation coefficient is about 0 (Table 2). A first important step toward synchronization occurred during the interwar period, when the share of negative correlations fell below 30 percent while the average correlation increased to about 0.15. A subsequent reversal during the Bretton Woods era was small, and correlations remained, on average, above those found for the Gold Standard era. A second important increase then occurred during 1973-2001, when less than 10 percent of all country pairs were characterized by negative output correlations and the average correlation was 0.33.

**Figure 2. Bilateral Output Correlation Coefficients By Percentile**



Are these changes over time statistically different? This question is relevant since the confidence intervals for the bilateral correlation coefficients are relatively wide given the few observations per era.[7] We used both nonparametric and parametric tests to address the issue. (Nonparametric) Wilcoxon Rank sum tests suggest that the upward shifts in the distribution of the correlation coefficients are significant at the 5 percent level for the interwar period (compared to the Gold Standard era) and for the modern floating era (compared to both the interwar and the Bretton Woods eras). The downward shift in the distribution of correlation coefficients from the interwar to the Bretton Woods eras is only significant at the 10 percent level.

The Wilcoxon rank sum test does not require any assumption about the underlying distribution of the correlation coefficients. In practice, however, it is typically assumed that the correlation coefficients of a vector series of log first differences of outputs reflect an underlying multivariate normal distribution, at least asymptotically. For this reason, we also formally tested for the equality of the covariance and correlation matrices over subsequent periods using the tests proposed by Jennrich (1970). In a first step, we tested for the significance of the changes in the average correlation coefficients. The changes between the Gold Standard and the interwar eras and between the Bretton Woods and the modern floating eras, respectively, are statistically significant while the decline between the interwar and the

[7] The sampling standard deviation of estimated correlation coefficients depends on the size of the estimated coefficient and the number of observations. Given the former, small samples tend to amplify the sampling uncertainty greatly. For example, for a correlation coefficient of 0.15—the average for the interwar period—the standard deviation for a sample of 20 observations is 0.23. With 50 and 100 sample observations, the standard deviations decline to 0.14 and 0.10, respectively.

Bretton Woods eras is insignificant (Table 3). Given the small number of observations per era, this is a strong result.

Beyond average correlations, however, the statistical significance of the changes in output comovements in general is more ambiguous. Regarding the six possible pairs of covariance matrices for the four eras, all but two are statistically significantly different from each other at the 5 percent level (Table 4), which bears on the factor model-based approach to measuring synchronization discussed in the next subsection. The pairs that are insignificantly different are 1880-1913 vs. 1926-1938 and 1880-1913 vs. 1951-1972, respectively. Regarding correlation matrices, only two are statistically different at the 5 percent level (1926-1938 vs. 1973-2001 and 1951-1972 vs. 1973-2001, respectively). This, however, reflects to some extent "substitution" among country pairs, as some correlations increased while others decreased.

For smaller groups, especially the (old and new) core countries and European countries, the changes in the correlation matrices from era to era are generally statistically significantly different from era to era, except for the pair 1880-1913 vs. 1926-1938. We attribute the fact that the changes between those two eras are often insignificant to the few number of observations for the interwar era, which tends to reduce the sampling precision (as noted in footnote 7). Overall, the results of all the tests support the notion of a secular increase in business cycle synchronization.

So far, we have looked at business cycle synchronization through a global lens, noting the increased synchronization without consideration for other factors. However, one would expect that synchronization patterns differ considerably across groups of countries, depending on factors such as “gravity” or country size. The evidence clearly illustrates that the extent to which gravity has shaped the synchronization trends depends on the region (Table 2). For core European countries (the old “EEC”) and Continental European countries, the increase in business cycle synchronization was clearly much sharper than the general increase. At the other end of gravitas, business cycle synchronization between Japan and the other countries in the panel has increased by less. In particular, there is no evidence for an increase between the Bretton Woods era and the modern floating rate period.

The fact that the increase for all Continental European countries was smaller than that for the Core European countries suggests that the forces of gravity are affected by common policies, preferential trading agreements, and specific currency arrangements. The increase in correlations among the Anglo-Saxon countries is also remarkable even though it seems more difficult to attribute this to forces of gravity.[8] While we do not believe that common institutions and heritage among the Anglo-Saxon countries account directly for the increased synchronization, as Otto et al. (2002) have argued, they likely have fostered similar patterns

[8] The emergence of strong business cycle linkages among core European countries and among the Anglo-Saxon countries was noted, among others, by Helbling and Bayoumi (2003) and Stock and Watson (2003).

in the transmission of shocks through what appear to be similar, market-based financial systems.[9]

While the regional perspective reinforces the notion of a trend increase, it should be noted that stark regional differences have only really emerged during the modern floating rate period. Forces of gravity were not a factor behind business cycle synchronization during the classical Gold Standard, as differences in correlations among regions were minor, with the high correlation between Canada and the United States and, to a much lesser extent, among the Scandinavian countries, being the main exceptions. During the Bretton Woods period, increased regional synchronization began to emerge in the core European countries. Interestingly, the increased synchronization during the interwar period was primarily on account of an increased synchronization between the cycles in the United States and other countries, which in turn seems to reflect the equity boom bust cycle and its effects from the mid-1920s to the mid-1930s.

So far, all the cycle correlations we have studied were based on log first differences of output. Does the detrending method matter for our findings? Naturally, high frequency noise is not a great concern, given our panels of annual data, but it is possible that the increases in cycle correlations from the 19$^{th}$ to the end of the 20$^{th}$ century really reflect changes in trend comovements. For this reason, we have also looked at correlation patterns in bandpass-

[9] See Bayoumi and Edison (2003) on the distinction between market- and bank-based financial systems.

filtered log output data (Table 5).[10] The results show that the detrending method makes little difference and that the same principal changes in the patterns of cycle synchronization are found with bandpass-filtered output data.

### 3. The Importance of Common Factors

Another measure of business cycle synchronization is related to the notion of common factors. Many macroeconomists would probably agree that international business cycle linkages are best understood as a small set of factors that are common to all countries and that explain a substantial fraction of fluctuations in major macroeconomic aggregates. Baxter (1995), for example, refers to the world component to business cycles.[11] The common factors themselves reflect a combination of global shocks affecting all countries and country-specific disturbances with significant spillover effects. It is, therefore, quite natural to examine international business cycle linkages with factor models.[12]

In our context, we have estimated the following factor model for each country $i$ in the sample:

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[10] We used the parameter setting recommended by Baxter and King (1999) for annual data.

[11] This is akin to Stock and Watson's (1989) suggestions to use the first common factor in a panel of business cycle indicators as the reference cycle.

[12] See, among others, Gerlach (1988), Gregory, Head, and Raynauld (1997), Lumsdaine and Prasad (1999), Kose, Otrok, and Whiteman (2002), Monfort and others (2002), and Helbling and Bayoumi (2003).

$$y_{it} = l_i' F_t + e_{it} \tag{2.1}$$

where $y_{it}$ denotes (log) output growth in country $i$ at time $t$, $F_t$ is a vector of factors, $l_i'$ is a vector of factor loading for country $i$, and $å_{it}$ an idiosyncratic error term. In this set-up, the factors in $F_t$ are common to all countries but each country has specific factor loadings, allowing for country-specific responses to shocks to the factor. Subsequently, we will call the product $l_i' F$ the common factor(s) for country $i$. This model is what is called a static approximate factor model in the literature (e.g., Bai and Ng, 2002). It is static because it does not allow for dynamics in the relationship between output and the factors. It is approximate because the set-up allows for serial correlation and heteroskedasticity in the idiosyncratic components. While seemingly restrictive, these assumptions seem compatible with the properties of the data we use, given little evidence of significant serial correlation in output growth rates during all four eras.

A frequently used measure of output synchronization is the fraction of the output variance that is explained by common factors. Following practices in the literature, we use the first common factor as a measure of the common elements in cross-country business cycle fluctuations (Table 6).[13] The results are based on log output growth rates (the results obtained with bandpass-filtered log output series are, again, very similar). The role that common factors play in explaining output fluctuations has clearly increased over time. The share of

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[13] Bai and Ng (2002) proposed to use information criteria to determine the appropriate number of factors. However, their Monte Carlo simulations show that for panel datasets where the cross-sectional and time dimensions are as low as in ours, the tests are not very reliable and tend to imply too high a number of factors.

the variance in output fluctuations attributable to fluctuations in the common factor doubled from about 20 percent during the Gold Standard era to about 40 percent during the modern era of flexible exchange rates.

### B. The Synchronization of Cycles in Industrial Production

Unfortunately, IP data are only available for some of the countries in our panel. Based on a small eight country panel, we calculated the same set of synchronization measures for fluctuations in IP indices over all four eras. With a larger panel, we calculated them for the post-World War II eras. The results are qualitatively very similar with those for output fluctuations, and we only show the correlation coefficients (Table 7). Compared with output fluctuations, all measures clearly suggest a much stronger increase in synchronization over time. Strikingly, the evidence for much stronger increases in regional synchronization compared to the average increase is weaker than in the case of output linkages.

## III. Explaining Increased Synchronization: The Role of Shocks

Using three measures of synchronization, we found evidence of increased cross-country business cycle synchronization over time among industrial countries. From an impulse-propagation perspective, the increased synchronization could reflect changes in the nature of the impulses (the “shocks”) driving the economies, especially changes in their cross-country correlations, changes in the transmission channels and mechanisms, resulting inter alia from increased integration, or, most likely, both.

Disentangling the relative contributions of the changes in the correlation of shocks and changes in the transmission channels to the changes in output correlations is difficult, however, as this would require a comprehensive structural model of the economy that can be estimated empirically. Such a model, which would need to allow for factors such as changes in trade and financial integration and a multitude of shocks, seems beyond our reach, given the current state of the art in multi-country modelling. Financial integration, for example, is not yet satisfactorily accounted for in any of the leading multi-country models.

Against this background, we will proceed with a more modest research agenda. In this section, we will focus on deriving measures of the impulses driving each economy and study the changes in their variance-covariance structure. On this basis, we will then attempt to assess the extent to which changes in this structure of shocks may help to explain the observed changes in the synchronization of cycles. In the next section, we will focus on how changes in integration and the policy environment may have shaped changes in business cycle synchronization.

### A. Global Shocks versus Idiosyncratic Shocks

In a first exercise, we will address the issue of whether there is evidence that global shocks have been driving the increased business cycle synchronization. This is a natural hypothesis, given the evidence of increased globalization, that is, economic interdependence through trade in goods, services, and assets. To structure our discussion, the following canonical, simple model is a helpful illustrative device:[14]

$$\begin{pmatrix} y_{1t} \\ y_{2t} \end{pmatrix} = \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix} \begin{pmatrix} y_{1t-1} \\ y_{2t-1} \end{pmatrix} + \begin{pmatrix} \boldsymbol{n}_{1t} \\ \boldsymbol{n}_{2t} \end{pmatrix} \tag{3.1}$$

where $y_{it}$ denotes the log output growth rate in country *i*. Following Stock and Watson (2003), the error vector í is assumed to be determined by the following factor structure:[15]

$$\begin{pmatrix} \boldsymbol{n}_{1t} \\ \boldsymbol{n}_{2t} \end{pmatrix} = G\boldsymbol{z}_t + \begin{pmatrix} \boldsymbol{h}_{1t} \\ \boldsymbol{h}_{2t} \end{pmatrix} \tag{3.2}$$

In this factor VAR model, *æ* is a global shock and *ç<sub></sub>*$_i$ is a country-specific idiosyncratic shock, both of which are assumed to be white noise processes. *G* is a vector of factor loadings. According to this model, increased output synchronization—as measured by output correlation—between countries 1 and 2 can, ceteris paribus, arise for three reasons:

- Increases in the variance of the global shock relative to the country specific idiosyncratic shocks.

[14] Canova and Dellas (1993) show how a very stylized two-country real business cycle model implies such a bi-variate vector autoregressive representation.

[15] In contrast, in the standard dynamic factor model, the coefficient matrix A is usually assumed to be zero and the dynamics arise from the factors, which are modelled as autoregressive processes.

- Increases in the covariance of the idiosyncratic shocks $ç_1$ and $ç_2$.

- Increases in the "transmission" coefficients $a_{12}$ and $a_{21}$, which determine the spillover effects that shocks in country 1 have on country 2 and vice versa.[16]

Extending the simple bivariate model (3.1) and (3.2) to a general model that includes all the 16 countries in our sample proved to be difficult. First, given few observations for each era, very few degrees of freedom would be left if the model were estimated even with one lag. For the interwar era, the number of common observations for all 16 countries is even less than the number of parameters, so that comparability of the model across eras could not be ensured. Second, in our annual data, the degree of autocorrelation in output growth is generally low in all eras. While there are some intrinsic dynamics in the cross-country output dynamics, they are generally insignificant, as the null hypothesis of $a_{11}=a_{12}=a_{21}=a_{22}=0$ could usually not be rejected in bi-variate VARs. In the circumstances, we used two simplified versions of the general factor VAR model for the analysis:

- *The center country model*. In this version, the equation for each country's real GDP growth includes lagged own GDP growth and lagged GDP growth in the center country (the United Kingdom in the Gold Standard era and the United States in the

[16] In the simple model of Canova and Dellas (op. cit.), these coefficients follow from the production structure, as foreign intermediate goods are needed to produce the final consumption goods.

other eras). The rationale behind this model is that idiosyncratic shocks in the center country can be transmitted through the traditional channels while idiosyncratic shocks elsewhere have only limited effects on other countries.

- *The trade linkages model*. In this version, the equation for each country's real GDP growth includes lagged GDP growth and lagged GDP growth in important trading partner countries (the ones reported in Mitchell (1998a, 1998b, and 1998c, as explained below). The rationale behind this model is, of course, straightforward.

We estimated the models with a two-step, semi-parametric procedure. In the first, we used SURE estimators to obtain the coefficients and the residual series $\text{í}_{it}$. In a second step, we used the approximate factor model approach of section 2.3 to obtain the global shock *æ* and the idiosyncratic shocks $\text{ç}_{i}$ from the residual series $\text{í}_{it}$. Both models turned out to be roughly similar in terms of information criteria for all eras, although the restrictions implied by the center model compared to the trade model were rejected by standard likelihood ratio tests (we note, though, that the values of log determinants of the variance-covariance matrix of the simple de-meaned log growth rates are generally lower than the information criteria for either model in each era).

A first set of results concerns the issue of whether the moderation in the volatility of national output growth rates over time in the postwar period reflects primarily reductions in the volatility of global shocks.[17] Table 8 shows the average standard deviations of the national output series (log growth rates) and compares them with the estimated variances of the global shock and the average idiosyncratic shock. As was to be expected, the direction of change in the standard deviations of both global and idiosyncratic shocks generally follows that of the average standard deviation of output growth. The most striking feature is the change in the difference between the standard deviations of the two types of shocks. While idiosyncratic shocks were clearly larger than global shocks during the Gold Standard and the interwar period, their magnitudes became about equal during the modern floating era. In this sense, global shocks have become more important in shaping national output dynamics.

A second set of results concerns the issue of whether global shocks or spillovers drive the observed increases in output synchronization. Table 9 presents averages of 4-step variance decompositions of the output growth rates for the four eras, distinguishing between the shares of total output variance explained by the global shock, the idiosyncratic shocks, and transmission. To have a variance decomposition based on orthogonal shocks, we used the diagonal of the variance-covariance matrix of the $ç_i$ series rather than the full matrix.[18] Hence, we assume that the transmission of shocks—which can be both due to global shocks

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[17] The general moderation in the amplitude of output fluctuations has been analyzed by Blanchard and Simon (2001).

[18] We note that the cross-correlation among idiosyncratic shocks is very minor.

and idiosyncratic shocks—occurs with a one-year lag.[19] Output synchronization due to a global shock is thus immediate, while that due to transmission occurs gradually.

The variance decomposition suggests the following. First, idiosyncratic shocks have become less important in shaping each country's output dynamics. Second, both global shocks and transmission have become more important. Third, the relative importance of the last two factors in accounting for the lesser role of idiosyncratic shocks depends on the model. The center model suggests that the increases in the variance share of the global shock account for most of the reduction in the variance share of the idiosyncratic shocks. The increase accounted for by the transmission of idiosyncratic shocks is minor. On the other hand, the trade model suggests that both increases in the variance shares of the global factor and the transmission account for the reduction in the variance share of idiosyncratic shocks. However, it is noteworthy that shock transmission appears more important for peripheral countries than for core countries in the trade model. For the core countries, the increase in the variance share related to global shocks accounts for most of the decrease in the share explained by idiosyncratic shocks.

Our factor VAR models thus suggest that the increased business cycle linkages among core countries, as measured by output correlations, largely reflect the dynamics of common global shocks. Remarkably, the increased importance of transmission for peripheral countries only

[19] The shares of the global shocks and the idiosyncratic shocks reported in Table 9 are the shares explained by each country's own autoregressive structure in response to each of the two shocks. We do not, therefore, distinguish between the transmission of global shocks and idiosyncratic shocks.

arises in the trade model. This suggests that it is not transmission from the center country, a channel that operates in both models, that accounts for the increased variance share of transmission. It is rather the intra-European transmission that matters in the modern floating era, a fact that seems consistent with the above average output synchronization among European countries reported in Table 2.

For a further understanding of the role of global shocks, Figure 3 is instructive. Each panel shows the global shocks from the trade model (solid lines) for an era, supplemented by dotted lines depicting the global shocks from the simple factor model discussed in section 2.3 and bars showing time dummies—the equivalent of global shocks—from an error components model that are significant at the 5 percent level.[20] [21] We estimated the latter to obtain a measure for large and important shocks. The global shocks implied by the two factor-based approaches—the trade model and the simple factor model—are surprisingly similar except for the interwar period, although the shocks from the trade model are generally smaller in magnitude. The latter finding is not that surprising since the possibility of transmission implies lower shock variances with equal output variances.

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[20] Naturally, only the product $ËF_t$ is identified in this factor model. We normalized the square of the factor loadings, i.e., $Ë'Ë/16=1$, to identify $F_t$. We believe this to be the natural normalization, as it allows for comparable variances between outputs and factors. The alternative would be to normalize the factor variance to 1 (Bai and Ng, 2002).

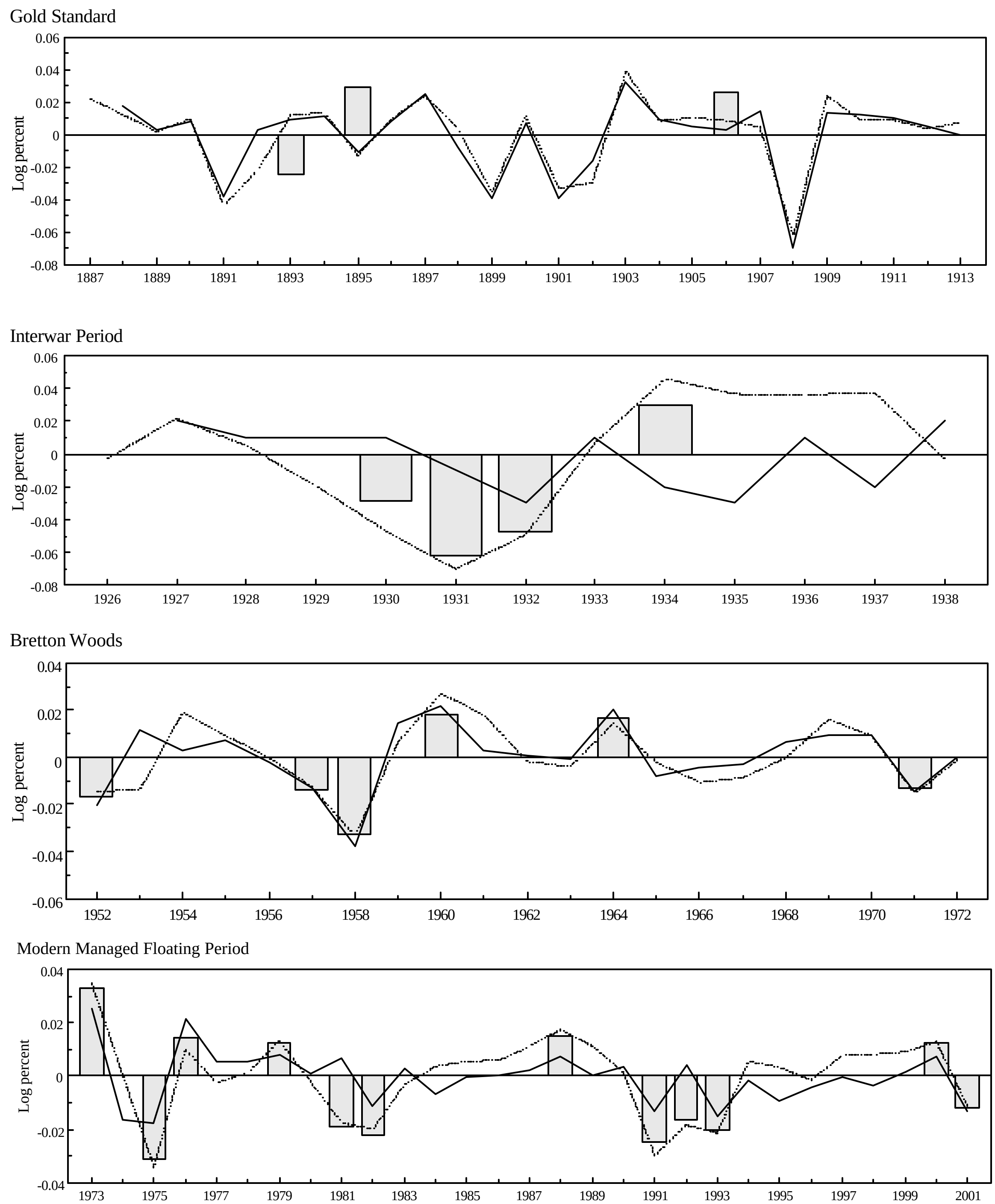
[21] We estimated the following traditional error component model with our panel dataset:

$$y_{i,t} = l_t + h_{i,t}$$

where $ë_t$ denotes a time dummy taking on the value 1 in time t and $ç_{i,t}$ a shock specific to country *i*.

Figure 3. Global Shocks, 1887-2001

(Solid Lines: trade model; dotted lines: simple factor model; and bars: significant time dummies)




Source: Authors' calculcations (see text for details).

The general picture emerging from Figure 3 is that global shocks appear noticeably important at times of world-wide downturns, suggesting an asymmetry between downturns and upturns.[22] However, as the differences in the global shocks for the interwar period illustrate, simple factor models can occasionally identify global shocks when it is really the transmission of a large idiosyncratic shock in one of the large economies that is behind synchronized output fluctuations, a problem noted by Stock and Watson (2003).[23] The Great Depression is a case in point. It originated largely in the United States, but was transmitted abroad rapidly through “the golden fetters” of the Gold Standard (Friedman and Schwartz, 1963, and Eichengreen, 1992). Similarly, the “Volcker shock” in 1981/82 also hit other countries hard at a time when the after effects of the 1979-80 oil shock were still widely felt. Indeed, this corroborates findings reported in IMF (2002), where episodes of broadly synchronized classical recessions were found to have generally coincided with recessions in the center country.[24]

Overall, our results are consistent with both globalization and business cycle moderation. With increased economic and financial interdependence through trade and financial linkages, the scope for global shocks or the rapid transmission of shocks in the center countries has

[22] This corroborates Helbling and Bayoumi (2003), who found a similar result for the G-7 countries during 1973-2001 using quarterly data and a dynamic factor model to isolate common cycles.

[23] While global shocks continue to be important in the trade model, it can be noted that the magnitudes of these shocks in this model are smaller compared to the simple factor model, highlighting the importance of transmission in output synchronization.

[24] Helbling and Bayoumi (2003) have shown that during 1973-2001, the tendency for major growth contractions to be synchronized and coincide with recessions in the center country also holds.

clearly increased. In addition, with global shocks, floating exchange rates do not provide much scope for insulation, since shocks affect all countries in similar ways. At the same time, business cycle amplitudes have clearly moderated during the post-World War II period, reflecting, among other factors, changes in sectoral structure, automatic stabilizers, the use of lender of last resort operations, and the use of discretionary counter-cyclical policies.[25] In this context, it is interesting to note that the volatility of idiosyncratic shocks has decreased more than that of global shocks. Among other factors, this finding is consistent with the notion that the changes in the sectoral structure and the use of automatic stabilizers as well as other counter-cyclical policies have been fairly similar across the industrial countries.

### B. Demand and Supply Shocks over Time

In a second exercise, we study what are often called “demand” and “supply” shocks by era for each country and analyze their correlation structure. Following Blanchard and Quah (1989), the two kinds of shocks are derived on the basis of the residuals of bivariate vector autoregressive (VAR) models for output growth and inflation for each country by era.[26] The demand and supply shocks are orthogonal and are identified by differences in their long-run

[25] See McConnell and Perez-Quiros (2000), Ahmed, Levin and Wilson (2002), Blanchard and Simon (2001), IMF (2002), Dalsgaard, Elmeskov, and Park (2002), and Stock and Watson (2002) for recent studies examining the volatility of output fluctuations in industrial countries, especially the United States, and the reasons thereof.

[26] Bordo and Jonung (1996) and Bayoumi and Eichengreen (1994) used the same approach with similar data for other purposes.

cumulative effects on growth.[27] Demand shocks only have temporary effects on real GDP while supply shocks have permanent effects on real GDP. We would like to emphasize, however, that we merely follow conventional wisdom in the terminology and that we will not focus on whether the impulse response functions meet the standard overidentifying restrictions.[28] We are primarily interested in obtaining shock series that can be used to analyze whether the increased importance of global shocks concerns primarily permanent or temporary shocks or both. In addition, we hope to corroborate our earlier findings using another approach that does not require assumptions about the factor structure and relies on country-specific information in the data only.

For the underlying VAR models, two versions were estimated for each country.[29] For the first version, the lag length for each country model was determined on the basis of the standard likelihood ratio tests, starting from an initial lag length of 3. For the second version, the lag length for each country model was determined on the basis of the Bayesian information criterion, also starting from an initial lag length of 3. The results are very similar, and we only show those based on information criteria minimization.

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[27] In addition to this long-run behavioral restriction, the identification requires additional restrictions. Following standard practice, the restrictions used ensure that the variance-covariance matrix of the orthogonalized shocks matches that of the VAR residuals.

[28] Bordo and Jonung (1996) discuss these and related issues in the context of a similar exercise that also uses long time series going back to the late 19th century.

[29] We note that, unlike in the case of bi-variate VARs for output growth, the autoregressive dynamics are more important in this set-up.

For supply shocks, the average correlation has clearly increased over time for both VAR versions (Table 10). From being roughly uncorrelated during the classical Gold Standard, they became more correlated over time, although the average correlation and the change between the Bretton Woods era and the modern floating rate era remain small in magnitude. With regard to the regional breakdown, we note that the increase in the correlations of supply or permanent shocks is particularly strong for the continental European and, even more so, for the core European countries and the United States and Canada. In addition, the increase in supply shock correlations has been steady for the case of the United States and Canada while for the European countries, it largely occurred between the Bretton Woods era and the modern floating rate era, which is consistent with the patterns in European integration.

For demand shocks, the average correlation followed a u-shaped pattern (Table 11). After the classical Gold Standard, the correlation of demand shocks fell during the interwar and the Bretton Woods eras and then increased again during the modern floating rate era to a level above the initial one. Regarding the regional breakdown, we note that the increase in the correlations of demand or temporary shocks is particularly strong for the continental European countries and, even more so, for the core European countries. This regional increase in correlations occurred largely between the Bretton Woods era and the modern floating rate era, which is consistent with the coordination of policies required with the pegging of currencies within Europe.

Overall, the direction of changes in the average correlation of the demand and supply shocks obtained with the first versions of the VAR seems consistent with the direction of changes in output correlations. Under the classical Gold Standard, output correlations were, on average, zero, which, according to our shock decomposition exercise, coincided with roughly uncorrelated shocks. The subsequent increase in output correlations reflected in part increased shock correlations. The fact that output correlations increased by more than shock correlations could be explained with larger spillover effects, as cross-border economic and financial linkages in the modern floating rate period have also risen compared to the classical Gold Standard era.

Finally, we also examined the factor structure in our panel of shocks series (Table 12). This analysis suggests that the shares of the shock variances explained by the global component contained in each national shocks series increased over time for demand or temporary shocks. For permanent shocks, the evidence is less clear-cut. In particular, relative to the Gold Standard era, the increase is very small only. The fact that the global component plays an increasing role in demand shock realizations is consistent with our earlier finding of the global synchronization of recessions, the effects of which are, after all, reversed in the subsequent recoveries.

## IV. Explaining Increased Synchronization II: Integration and Policies

Changes in the nature of shocks are only one reason for the observed increased in business cycle synchronization. In parallel with the factor model-based literature on the sources of international business cycle linkages, the role of structural factors such as trade integration and exchange rate regimes in determining shock correlations as well as the transmission coefficients $a_{ij}$ has also been studied. For example, as noted by Canova and Dellas (1993), the extent to which two or more countries are linked through trade is an important determinant of the $a_{ij}$ coefficients and thus of the strength of business cycle linkages. In this section, we will focus on how changes in integration and the policy environment may have shaped changes in business cycle synchronization.

### A. Trade Integration

Starting with Canova and Dellas (1993), the role of trade interdependence in explaining international business cycle linkages has received considerable attention in the literature.[30] Frankel and Rose (1998) found that in the period from 1959-1993, OECD countries with closer trade links tended to have more tightly correlated business cycles. In this subsection, we follow Frankel and Rose's methodology and examine the linkages between business cycle synchronization and trade links for the four eras that are the subject of our paper. In particular, we will try to address the questions of whether and to what extent the observed changes in trade linkages can explain the changes in business cycle synchronization.

Using data from Mitchell (1998a,1998b, and 1998c) for the Gold Standard and the interwar period and from the *IMF's* Direction of Trade Statistics for the Bretton Woods era and

[30] See, for example, Frankel and Rose (1998), Otto, Voss, and Willard (2001), and Imbs (2003).

modern floating era, we constructed a measure of the trade intensity between countries *i* and *j* proposed by Frankel and Rose (1998):[31]

$$w_{ijt}^{Y} = \frac{X_{ijt} + M_{ijt}}{Y_{it} + Y_{jt}} \tag{4.2}$$

where *X* and *M* denote exports and imports, respectively, and where a double subscript *ij* stands for bilateral trade values. Y denotes nominal GDP. Unfortunately, based on the limited information provided in Mitchell (1998a, 1998b, and 1998c), we could not construct the two measures for all the 120 correlation observations that we have for each era. We will thus show results for a sample of 59 observations for each of the four eras and for a sample of all the 120 country pair combinations for the post-World War II eras.

With these measures, we estimated cross-sectional regressions of the form for each era ô:

$$f\,(corr_{ij,t}) = a + b\,\ln(\bar{w}_{ij,t}^{k}) + e_{ij,t} \tag{4.3}$$

where the bar over the trade intensity measure $w^k$ (k=T, Y) indicates that it is an era average and where the function f(·) maps the output correlation coefficient $corr_{ij}$ from the interval [-

[31] We also constructed the other measure proposed by Frenkel and Rose (1998), where bilateral trade is normalized by total trade:

$$w_{ijt}^{T} = \frac{X_{ijt} + M_{ijt}}{X_{it} + M_{it} + X_{jt} + M_{jt}}$$

where simple subscripts *i* or *j* indicates aggregate trade values. We found that this measure lacked a clear time trend, which suggests that the direction of trade has been quite stable over time.

1,1] to [-  ,  ] so that standard assumptions in the linear regression model are met.[32] As do Frankel and Rose, we treat trade intensity as an endogenous variable and use gravity variables as instrumental variables. Specifically, the following three instruments are used: the natural logarithm of the distance between the main business centre in each country, a dummy variable for common borders, and a dummy variable for common language.

Table 13 shows the results for both trade intensity measures and for both the smaller and larger sample, as discussed earlier. The estimated $\hat{a}$ coefficients all have the right sign but they are not always significant. However, the insignificant coefficients are confined to the smaller panel covering all four eras, which may reflect small sample problems, as we are missing bilateral trade relations among the smaller European countries in particular. The simple regressions suggest the following observations.

- There is substantial variation in the slope coefficients (and the constant terms) across eras. In fact, the assumption of identical slope coefficients and constant terms is rejected for both samples.

- As do Frankel and Rose, we find that changes in bilateral trade intensities are estimated to have large effects on output correlations. For example, a one percentage

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[32] We use the function $f(x) = \ln\left(\frac{1+x}{1-x}\right)$ following Otto, Voss, and Willard (2001). Frankel and Rose (1998) used raw correlation measures.

point increase in trade intensity generally raises output correlations by more than 3½ percentage points in the case of bilateral trade normalized by GDP.

- The difference in the estimated â coefficients for the Gold Standard and interwar eras is small. Between the Bretton Woods era and the modern floating rate era, however, the synchronization effects of higher trade intensity appear to have increased, which may suggest that the synchronization effects of trade intensity depend on other factors as well. We will take up this issue below, but note in the meantime that this finding is consistent with the argument that stronger trade linkages in recent years have reflected common policies, including, for example, those related to European monetary integration.[33] It is also consistent with our earlier finding of increased structural shock correlations, since bilateral output correlations are determined by the product of the transmission coefficients, which in turn can be considered to be functions of trade intensities, and shock correlations.[34]

- Finally, similar to Otto, Voss, and Willard (2001), we find that while they are significant determinants, trade intensities alone explain relatively little of the overall variation in bilateral output correlations, especially for the interwar and the Bretton Woods eras. For the Gold Standard era, bilateral trade intensities appear to explain about the same share of output correlations as during the modern floating era with the

[33] This argument was advanced by Bayoumi and Eichengreen (1993).

[34] In the general autoregressive model (3.1), the positive interaction between transmission coefficients and shock correlations holds for a wide range of the admissible parameter values.

first measure. Interestingly, the explanatory power of the gravity variables in the first stage regressions appears to increase over time.

So far, the results suggest that bilateral trade intensities explain only a small share of the differences found in bilateral output correlations for each era. Can changes in trade intensities from era to era explain the increased international business cycle synchronization? In Table 14, we show the estimated impact of the actual changes in trade intensities between eras on bilateral output correlations (the second set of columns in the table corresponds to the larger sample for the postwar eras). Since the restriction of identical slope coefficients between eras can generally be rejected, we use both the estimated values for the current and the previous era to generate predicted changes.[35]

The actual changes in the average bilateral output correlation are generally much larger than the changes predicted by changes in trade intensities. The model is successful in explaining the increase in output correlations from the interwar to the Bretton Woods eras, where the rise in trade intensity explains roughly half of the increase or more, depending on the â coefficient used in the calculations. This finding is consistent with the notion that trade liberalization after World War II, when the earlier increase in the restrictiveness of trade regimes was reversed under the umbrella of the General Agreement on Tariffs and Trade (Irwin, 1995), has played an important role in shaping business cycle synchronization. The

[35] We do not use changes in the constant between eras in the calculation of the predicted changes.

model is less successful in explaining the strong increase in output correlations from the Bretton Woods era to the modern floating rate period. We interpret this as possible evidence that the momentum of trade liberalization has slowed down in recent decades, as efforts in the most recent GATT/WTO rounds have shifted from tariff reductions on industrial goods towards nontariff barriers and trade in agriculture and services (Irwin, op. cit.). It also may be that increased financial integration played a bigger role in shaping output synchronization than trade integration during recent decades.

## B. Asset Market Integration

Increasing asset market integration is another channel through which globalization can affect international business cycle synchronization. Baxter and Crucini (1995), for example, show how asset market integration affects the spillover effects of country-specific shocks and, thus, output correlations. Unfortunately, asset market integration remains difficult to measure. Bilateral data in (net) asset trade is all but unavailable, especially for a period covering our four eras. Measuring asset market integration through the correlation of asset returns may suffer from problems of reverse causality. We decided, therefore, to use a very crude indicator to measure asset market integration. Based on the annual capital control dummy variables for each country prepared by Bordo et al (2001), we derived an indicator, which we will refer to as *CC*, that measures the number of years (as a fraction of the total number of years covered by each era) during which capital flows were subject to restrictions.

With this indicator, we first estimated the simple regression equation:

$$f\,(corr_{ij,t}) = a + d\,\overline{CC}_{ij,t} + e_{ij,t} \tag{4.4}$$

where the function $f\,(\cdot)$ is the same as above for each era but the Gold Standard era. The latter is because our dummy variables suggest that none of the countries in our sample imposed capital controls. Ordinary least squares was used to estimate the coefficients. Subsequently, we also estimated the following equation that combines trade and asset market integration:

$$f\,(corr_{ij,t}) = a + b\ \ln(\overline{w}^{Y}_{ij,t}) + d\,\overline{CC}_{ij,t} + e_{ij,t} \tag{4.5}$$

for both post-World War II eras. By focusing on the postwar period only, we have 120 observations for the trade intensity measures, which is preferable, given that we suspect small sample problems for the data set covering all four eras. As above, we treat the trade intensity measure $w^{Y}$ as an endogenous variable using the gravity variables discussed above as instruments (and our capital control indicator *CC*).

The results are shown in Table 15. The ä coefficients have the right signs for the interwar and modern floating rate eras but are insignificant. The conclusion is that adding the capital control indicator to the trade equation does not add to the explanatory power of the equation. We interpret these results mainly as a reflection of possible shortcomings of our dummy variable rather than as a rejection of the hypothesis that financial market integration has been a major factor behind international business cycle synchronization.[36]

### C. Exchange Rate Policy

One of the main arguments, if not the main argument, in favor of flexible exchange rates has been that they allow a country to insulate itself from external shocks through an independent monetary policy. With fixed exchange rates, the scope for pro-cyclical spillovers can be expected to be larger, unless capital controls or other restrictions allow for some monetary independence.

To some extent, we have accounted for differences in the exchange rate and external policy environment by distinguishing between four eras with different international monetary regimes. Nevertheless, there are interesting cross-sectional variations in the exchange rate regime in each era. For example, Italy went off the Gold Standard in 1894 and did not return to a gold parity until 1928. To the extent that the cross-sectional variation in exchange rate regimes has changed over time, this may be a factor that could explain changes in the

[36] Other measures of financial market integration have been suggested to us. In particular, our discussant and other conference participants suggested that we use Lewis' (1996) consumption sharing "betas" to construct an alternative measure. The consumption betas are the coefficients from a panel regression of consumption growth on GDP growth, which, with perfect cross-country consumption insurance, should not be statistically significantly different from zero. However, we faced the problem that the betas are not only positive and statistically significantly different from zero (a problem also faced by Lewis) but also increasing in magnitude from era to era, which contradicts the widely-accepted hypothesis of a u-shaped pattern in international financial integration (Obstfeld and Taylor, 1998 and 2003, and Bordo, Eichengreen, and Kim, 1998). We intend to explore this puzzle in future research.

average business cycle synchronization from era to era. To estimate the effects of exchange rate regimes, we have constructed a dummy variable, which, again, is based on the annual exchange rate regime indicators for each country prepared by Bordo et al (2001). The indicator, which we will refer to as *Z*, measures the number of years (as a fraction of the total number of years covered by each era) during which the exchange rate between two countries was pegged (disregarding re-alignments in the case of fixed but adjustable rates).

With this indicator, we estimated the simple equation:

$$f(corr_{ij,t}) = a + g\,\overline{Z}_{ij,t} + d\,\overline{CC}_{ij,t} + e_{ij,t} \qquad (4.6)$$

for each era.[37] We would expect the *ã* coefficient to be positive for the reasons noted above. Despite earlier problems, we kept the capital control indicator *CC* in the equation, since it is important to control for these effects. Subsequently, we also added the exchange regime indicator to the trade equation but only for the two post-World War II eras for the reason explained above.

The results are shown in Table 16. The *ã* coefficients generally have the right sign and are significant at the 10 or 5 percent levels, except for the Gold Standard era. The latter may be explained by the fact that only four out of the 16 countries had not pegged their currencies to gold (at least during some time) in this era. For the interwar period, the *ã* coefficient has the

[37] As explained above, the capital control indicator *CC* is redundant for the Gold Standard era.

wrong sign, which may reflect the fact that the countries hanging on to the Gold Standard after the United States and other countries went off resorted to exchange controls and increased trade restrictions to maintain their peg. The *ä* coefficients are generally of the right sign, except for the Bretton Woods period.

The exchange rate regime effect is not robust to small variations in the specification.[38] For example, adding the trade intensity variable reduces the size of the exchange rate regime effect considerably. Similarly, the significance also changes once that additional variable is included. Finally, we would like to note one interesting aspect for the modern floating rate period. If a European country dummy is included, the exchange rate regime indicator becomes insignificant, which can be interpreted as evidence that the exchange rate pegging within Europe (all bilateral fixed exchange rate regimes in the modern floating rate period are confined to that continent for our panel of countries) has been really important or that the exchange rate indicator picks up the effects of European integration more generally.

[38] Frankel and Rose (1998), who in one specification added a dummy variable for country pairs with fixed exchange rates, argued that the exchange rate regime choice is also endogenous and used the same gravity variables as instruments to correct for this endogeneity. We remain somewhat skeptical about the relevance of these instruments and note the technical problems in using a 1-0 variable as an endogenous variable in a linear regression. We have, however, used instrumental variable estimators with the same instruments to check the robustness of our results. For the simple equation (0.6), the results are robust. In fact, the size of the exchange rate effect increases somewhat, and the standard errors of the *ã* coefficients decrease. For the interwar period, the sign of the *ã* coefficient changes. If the trade intensity variable is included, the results do not remain robust. In particular, as Frankel and Rose, we find that the *ã* coefficients are negative and insignificant.

## V. Conclusion

In this paper, we have documented evidence on the synchronization of business cycles across 16 countries over the past century and a quarter, demarcated into four exchange rate regimes. We find using three different methodologies that there is a secular trend towards increased synchronization for much of the twentieth century and that it occurs across diverse exchange rate regimes. This finding is of interest because it is in marked contrast to much of the recent literature, which has focused primarily on the evidence for the past 20 or 30 years and which has produced mixed results.

We then considered a number of possible explanations for the observed pattern of increased synchronization. We first ascertained the role of shocks demarcated into country-specific (idiosyncratic) and global (common). Our key finding here is that global (common) shocks are the dominant influence across all regimes. We note, however, that it remains difficult to distinguish between "true" global shocks and major shocks in the center country(ies). For example, shocks in the largest country, the U.S., were, unsurprisingly, a key factor in the worldwide Great Depression.

This finding coupled with earlier evidence produced by ourselves and others that business cycles since World War II have become less volatile, less frequent and asymmetric with a tendency towards recoveries exceeding downturns in duration, has some interesting implications. We suggest that what may be occurring is that the weakening in national business cycles since World War II coupled with the diminution of idiosyncratic shocks reflect the forces discussed by Zarnowitz (1992) and others, such as changes in the composition of output, the advent of automatic stabilizers, improvements in discretionary monetary and fiscal policy, the implementation of effective lenders of last resort and a financial safety net, and the proliferation of private risk sharing instruments.[39]

At the same time, the increasing importance of global shocks we posit reflects the forces of globalization, especially the integration of goods and services through international trade (Findlay and O'Rourke 2003) and the integration of financial markets (Obstfeld and Taylor 2003). We present evidence showing a modest role for increasing bilateral trade in explaining synchronization, with stronger evidence for regional integration in Europe and North America. Evidence for the role of financial integration proxied by the removal of capital controls is inconclusive.

Finally, we began considering explicitly the role of the policy regime in explaining the pattern of synchronization. We find little evidence for the prediction that adhering to fixed

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[39] We would like to emphasize that the decrease in idiosyncratic shock volatility does not necessarily imply that it is the volatility of the underlying "deep structural" shocks (e.g., a widespread drought) that we are measuring. Given that we derive the shocks on the basis of GDP series alone, it means that the effects of these deep shocks on output have diminished. There is, however, some evidence that shocks that are widely perceived as being exogenous may have an endogenous component. Barsky and Kilian (2001), for example, argue that the sharp increase in real oil prices in the 1970s was in part a reaction to the earlier massive expansion of the world money supply.

exchange rates fosters synchronization except in the period since 1973, and these results appear driven largely by the process of European Monetary Union.

What are the policy lessons to be gleaned from these findings? One lesson from the dampening of national cycles and the diminution in idiosyncratic shocks is that to the extent that they reflect sound domestic macro policy and the creation of an environment conducive to the development of both private and public risk sharing institutions and instruments that these policies should continue to be fostered.

A second lesson is that globalization seems to be associated with the creation of a global business cycle. Whether policies should be developed at the global level to counter it is another matter. Regarding monetary policy, experience suggest that the key to success has been that policy makers get their policy objectives right. In addition, for monetary policy to be effective, it is also critical that policy makers get the shocks right. In this sense, policy coordination is likely to be very important. As Obstfeld and Rogoff (2002) have argued, there is probably little gain from policy coordination beyond this, provided that central banks have the right policy objectives. The generally negative experience with policy coordination in the 1970s and 1980s supports this view (Frankel, 1990).

A third lesson, to paraphrase Forrest Gump, is that shocks happen! We live in a stochastic world and shocks generate business cycles via diffuse propagation mechanisms. Moreover productivity shocks occur in a non linear fashion, asset markets overshoot and people are at times over optimistic and other times over pessimistic. In this reality the best strategy is to encourage the development of private market mechanisms to insure against cyclical risks, in the case of incomplete markets related to market failures to provide public insurance, and to maintain stable and predictable macro policies.

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Table 1. Average Concordance Correlations By Region and Era
(Based on first differences of log output)

| Region (Number of obs.) | 1880-1913 | 1920-38 | 1973-2001 |
|---|---|---|---|
| All countries (120) | 0.02 | 0.19 | 0.20 |
| Core countries (15)[a] | -0.01 | 0.14 | 0.28 |
| Peripheral countries (45)[a] | 0.01 | 0.19 | 0.13 |
| One country core, one periphery (60) | 0.03 | 0.21 | 0.21 |
| Continental European countries only (55) | -0.01 | 0.05 | 0.51 |
| One country Continental European (55) | 0.01 | 0.20 | 0.10 |
| Core European countries only (6)[b] | -0.02 | 0.08 | 0.69 |
| One country Core European (48)[b] | 0.00 | 0.15 | 0.25 |
| One Country North America 28) | 0.03 | 0.35 | 0.18 |
| One country Japan (15) | 0.03 | 0.20 | -0.01 |
| Anglo-Saxon Countries (6)[c] | 0.16 | 0.26 | 0.37 |
| Memorandum items: | | | |
| USA-Canada | 0.71 | 0.77 | 0.44 |
| Scandinavian countries only (6) | 0.08 | 0.36 | 0.04 |
| One country Scandinavian (48) | 0.02 | 0.25 | 0.12 |

[a] Core countries comprise France, Germany, the United Kingdom, the United States, the Netherlands, and Switzerland during 1880-1913 and 1920-39 and the G-7 countries afterwards.

[b] Core European countries comprise France, Germany, the Netherlands, and Switzerland during 1880-1913 and 1920-39 and the EEC countries in the panel (France, Germany, Italy, the Netherlands) afterwards.

[c] Comprises Australia, Canada, the United Kingdom, and the United States.

Table 2. Average Output Correlations By Region and Era
(Based on first differences of log output)

| Region (Number of obs.) | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 |
|---|---|---|---|---|
| All countries (120) | 0.01 | 0.15 | 0.12 | 0.33 |
| Core countries (15) | 0.02 | 0.13 | 0.15 | 0.45 |
| Peripheral countries (45) | 0.02 | 0.21 | 0.11 | 0.27 |
| One country core, one periphery (60) | 0.01 | 0.12 | 0.11 | 0.33 |
| Continental European countries only (28) | 0.01 | 0.08 | 0.10 | 0.49 |
| One country Continental European (64) | 0.01 | 0.17 | 0.13 | 0.26 |
| Core European countries only (6) | 0.07 | 0.10 | 0.25 | 0.59 |
| One country Core European (48) | -0.02 | 0.09 | 0.14 | 0.36 |
| One Country North America 28) | 0.00 | 0.28 | 0.04 | 0.38 |
| One country Japan (15) | -0.08 | 0.05 | 0.16 | 0.17 |
| Anglo-Saxon Countries (6) | 0.12 | 0.29 | 0.17 | 0.63 |
| Memorandum items: | | | | |
| USA-Canada | 0.55 | 0.86 | 0.61 | 0.77 |
| Scandinavian countries only (6) | 0.11 | 0.48 | 0.17 | 0.31 |
| One country Scandinavian (48) | 0.00 | 0.17 | 0.11 | 0.25 |

Note: See Table 1 for definitions of regions and country groups.

Table 3. Jennrich Test for Equality of Average Correlation Coefficients
(Marginal significance levels)

| Region (Number of obs.) | 1880-1913 vs. 1926-38 | 1926-38 1952-72 | 1952-72 1973-2001 | 1880-1913 1952-72 | 1880-1913 1973-2001 | 1926-38 1973-2001 |
|---|---|---|---|---|---|---|
| All countries (120) | 0.01 | 0.85 | 0.06 | 0.02 | 0.00 | 0.03 |
| Core countries (15) | | | | | | |
| Old | 0.30 | 0.40 | 0.25 | 0.02 | 0.00 | 0.08 |
| New | 0.04 | 0.85 | 0.02 | 0.05 | 0.00 | 0.09 |
| Continental European countries (28) | 0.26 | 0.19 | 0.01 | 0.01 | 0.00 | 0.00 |
| Core European countries (6) | | | | | | |
| Old | 0.92 | 0.12 | 0.50 | 0.06 | 0.00 | 0.05 |
| New | 0.98 | 0.15 | 0.01 | 0.07 | 0.00 | 0.00 |
| Anglo-Saxon Countries (6) | 0.49 | 0.76 | 0.01 | 0.68 | 0.00 | 0.04 |
| Memorandum items: | | | | | | |
| USA-Canada | 0.18 | 0.48 | 0.66 | 0.34 | 0.14 | 0.70 |
| Scandinavian countries only (6) | 0.11 | 0.25 | 0.57 | 0.60 | 0.25 | 0.51 |

Note: See Table 1 for definition of regions and country groups and Jennrich (1970) on the computation of the variance-covariance matrix of the correlation coefficients.

Table 4. Jennrich Test for Covariance and Correlation Matrix Equality
(Marginal significance levels)

| Region (Number of obs.) | 1880-1913 vs. 1926-38 | 1926-38 1952-72 | 1952-72 1973-2001 | 1880-1913 1952-72 | 1880-1913 1973-2001 | 1926-38 1973-2001 |
|---|---|---|---|---|---|---|
| | | | Covariance matrices | | | |
| All countries (120) | 0.86 | 0.01 | 0.01 | 0.28 | 0.00 | 0.00 |
| Core countries (15) | | | | | | |
| old | 0.04 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| new | 0.07 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| Continental European countries (28) | 0.06 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| Core European countries (6) | | | | | | |
| old | 0.00 | 0.00 | 0.02 | 0.00 | 0.00 | 0.00 |
| new | 0.01 | 0.00 | 0.02 | 0.00 | 0.00 | 0.00 |
| Anglo-Saxon Countries (6) | 0.44 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| Memorandum items: | | | | | | |
| USA-Canada | 0.36 | 0.00 | 0.96 | 0.00 | 0.00 | 0.00 |
| Scandinavian countries only (6) | 0.08 | 0.01 | 0.53 | 0.05 | 0.00 | 0.02 |
| | | | Correlation matrices | | | |
| All countries (120) | 0.86 | 0.01 | 0.01 | 0.28 | 0.00 | 0.00 |
| Core countries (15) | | | | | | |
| old | 0.04 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| new | 0.07 | 0.00 | 0.01 | 0.00 | 0.00 | 0.00 |
| Continental European countries (28) | 0.06 | 0.01 | 0.00 | 0.00 | 0.00 | 0.00 |
| Core European countries (6) | | | | | | |
| old | 0.00 | 0.00 | 0.02 | 0.00 | 0.00 | 0.00 |
| new | 0.01 | 0.00 | 0.02 | 0.00 | 0.00 | 0.00 |
| Anglo-Saxon Countries (6) | 0.44 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| Memorandum items: | | | | | | |
| USA-Canada | 0.36 | 0.00 | 0.96 | 0.00 | 0.00 | 0.00 |
| Scandinavian countries only (6) | 0.08 | 0.01 | 0.53 | 0.05 | 0.00 | 0.02 |

Note: See Table 1 for definition of regions and country groups and Jennrich (1970) on the computation of the variance-covariance matrix of the correlation coefficients.

Table 5. Average Output Correlations By Region and Era
(Based on bandpass-filtered log output series)

| Region (Number of obs.) | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 |
|---|---|---|---|---|
| All countries (120) | 0.01 | 0.21 | 0.16 | 0.30 |
| Core countries (15) | -0.01 | 0.17 | 0.21 | 0.42 |
| Peripheral countries (45) | 0.01 | 0.22 | 0.15 | 0.23 |
| One country core, one periphery (60) | 0.02 | 0.22 | 0.15 | 0.30 |
| Continental European countries only (28) | 0.00 | 0.09 | 0.22 | 0.44 |
| One country Continental European (64) | -0.01 | 0.22 | 0.17 | 0.21 |
| Core European countries only (6) | 0.06 | 0.11 | 0.27 | 0.59 |
| One country Core European (48) | -0.01 | 0.14 | 0.22 | 0.34 |
| One Country North America 28) | -0.01 | 0.36 | 0.01 | 0.35 |
| One country Japan (15) | -0.08 | 0.18 | 0.24 | 0.16 |
| Anglo-Saxon Countries (6) | 0.16 | 0.38 | 0.11 | 0.63 |
| Memorandum items: | | | | |
| USA-Canada | 0.52 | 0.86 | 0.68 | 0.75 |
| Scandinavian countries only (6) | 0.11 | 0.31 | 0.16 | 0.26 |
| One country Scandinavian (48) | 0.00 | 0.24 | 0.13 | 0.24 |

Note: See Table 1 for definitions of regions and country groups.

Table 6. Fraction of Variance of Output Fluctuations Explained by First Principal Component
(Based on first differences of log output)

| Country | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 |
|---|---|---|---|---|
| Denmark | 0.51 | 0.30 | 0.17 | 0.31 |
| Finland | 0.55 | 0.55 | 0.45 | 0.33 |
| Norway | 0.00 | 0.22 | 0.00 | 0.06 |
| Sweden | 0.14 | 0.64 | 0.31 | 0.31 |
| France | 0.15 | 0.14 | 0.13 | 0.69 |
| Germany | 0.44 | 0.47 | 0.17 | 0.45 |
| Italy | 0.05 | 0.28 | 0.20 | 0.63 |
| Switzerland | 0.08 | 0.31 | 0.78 | 0.35 |
| Netherlands | 0.31 | 0.46 | 0.44 | 0.49 |
| Portugal | 0.04 | 0.01 | 0.07 | 0.38 |
| Spain | 0.04 | 0.00 | 0.05 | 0.44 |
| UK | 0.06 | 0.05 | 0.43 | 0.51 |
| Australia | 0.25 | 0.11 | 0.10 | 0.23 |
| US | 0.01 | 0.71 | 0.09 | 0.51 |
| Canada | 0.02 | 0.82 | 0.00 | 0.50 |
| Japan | 0.55 | 0.12 | 0.19 | 0.11 |
| Average | 0.20 | 0.32 | 0.22 | 0.39 |

Table 7. Industrial Production: Average Correlations By Region and Era
(Based on first differences of log series)

| Region (Number of obs.) | 8 Country Panel [a] | | | | 14 Country Panel [b] | |
|---|---|---|---|---|---|---|
| | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 | 1948-72 | 1973-2001 |
| All countries (28/91) | 0.11 | 0.35 | 0.23 | 0.57 | 0.27 | 0.50 |
| Core countries (6/21) | 0.28 | 0.38 | 0.30 | 0.63 | 0.30 | 0.61 |
| Peripheral countries (6/21) | 0.03 | 0.30 | 0.15 | 0.52 | 0.28 | 0.42 |
| One country core, one periphery (16/49) | 0.08 | 0.35 | 0.16 | 0.49 | 0.25 | 0.49 |
| Continental European countries only (10/15) | 0.16 | 0.35 | 0.19 | 0.62 | 0.15 | 0.65 |
| One country Continental European (15/48) | 0.04 | 0.34 | 0.25 | 0.53 | 0.28 | 0.47 |
| One Country North America (7/24) | 0.04 | 0.47 | 0.17 | 0.58 | 0.20 | 0.54 |
| One country Japan (7/13) | 0.03 | 0.14 | 0.32 | 0.55 | 0.28 | 0.52 |
| Anglo-Saxon Countries (1/6) | 0.30 | 0.67 | 0.32 | 0.65 | 0.37 | 0.66 |
| Memorandum items: | | | | | | |
| Output correlations for identical groups | | | | | | |
| All countries (28/91) | -0.01 | 0.05 | 0.13 | 0.38 | 0.13 | 0.34 |
| Core countries (6/21) | -0.03 | 0.07 | 0.16 | 0.46 | 0.15 | 0.45 |
| Peripheral countries (6/21) | 0.09 | 0.15 | -0.07 | 0.40 | 0.12 | 0.29 |
| One country core, one periphery (16/49) | -0.04 | 0.00 | 0.12 | 0.29 | 0.12 | 0.31 |

[a] In the panel of 8 countries, core countries comprise France, Germany, the United Kingdom, and the United States during 1880-1913 and 1920-39 and the G-7 countries excluding Canada afterwards while the Anglo-Saxon countries include the United Kingdom and the United States.

[b] Excludes Portugal and Switzerland. See Table 1 for country groupings.

Table 8. Output and Shock Standard Deviations By Era
(Based on first differences of log output)

| Aggregates[1] | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 |
|---|---|---|---|---|
| Average output growth | 0.047 | 0.057 | 0.027 | 0.023 |
| *Center Model* | | | | |
| Global factor | 0.023 | 0.028 | 0.014 | 0.012 |
| Idiosyncratic shocks | 0.037 | 0.040 | 0.020 | 0.016 |
| *Trade Model* | | | | |
| Global factor | 0.022 | 0.017 | 0.013 | 0.010 |
| Idiosyncratic shocks | 0.034 | 0.024 | 0.018 | 0.014 |

[1] Unweighted averages across countries.

Table 9. Variance Decomposition of Output Growth
(Fractions of forecast error variance; based on first differences of log output; simple averages over countries)

| Countries (Number of obs.) | 1887-1913 | | | 1926-38 | | | 1952-72 | | | 1973-2001 | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Global Factor | Trans-mission | Idio-syncratic Shocks | Global Factor | Trans-mission | Idio-syncratic Shocks | Global Factor | Trans-mission | Idio-syncratic Shocks | Global Factor | Trans-mission | Idio-syncratic Shocks |
| Center model | | | | | | | | | | | | |
| All countries (16) | | | | | | | | | | | | |
| 1-step ahead | 0.22 | 0.0 | 0.78 | 0.24 | 0.0 | 0.76 | 0.28 | 0.0 | 0.72 | 0.36 | 0.0 | 0.64 |
| 2-step ahead | 0.22 | 0.03 | 0.75 | 0.21 | 0.12 | 0.67 | 0.27 | 0.04 | 0.69 | 0.33 | 0.10 | 0.57 |
| 3-step ahead | 0.21 | 0.05 | 0.74 | 0.21 | 0.14 | 0.65 | 0.27 | 0.04 | 0.69 | 0.33 | 0.10 | 0.57 |
| 4-step ahead | 0.21 | 0.05 | 0.74 | 0.21 | 0.15 | 0.64 | 0.27 | 0.04 | 0.69 | 0.33 | 0.10 | 0.57 |
| New core countries (7) | | | | | | | | | | | | |
| 1-step ahead | 0.19 | 0.0 | 0.81 | 0.31 | 0.0 | 0.69 | 0.24 | 0.0 | 0.76 | 0.45 | 0.0 | 0.55 |
| 2-step ahead | 0.19 | 0.02 | 0.79 | 0.29 | 0.07 | 0.64 | 0.23 | 0.02 | 0.75 | 0.42 | 0.08 | 0.51 |
| 3-step ahead | 0.19 | 0.03 | 0.78 | 0.29 | 0.09 | 0.62 | 0.23 | 0.02 | 0.75 | 0.42 | 0.08 | 0.51 |
| 4-step ahead | 0.19 | 0.03 | 0.78 | 0.28 | 0.10 | 0.62 | 0.23 | 0.02 | 0.75 | 0.42 | 0.08 | 0.51 |
| Trade model | | | | | | | | | | | | |
| All countries (16) | | | | | | | | | | | | |
| 1-step ahead | 0.23 | 0.0 | 0.77 | 0.24 | 0.0 | 0.76 | 0.32 | 0.0 | 0.68 | 0.31 | 0.0 | 0.69 |
| 2-step ahead | 0.20 | 0.13 | 0.67 | 0.17 | 0.40 | 0.43 | 0.28 | 0.16 | 0.56 | 0.25 | 0.24 | 0.51 |
| 3-step ahead | 0.19 | 0.17 | 0.64 | 0.16 | 0.47 | 0.37 | 0.27 | 0.19 | 0.54 | 0.24 | 0.29 | 0.47 |
| 4-step ahead | 0.19 | 0.18 | 0.63 | 0.15 | 0.52 | 0.33 | 0.27 | 0.20 | 0.53 | 0.23 | 0.32 | 0.45 |
| New core countries (7) | | | | | | | | | | | | |
| 1-step ahead | 0.27 | 0.0 | 0.73 | 0.22 | 0.0 | 0.78 | 0.27 | 0.0 | 0.73 | 0.46 | 0.0 | 0.54 |
| 2-step ahead | 0.25 | 0.12 | 0.63 | 0.18 | 0.26 | 0.56 | 0.24 | 0.11 | 0.65 | 0.38 | 0.17 | 0.45 |
| 3-step ahead | 0.24 | 0.16 | 0.60 | 0.17 | 0.33 | 0.50 | 0.24 | 0.12 | 0.64 | 0.37 | 0.19 | 0.44 |
| 4-step ahead | 0.24 | 0.16 | 0.60 | 0.15 | 0.40 | 0.45 | 0.24 | 0.13 | 0.63 | 0.37 | 0.19 | 0.44 |

Note: See Table 1 for regions and country groupings.

Table 10. Average Supply Shock Correlations By Region and Era
(Based on first differences of log output; lag length determined with BIC minimization)

| Region (Number of obs.) | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 |
|---|---|---|---|---|
| All countries (120) | 0.01 | 0.09 | 0.09 | 0.11 |
| Core countries (15) | -0.02 | -0.04 | 0.14 | 0.14 |
| Peripheral countries (45) | -0.01 | 0.13 | 0.08 | 0.09 |
| One country core, one periphery (60) | 0.04 | 0.08 | 0.09 | 0.12 |
| Continental European countries only (28) | 0.06 | 0.02 | 0.09 | 0.24 |
| One country Continental European (64) | 0.00 | 0.08 | 0.11 | 0.04 |
| Core European countries only (6) | 0.07 | 0.04 | 0.07 | 0.35 |
| One country Core European (48) | 0.01 | 0.02 | 0.13 | 0.11 |
| One Country North America 28) | -0.03 | 0.10 | 0.02 | 0.10 |
| One country Japan (15) | -0.02 | 0.06 | 0.18 | 0.07 |
| Anglo-Saxon Countries (6) | 0.01 | 0.12 | 0.13 | 0.22 |
| Memorandum items: | | | | |
| USA-Canada | 0.13 | 0.30 | 0.55 | 0.52 |
| Scandinavian countries only (6) | -0.01 | 0.44 | 0.02 | 0.12 |
| One country Scandinavian (48) | -0.01 | 0.11 | 0.09 | 0.08 |

Note: See Table 1 for definition of regions and country groups.

Table 11. Average Demand Shock Correlations By Region and Era
(Based on first differences of log output; lag length determined with BIC minimization)

| Region (Number of obs.) | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 |
|---|---|---|---|---|
| All countries (120) | 0.16 | 0.05 | 0.20 | 0.25 |
| Core countries (15) | 0.11 | 0.11 | 0.08 | 0.22 |
| Peripheral countries (45) | 0.14 | 0.05 | 0.32 | 0.25 |
| One country core, one periphery (60) | 0.18 | 0.04 | 0.18 | 0.25 |
| Continental European countries only (28) | 0.19 | 0.04 | 0.13 | 0.35 |
| One country Continental European (64) | 0.15 | 0.07 | 0.21 | 0.23 |
| Core European countries only (6) | 0.26 | 0.15 | 0.05 | 0.61 |
| One country Core European (48) | 0.20 | 0.08 | 0.20 | 0.29 |
| One Country North America 28) | 0.07 | -0.20 | 0.23 | 0.12 |
| One country Japan (15) | 0.09 | 0.14 | -0.23 | 0.10 |
| Anglo-Saxon Countries (6) | 0.00 | -0.13 | 0.36 | 0.13 |
| Memorandum items: | | | | |
| USA-Canada | -0.17 | 0.35 | 0.13 | -0.19 |
| Scandinavian countries only (6) | 0.41 | 0.24 | 0.50 | 0.41 |
| One country Scandinavian (48) | 0.17 | 0.09 | 0.26 | 0.27 |

Note: See Table 1 for definition of regions and country groups.

Table 12. Fraction of Demand And Supply Shock Variance Explained by First Principal Component
(Lag length in underlying VAR determined with BIC minimization)

| Type of Shocks (Model Version) | 1880-1913 | 1920-38 | 1948-72 | 1973-2001 |
|---|---|---|---|---|
| Demand shocks | | | | |
| Average | 0.25 | 0.40 | 0.33 | 0.35 |
| Supply shocks | | | | |
| Average | 0.18 | 0.27 | 0.22 | 0.21 |

Table 13. Bilateral Trade-Output Regressions[a]

(Standard errors in parenthesis; coefficients significant at the 5% level are bolded)

| | Limited Sample (59 Observations) | | | | | Full Sample (120 Observations) | | |
|---|---|---|---|---|---|---|---|---|
| | 1880-1913 | 1920-38 | 1950-72 | 1973-2001 | Panel | 1950-72 | 1973-2001 | Panel |
| $\hat{a}$[b] | 0.120 | 0.134 | 0.087 | **0.210** | **0.138** | **0.123** | **0.205** | **0.168** |
| | (0.071) | (0.154) | (0.099) | (0.088) | (0.057) | (0.056) | (0.053) | (0.042) |
| R-square [c] | -0.013 | -0.017 | 0.022 | 0.071 | 0.028 | 0.066 | 0.16 | 0.123 |
| R-square first-stage regressions | 0.179 | 0.187 | 0.398 | 0.460 | 0.271 | 0.431 | 0.551 | 0.484 |
| F-test[d] | | | | | 0.000 | | | 0.000 |
| Number of observations | 59 | 59 | 59 | 59 | 236 | 120 | 120 | 240 |

[a] Constant is not reported. Standard errors are heteroskedasticity-consistent.

[b] Instrumental variable estimate.

[c] From an OLS regression of the trade equation.

[d] Marginal significance level from an F-test of the restriction that the panel coefficients are not significantly different from those obtained for each period.

Table 14. Actual and Predicted Change in Output Correlations[a]
(All variables in percent or percentage points)

| | Limited Sample (59 Observations) | | | | Full Sample | |
|---|---|---|---|---|---|---|
| | 1880-1913 | 1920-38 | 1950-72 | 1973-2001 | 1950-72 | 1973-2001 |
| Average trade intensity | 0.757 | 0.491 | 0.698 | 0.956 | 0.460 | 0.612 |
| *Change in average trade intensity* | | -0.266 | 0.207 | 0.258 | | 0.152 |
| Average actual change in output correlations | | 13.409 | 2.332 | 22.815 | | 21.615 |
| Predicted change | | | | | | |
| Based on *â* of previous era | | -1.962 | 1.695 | 0.992 | | 1.271 |
| Based on *â* of current era | | -2.198 | 1.110 | 2.348 | | 2.093 |
| Based on â of panel | | -2.265 | 1.744 | 1.560 | | 1.724 |

[a] All data including correlation coefficients are reported in percent (percentage points). The *â* coefficients are those shown in Table 15.

Table 15. Asset Market Integration, Bilateral Trade and Output Correlations[a]
Bolded signifies significance at the 5% level and bolded-italicized significance at the 10 percent level)

| | 1920-38 | 1950-72 | 1973-2001 | 1950-72 | 1973-2001 |
|---|---|---|---|---|---|
| ä | -0.083 | 0.066 | -0.089 | 0.034 | -0.039 |
| | ( 0.165) | ( 0.108) | ( 0.110) | ( 0.100) | ( 0.104) |
| *â*[b] | | | | ***0.119*** | **0.191** |
| | | | | ( 0.059) | ( 0.047) |
| R-square [c] | 0.002 | 0.003 | 0.007 | 0.059 | 0.155 |
| R-square first-stage regressions | | | | 0.324 | 0.447 |
| Number of observations | 120 | 120 | 120 | 120 | 120 |

[a] Constant is not reported. Standard errors in parenthesis are heteroskedasticity-consistent.

[b] Instrumental variable estimate.

[c] From an OLS regression of the fourth and fifth equation.

Table 16. Exchange Rate Regime, Asset Market Integration, Bilateral Trade and Output Correlations[a]
Bolded signifies significance at the 5% level and bolded-italicized significance at the 10 percent level)

| | 1880-1913 | 1920-38 | 1950-72 | 1973-2001 | 1950-72 | 1973-2001 |
|---|---|---|---|---|---|---|
| ã | 0.122 | ***-0.417*** | **0.603** | **0.624** | ***0.42*** | 0.332 |
| | ( 0.085) | ( 0.233) | ( 0.187) | ( 0.176) | ( 0.231) | ( 0.205) |
| ä | | -0.132 | 0.001 | ***-0.183*** | -0.009 | -0.1 |
| | | ( 0.167) | ( 0.114) | ( 0.103) | ( 0.105) | ( 0.114) |
| *â*[b] | | | | | ***0.109*** | **0.15** |
| | | | | | ( 0.063) | ( 0.059) |
| R-square [c] | 0.006 | 0.012 | 0.064 | 0.092 | 0.09 | 0.172 |
| R-square first-stage regressions | | | | | 0.363 | 0.364 |
| Number of observations | | 120 | 120 | 120 | 120 | 120 |

[a] Constant is not reported. Standard errors in parenthesis are heteroskedasticity-consistent.
[b] Instrumental variable estimate.
[c] From an OLS regression of the fifth and sixth equation.