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The Transmission of International Business Cycles Shocks: Evidence from the European Periphery, 1861-1913

CONCEPCIÓN GARCÍA-IGLESIAS

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European University Institute
Badia Fiesolana
I – 50016 San Domenico (FI)
Italy

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by

Concepción García-Iglesias
Marie Curie Fellow

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Concepción García-Iglesias
European University Institute
Via Boccaccio, 121
50133 Firenze
Florence – ITALY
Email: garcia-i@datacomm.iue.it

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Abstract

This study compares the international transmission of both nominal and real business cycle shocks from 1861 to 1913 in Scandinavia and the Southern European countries of Italy, Portugal, and Spain. Cointegration analysis and estimated vector autoregression for real GDP and inflation reveal the presence of a real European business cycle. Specifically, I highlight the importance of exchange rate systems in the transmission of business cycles shocks. A remarkable result is that British inflation has a contemporaneous influence on the Scandinavian rates. This finding is consistent with the view that countries on the gold standard and thus under fixed exchange rates (i.e., Scandinavia) do not absorb shocks and reflect the transmission of nominal business cycles disturbances. On the contrary, the Southern European countries, not in the gold standard, do not experience a strong impact of international disturbances. Flexible exchange rates insulate the three countries from both foreign nominal and real shocks.

I. Introduction

During the last years of the nineteenth century all the European countries adopted the gold standard, a regime that worked well with no major changes until World War I. The world economy suffered great shocks during the 1870s decade, and the gold standard was considered as the solution to mitigate those shocks. The gold standard was thought to be the "perfect" mechanism to avoid pressure over the balance of payments, to satisfy the various objectives of the national monetary authority, and to provide price and exchange rate stability. The system could work as long as the key international currency operating with gold, the pound sterling, was strong and backed by financial institutions assuring its convertibility in gold. Essentially the gold standard was a system of international balance of payments adjustment, capable of adjusting short-term deficits or surpluses in the external balance of individual countries using gold as the basis of the circulating medium.

The classical gold standard regime was a credible commitment mechanism characterized by stable economic growth, relative political stability, and the belief in free trade. Moreover, it provided access to the international capital markets of the core countries. It was a system of fixed exchange rates, and so international disturbances were transmitted easily from the main financial centers to the rest of the countries in the system. In contrast, countries that did not adhere strictly to the rule could suffer from no access to international capital markets and low growth, lacking then stability in the system. Nevertheless, the presence of flexible exchange rates allowed them to be insulated from external shocks. Indeed, countries not in the gold standard did not experience transmissions of international business cycles. Therefore, I expect that belonging to the international monetary regime reflect the degree of synchronization of national business cycles with the international one, economic development, or simply political will to foster a closer economic union. It is this transmission of business cycles shocks what this study intends to investigate, following the line set up by Craig and Fisher (1992), and therefore by Friedman and Schwartz (1982).

Traditional studies of the business cycle in late 19th century Europe claim that financial crises in one country expand to others, inducing financial and real effects over a number of countries. Money is, therefore, not neutral in each country and money supplies are all linked through the international monetary rule (Friedman and Schwartz, 1982; Huffman and Lothian, 1984). However, empirical results on modern data as well as on some available historical data for few European countries (Bordo, 1986) show that financial crises in one country spread to others affecting them just in real terms. Money is then neutral. In the spirit of Craig and Fisher (1992), this same real business cycle explanation holds and there is monetary neutrality. Moreover, there exists a close financial integration of the European economy on the late years of the nineteenth century. At the same time, discussions of the international transmission of business cycles also involve the exchange rate mechanism of the time. Theoretical work in macroeconomics argues that flexible exchange rates, as opposed to the fixed exchange rate system of the gold standard, do not prevent the transmission of international business cycle shocks (see Fleming, 1962; Mundell, 1968). Nevertheless, some more recent econometric results for historical data defend the idea that flexible exchange rates protect a country from foreign business cycle disturbances as opposed to what has been previously said (Choudhri and Kochin, 1980; Bordo, 1985).

This paper then proposes to study the international transmission of business cycle shocks during the classical gold standard period, 1861-1913, and its effects for two different groups of countries - Scandinavia and the southern European countries of Italy, Portugal, and Spain. The former strictly adhered to the gold standard and had a fixed exchange rate regime, whereas the Southern European countries did not. An international crisis might be a disturbance independent of economic domestic conditions. Therefore, by looking at these two different groups of countries at the time of the gold standard, I pay especially close attention to exchange rate systems and their ability to absorb shocks, or on the contrary, to transmit business cycles disturbances from abroad. Basically, the two objectives of this paper are: to study the dynamics of the business cycles in these different countries; and, to see whether a flexible exchange rate system insulates an economy from international shocks.

This analysis is new in what it considers a set of countries that have not historically been analyzed and compared in this sense. The study is largely made possible by the development of national income accounting in Scandinavia as well as in Portugal and Spain. Moreover, transmissions of business cycles lead to possible interactions among these countries and Great Britain and the United States, which are the two references for international shocks. This allows to analyze whether there was a European business cycle among the countries that adopted the classical gold standard and that shared the same monetary policy. From this historical data, we can learn some lessons when compared to the current debate on European business cycles disturbances (see Artis and Zhang, 1997).

To investigate the transmission of international business cycle shocks, I analyze the behavior of GDP and change in prices over time using cointegration tests and vector autoregressions (VARs). I conduct Granger causality tests and innovation accounting – impulse response analysis and variance decomposition – to study the interrelationship between GDP and changes in price level over time. A test of Granger causality looks for the causation direction of the variables considered, one way or bi-directional. On the other hand, cointegration reveals whether there is a relationship between real business cycles of the different countries implied in the study. The empirical results are then used to compare the extent to which not following the gold standard insulated small countries from international disturbances (Huffman and Lothian, 1984; Bordo, 1985). If flexible exchange rates insulate an economy, we would expect a given country's output and prices to be largely unaffected by international shocks. On the other hand, if flexible exchange rates fail to insulate an economy, then we would expect international shocks to account for a large portion of movements in domestic output and prices.

We might expect then the Southern European countries to be insulated from both nominal and real shocks due to the flexible exchange rate regime. In contrast, I expect the Scandinavian countries not to be insulated from shocks, and reflect the international transmission of the business cycles through fixed exchange rates and its tight relationship with the gold standard regime. Therefore, this paper provides additional insight into the operation of the classical gold standard in the European periphery.[1] In addition, it allows us to show how insulated the countries that did not follow the gold standard were relative to countries that strictly adhered to the international monetary rule based on the importance of exchange rate regimes.

The paper is organized as follows. First, I present some historical background on the operation of the monetary regimes followed by the countries used in this study. Then I discuss the data used in this empirical analysis. This is followed by a description of the econometrics approach and the empirical results. The results are then placed within the framework of different exchange regimes to test the insulation issue. The paper concludes by summarizing the main results of this study.

II. The Gold Standard Experience of the European Periphery

To examine the transmission of international business cycles shocks, I compare a number of small European countries that operated under different monetary regimes during the last years of the nineteenth century and the beginning of the twentieth century. This sample has been chosen on the basis of adherence to the monetary rule in addition to data availability.

Recent research in monetary history confirmed that the gold standard worked successfully for the core countries on the classical gold standard: Britain, France, and the United States. A number of other countries also followed the rule. These included Australia, Canada, Japan, the Netherlands, the Scandinavian countries, and Switzerland. The last group is the countries of Latin America and the Southern European countries for which gold convertibility was the exception rather than the rule. This paper compares the effects of shocks on GDP and prices for two groups of European countries. The first group

[1] European periphery in this context means that these countries were not the main financial centers of the time.

contains countries that followed the gold standard, the Scandinavian countries of Denmark, Norway, and Sweden. In contrast, the second group of countries, Italy, Portugal and Spain, did not strictly adhere to the international regime but had an independent monetary policy.

The gold standard experience of the three Scandinavian countries can be analyzed as a whole just by looking at the development of the Scandinavian Currency Union which was established during the first years of the 1870s. Prior to the adoption of the gold standard, Denmark, Norway, and Sweden operated under a silver standard. In 1872, an agreement to form a monetary union was signed. Denmark and Sweden joined the Union that year while Norway did it three years later, in 1875. The Scandinavian Monetary Union was based on gold and adopted a common unit, the krona, which circulated equally in the three countries. However, there was not a gold coin in circulation due to the public preferences of using notes instead of gold. The Swedish constitution guaranteed the convertibility of central bank's notes into gold. The three central banks agreed to allow each other to draw drafts on each other at par. All notes, gold coins, and token coins were accepted at par in the other countries that allowed the Scandinavian Monetary Union to be very successful. In 1905, a political conflict between Norway and Sweden concerning the abolition of their political union led to a reduced monetary cooperation among the Scandinavian countries. The central banks were still accepting each other drafts but not necessarily at par. The three Scandinavian countries showed a common pattern of adoption of the gold standard between 1872 and 1875 and adhered to the system until the outbreak of the World War I in 1914 (Bergman *et al*., 1993; Henriksen and Kærgard, 1995; Jonung, 1984).

Italy, Portugal, and Spain instead did not strictly adhered to the monetary rule. Italy adopted a bimetallic standard in 1862, although the monetary regime was a de facto gold standard. Three years later, Italy became a member of the Latin Monetary Union but war against Austria and fiscal profligacy forced Italy to abandon convertibility. Fiscal and monetary discipline was restored along with exchange rate parity in 1874. The central authorities resumed convertibility on April 12, 1884, but money was only convertible into silver because silver was overvalued. Italy adopted a fiduciary standard in 1894 and remained on this system until 1914 (Fratianni and Spinelli, 1984).

Portugal operated under a bimetallic standard since the 1680s, alternating between gold and silver. In 1854, Portugal joined the international monetary regime and was a member until 1891. The Baring international crisis in conjunction with poor government policies forced Portugal to abandon the gold standard in 1891. Portugal suffered from a deep political, economic, and financial crisis during the 1890s and remained inconvertible until after World War I (Reis, 1996, 1999).

Spain did really never adopt the gold standard for the entire period and operated under different monetary and fiscal regimes during the last half of the nineteenth and the beginning of the twentieth centuries. In 1848, Spain adopted a bimetallic standard that did not become fully operational until a monetary reform in 1868. During the 1870s, when many countries adopted the gold standard, Spain moved towards a fiduciary system with flexible exchange rates. Silver production increased and the price of silver in terms of gold fell during these years. Spain, which was then on a bimetallic system, operated under a de facto silver standard where the intrinsic value of the coin was smaller than the face value. Convertibility of paper was finally suspended in 1883 and resumption never took place before or after World War I. However, Spain did enact fiscal and monetary reforms in the early part of the twentieth century to reverse the fiscal problems that emerged as a result of the Spanish-American War.

The objective of this paper is to investigate the incidence of cyclical fluctuations across countries in the gold standard – Scandinavia – and countries not in the gold standard – Italy, Portugal, and Spain. The Friedman-Schwartz work (1982) is the point of departure of almost all studies about the interrelationship between real and monetary issues. Any possible relationship between these two issues exists only in the short-run. In the long run, money is neutral with respect to output (Capie and Wood, 1994). However, and as noted by Craig and Fisher (1992), there is a connection between money and real events for business cycles. They also show the existence of strong financial linkages between European markets in the late years of the nineteenth century. I then propose to extent their study on business cycle transmissions to a new group of European countries where the main characteristic is their different experience with the international monetary rule, the classical gold standard.

III. Data

To test the transmission of international business cycle shocks, I undertake an empirical analysis of the latest gross domestic product (GDP) estimates and wholesale price indices for the eight countries of my sample. The quality of nineteenth century GDP figures is quite variable, especially for countries in the European periphery such as Italy, Portugal, and Spain (see Bardini *et al*., 1995). GDP series for each country are normalized so that 1913=100. For Italy, I employ Fuà and Gallegati's (1996) new estimates of Italian GDP for the years 1861-1913. For Portugal and Spain I use Nunes *et al*. (1989) and Prados (1995), respectively. Recent estimates of GDP developed by Krantz (1997) are employed for Denmark and Sweden. For Norway, the absence of new series forced me to use Mitchell's (1978) which spans the period 1865-1913. For the two references, the U.K. and U.S., GDP series are drawn from Feinstein (1972, 1988) and Romer (1989), respectively. The U.S. data are for the years 1869-1913. Real GDP estimates for Scandinavia, and Italy, Portugal, and Spain with respect to the U.K. and the U.S. are plotted in Figures 1 and 2, respectively.

The so-called inflation rates (first differences in the logarithms of price levels) have been calculated from the wholesale price indices. The data of these series come from the following sources: Denmark, Norway, Sweden, and Spain (Mitchell, 1978). The Danish series covers the years 1876-1913 while the Norwegian spans the period 1891-1913. For the other countries, the references are: Italy (Fratianni and Spinelli, 1984), Portugal (Nunes *et al*., 1989), the U.K. (Capie and Weber, 1985), and the U.S. (Balke and Gordon, 1989). Figures 3 and 4 present inflation rates for both groups of countries versus Great Britain and the United States.

Therefore and by looking at these four figures, it seems that the three Scandinavian countries display a higher degree of co-movement, being more synchronized with respect to Great Britain and the United States than the Southern European countries for the two variables considered. This confirms the differences between countries under the gold standard regime with fixed exchange rates – Scandinavia – and countries that were not in the gold standard system with flexible exchange rates - the Southern European countries of Italy, Portugal, and Spain.

IV. Empirical Analysis of GDP and Inflation Time Series – Cointegration Tests

To investigate the transmission of international business cycle shocks in the European periphery, I analyze the behavior of GDP and inflation rates over time using cointegration tests and vector autoregressions VARs. I use this methodology because it allows me to measure the causality direction and the possible interaction that exists among the variables. At the same time, cointegration suggests the existence of contemporaneous deviations from trend that implies the existence of a common business cycle among the different economic powers.

A large number of procedures have been suggested for estimating and testing stationary long-run relationships between variables. Engle and Granger (1987) introduced the concept of cointegration. Afterwards, several procedures were developed to avoid the defects of the Engle-Granger estimation. Therefore, I will use the two most popular cointegration tests to examine the proper way to estimate a system of cointegrated variables – Engle-Granger and Johansen procedures.

The cointegration technique developed by Engle and Granger (1987) offers a new possibility to test the long run relationship between two (or more) nonstationary processes. Then cointegration exists if deviations from the presumed relationship have bounded variability.

The most recent studies have shown that real GDP time series are nonstationary processes but the cointegration method establishes that even in this case, a linear combination of real GDP with respect to UK and US real GDP would be stationary. On the contrary, we would expect the differenced data of the changes in price levels to be stationary, and thus any linear combination of this variable with the two reference countries considered in these study, would be stationary as well. Therefore, the first step is to show that these series are I(1) or nonstationary. That is to test whether the series contain a unit root or not. To test for unit roots, the method used will be the Dickey-Fuller – DF – (1979), augmented Dickey-Fuller – ADF – (1984), and Phillips-Perron – PP – (1988) tests.

This econometric technique requires a regression of the variables (GDP or inflation), on a constant term, α, lagged variables (GDP or inflation), lagged changes in latter variables, and a time trend, T:

$$\Delta x_t = \alpha + \theta x_{t-1} + \Sigma_i \beta \Delta x_{t-i} + \delta T + \mu_t \qquad i = 1 \ldots p \qquad (1)$$

where p is the number of lagged dependent variables. Table 1 gives the result of the unit root tests for the differenced data for the real GDP. From these results we can see that the value for the t-statistic with a constant and time trend is bigger than the critical value for the augmented Dickey-Fuller, the Dickey-Fuller, and the Phillips-Perron tests. Hence I accept the null hypothesis: real GDP has a unit root, and thus the series are non stationary for all the countries in the sample.[1] On the contrary, the results obtained for the inflation rates differenced data reveal that there is a tendency of the inflation rate to revert

[1] The same results have been obtained when running the regression in levels.

to a long run value (Table 2). I can reject the presence of a unit root at the five per cent level, and hence the inflation rate time series are stationary.[1]

The cointegration technique is designed to test for long run equilibrium relationships for which no adjustment mechanism has been specified. The cointegrating equation is then

$$y_t = \alpha + \beta x_t + \gamma z_t + \mu_t$$
(2)

where y_t, x_t and z_t are the measures of the business cycles activity in different countries. The test asks whether a group of nonstationary variables can be linearly combined to produce a stationary variable. If so, the nonstationary variables are said to be cointegrated (Engle and Granger, 1987). Therefore, I test μ_t for a unit root. If the series are cointegrated, residuals from (2) should be stationary. Then one can say that a linear relationship between the variables over the long run exits. If, by the contrary, the residuals have a unit root, the long run relationship is not stationary.

In order to check for cointegration, there is a need to verify if all series, individually, are or not stationary. The finding of unit root in a time series indicates non stationarity which implies that shocks might have a permanent effect on the economy. As this has been the case for the real GDP – I run the cointegration regression (2) to show that a linear combination of these non-stationary variables with respect to the UK and US could be stationary. The procedure is to regress a variable in which we are interested – let say the Spanish real GDP – on the variables that are considered references– British and American real GDP. The results are given in Table 3.[2]

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[1] The primary concern about this random walk test hypothesis relies on the lack of sufficient power to reject. Therefore, this test may give us little information against the relevant alternative hypothesis. A cointegration test is then required to confirm the results.

[2] As expected, the linear combination between inflation rates series and the two references is stationary. Hence, there exists a cointegration relationship (see table 4).

Hence, the cointegrating regression gives the values for the Dickey-Fuller test on residuals. In all cases, the null hypothesis that the series are not cointegrated should be rejected. The null hypothesis for no cointegration does not hold which lead us to the conclusion that there is a long run relationship. Both series, real GDP and inflation for the six countries, are cointegrated with respect to the two references, the U.K. and the U.S., and so there is a linear relationship among the variables. The series do no wander apart over time and therefore, they share a long run relationship.

The other cointegration method was developed by Johansen (1988). The estimation of the long-run equilibrium regression requires the use of one variable on the left hand-side and the others as regressors. However, it is possible to find that one way of regressing the variables indicates cointegration whereas reversing the order indicates no cointegration. This is a serious problem because we expect the results to be consistent without taking into account the choice of the variable selected for normalization. The problem is even more important when more than three variables can be selected as the left hand-side variable.

Another problem of the Engle-Granger method is that it relies on a two step procedure. First, it generates the error series and second, uses these generated errors to estimate a regression of $\Delta \hat{e}_t = a_1 \hat{e}_{t-1} + \dots$. So, the coefficient a_1 is obtained by estimating a regression using the residuals from another regression. Therefore, any error that can appear in step one is carried into step two. Johansen procedure avoids this kind of problem. It can estimate for the presence of multiple cointegrating vectors and does not use the two step estimator by using the maximum likelihood estimators. Furthermore, these tests allow the researcher to test restricted versions of the cointegrating vector(s) and speed of adjustment parameters (Enders, 1995).

Using the time series data for real GDP and inflation rates, I test the null hypothesis of no-cointegrating vectors against the alternative of cointegrating vectors. As we are simply interested in the hypothesis that the variables are not cointegrated, I test the null hypothesis of no cointegrating vectors against the alternative of one or more cointegrating vectors. Tables 5 and 6 give us the calculated values of real GDP and inflation, respectively, for both λ trace and λ max tests. These values are then compared to the corresponding critical values. Thus, at the 90% level, the restriction is binding for all six countries, so that the variables are cointegrated using these tests. In both λ trace and λ max tests, we can reject the null hypothesis and confirm the results of our previous tests. There is a cointegration relationship among the real GDP and inflation rates for all six countries with respect to Great Britain and the United States.

Therefore, the results of both cointegration techniques reveal that there is no difference between the experiences of Scandinavia and the Southern European countries. The existence of cointegration for both variables, GDP and inflation rates, confirms the presence of real business cycles among the European periphery and the economic powers of the time, Great Britain and the United States.

V. VAR Analysis and Innovation Accounting

We are interested in the dynamic relationship between real GDP and changes in price levels time series with respect to the U.K. and U.S. in order to analyze the transmission of business cycles shocks from 1861 to 1913. Therefore, I first conduct a VAR analysis that provides a Granger causality test, and second innovation accounting – impulse response analysis and variance decomposition to study the interrelationships among the variables over time.

VAR methodology allows us to examine the reactions of one variable to the other and vice versa. It postulates that all the variables in the system are endogenous and that each can be written as a linear function of its own lagged values and the lagged values of all other variables in the system. If all variables are joined in one single equation, or vector, this is called a vector autoregression. This vector represents a linear function of its own lagged values plus an error vector. However, it is inappropriate to estimate a VAR of cointegrated variables using only first differences. Therefore, I include the error-correction portion of the model. The error correction model is particularly interesting because it provides a "reconciliation" between short run and long run behavior. Hence, cointegration tests based on error correction models give real support for a long run equilibrium relationship.

In general, both variables in a cointegrated system respond to a deviation from long-run equilibrium. But, it can be possible that one of them won't. In such a case, that variable does not respond to the discrepancy from long-run equilibrium and the other variable does all the adjustment. A new reinterpretation of Granger causality is then needed in a cointegrated system. As Enders specifies, in a cointegrated system, $\{z_t\}$ does not Granger cause $\{y_t\}$ if lagged values Δz_{t-1} do not enter the Δy_t equation and if y_t does not respond to the deviation from long-run equilibrium (pp. 368-372). Hence the basic idea is that if z causes y, the changes in z should precede changes in y. The purpose is then looking for the relationship between these two variables.

Table 7 provides the results of the Granger causality test for real GDP. The F-statistic values reported refer just to one of the causality directions. I just look at the effects from Great Britain and the United States to the peripheral countries in order to analyze the influence they have over Scandinavia and the Southern European countries of Italy, Portugal, and Spain. The findings show that real GDP does not respond to deviations from long-run equilibrium and the UK and US real GDP do not Granger-cause real GDP in none of the six countries of the sample. The VAR results for inflation rates do not differ from what has been obtained for real GDP time series. Table 8 gives us the non-significant relation between inflation rates for the six countries and Great Britain and the United States inflation data. In essence and except for the cases of Sweden and Denmark (GDP and inflation series, respectively), no relevant results have been found in any of the countries, which means that there is not a causality relationship among the variables.

Once the VAR has been estimated (see Tables 9 and 10), it is important to find its dynamic structure. Innovation accounting does this through the impulse response function and variance decomposition. Impulse response function analysis determines how each endogenous variable responds over time to a shock in that variable and in every other endogenous variable. Thus the impulse response function traces the response of the endogenous variable to such shocks. In the same way, the decomposition of variance tells us the proportion of the movements in a variable due to its own shocks versus shocks to the other variable (see Statistical Appendix).

The impulse response analyses as well as the decomposition of variance of real GDP leads us to the same conclusion for all countries (see Figures 5-10). The series are exogenous to each other and thus there is not a strong correlation between real GDP and the reference countries. Therefore, the use of vector autoregressions and innovation accounting assesses that there are no dynamic responses of the real GDP estimates.

However, the innovation accounting results obtained for the inflation rates reveal a different pattern (see Figures 11-16). The British inflation essentially causes the main movements in the three Scandinavian countries inflation rates. The decomposition of variance for the Scandinavian inflation rates shows that the British inflation explains partially the forecast error variance of each one. This alternative way of finding out the dynamic structure of a VAR leads us to the conclusion that British inflation has a contemporaneous influence on Scandinavian inflation. I can then conclude that out of the six countries in the sample, Denmark, Norway, and Sweden show a clear short-run relationship between inflation and the British counterpart. Nevertheless, this is not so clear for Italy, Portugal, and Spain. Such a strong relationship between the Southern European and the British changes in price levels do not exist.

VI. Transmission of Cycles under Different Exchange Rates Regimes

Empirical evidence provided in the previous section indicates that shocks to real GDP and inflation had different short and long run effects in Scandinavia as compared with the Southern European countries of Italy, Portugal, and Spain. In both groups of countries, VAR fails to identify a dynamic relationship between real GDP and inflation with respect to Great Britain and the United States. The F-statistics, reported on tables 7 and 8 for the Granger causality tests, reveal that there is not a relevant influence from the references' variables to the peripheral variables.[1] However, there is one exception. Innovation accounting analysis leads us to the conclusion that inflation rates for the Scandinavian countries are responsive to British inflation shocks over the short-run. Thus, British inflation has a short-run influence on Scandinavian inflation. Denmark, Norway, and Sweden inflation rates show a short-run relationship with the British data, but there are no deviations from the long run equilibrium.

The cointegration tests, however, confirm the proposition that there was a real business cycle among the European periphery and the economic powers of the last years of the nineteenth and beginning of the twentieth centuries, Great Britain and the United States. The existence of business cycles relationships indicates that there must be a relationship between deviations from long-run trends in real GDP and inflation for the countries we are studying and the two economic powers of the time.

The cointegration tests of real GDP show a real interaction between these two groups of European countries and the two references I have used, mainly with the British real GDP time series. There is a linear combination among the variables that indicates the presence of real business cycles. However, real GDP series of Great Britain and the United States do not Granger cause any of the real GDP estimates of the other countries (see table 7).[2] Therefore, there is a contemporaneous evidence of real interaction but not Granger causality.

The cointegration test for the inflation time series reveal a deeper interaction of the European periphery series with respect to Great Britain and the United States. These results are also consistent with the VAR analysis for the Scandinavian countries. I cannot state that the inflation rates of neither Great Britain nor of the United States Granger-cause

[1] As I want to study the effects that Great Britain and the United States have over peripheral countries, I just look at the results of the Granger causality test in this direction. I do not analyze the possible effects of the variables the other way around.

[2] The F-tests indicate that, at conventional significance level, the U.K. and the U.S. Granger-cause themselves (i.e., the existence of cointegrating vector necessarily implies Granger causality in, at least, one direction).

the inflation rates in the European periphery.[1] However, a significant effect of British inflation on the European periphery's inflation appears in relatively few instances. This is the case for the Scandinavian countries. The results for Denmark, Norway, and Sweden suggest that the rate of inflation is not necessarily exogenous in monetary models. The cointegration tests depict the existence of a contemporaneous relationship among the Scandinavian and the British inflation. Moreover, British inflation rates influenced the Scandinavian rates over the short run, and thus there was a correlation among them.

To explain the previous results based on the transmission of business cycles under different exchange rate regimes, we need to highlight the difference between flexible and fixed exchange rate systems. Advocates of flexible exchange rate regimes have often claimed that this regime better insulates the economy from foreign nominal business cycles disturbances, whereas real shocks are likely to be transmitted among countries under both regimes (Robertson and Wickens, 1997). A country that has a flexible exchange rate is able to independently determine its own monetary policy, and therefore there would not be affected by international disturbances. Flexible exchange rates will absorb international shocks and the country would only suffer from temporary disturbances. On the contrary, fixed exchange rates will allow a much easier transmission of disturbances both nominal and real from one country to another avoiding the insulation of the domestic economy.

From the point of view of small countries, any international financial crisis could have had an important effect on the economy. Thus, by focusing on these two groups of countries and its relationship with the classical gold standard, I emphasize the importance of exchange rates regimes in the transmission of international business cycles shocks. Theory predicts that nominal shocks are least transmitted among countries under a flexible than a fixed exchange rate regime. Therefore, I expect Southern European countries to be

[1] The F-tests depict that, at conventional significance level, the U.K. Granger-causes itself and the U.S. in most of the cases. On the other hand, the U.S. Granger-causes itself.

better insulated from foreign nominal shocks due to the flexible exchange rate system than Scandinavia.

Figures 1 to 4 present a striking contrast between the two exchange rate regimes. All three gold standard countries, Denmark, Norway, and Sweden experienced same movements in real GDP and inflation as the ones experienced by Great Britain and the United States during this period. Each gold country followed the British pattern closely as well as smoothly. This could be easily explained through trade and industrial relationships they shared at the turn of the century. Flourishing demand for Scandinavian products led to a close relationship with mainly Great Britain, and hence to a deeper interrelationship among these economies. In contrast, the Southern European series experienced less synchronized movements, for the most part, during this period. The only country that shared an important political and economic relationship with Great Britain was Portugal. We could have then expected Portuguese real GDP and inflation rates to follow their British counterparts. However, the results show that it did not happen this way. Portugal's strong relationship with international markets reached an end with the international financial crisis of 1891. Trade relationships diminished considerably and Portugal abandoned the gold standard regime. Therefore, I can conclude that the post-gold period had a stronger impact on Portuguese real GDP and inflation, when compared with the gold standard period.

I have compared movements in real GDP and change in price levels in each of the six countries with the movements of their counterparts in Great Britain and the United States. The empirical findings were quite consistent with what it could be expected (Huffman and Lothian, 1984). Countries on the gold standard and with fixed exchange rates are not totally insulated from shocks. They reflect the transmission of real and nominal business cycles disturbances. On the contrary, countries out of the international monetary regime are able to absorb both real and nominal shocks due to the flexible exchange rate regime.

If fixed exchange rates did not insulate an economy from foreign disturbances, I would expect to observe the same movements in real GDP and inflation for the

Scandinavian countries as well as for Great Britain and the United States. This is true when analyzing inflation rates but a slightly different picture emerges from the output data. There was not a real transmission of shocks from the U.K. and the U.S. to the Scandinavian countries when looking at real GDP series. The effect of Great Britain and the U.S. output is negligible for the three Scandinavian countries. International real disturbances (i.e., the financial crisis of 1891 and its effects) did not account for any movements in domestic output. One possible explanation for this phenomenon would be the slightly independent monetary policy followed by Scandinavia when they were under the Scandinavian Currency Union.

To obtain further evidence on the effectiveness of exchange rate regimes, I also examine the experience of Italy, Portugal, and Spain. The first two countries started the period being part of the gold standard regime but switched to flexible exchange rates when they went out of the international system. Spain remained always at the margin of such a system. If the exchange rate system does not matter in the transmission process, we would expect the output and price behavior to be not much different from the gold countries. Looking at the results of output and inflation rates, it is clear that the Southern European countries did not experience a strong impact of international shocks. The effect of Great Britain and the United States on these three countries is not noticeable. Being out of the gold standard helped these three economies to be insulated from both real and nominal shocks and therefore, there was no effect on output and prices indices. International shocks then did not account for any movements in domestic output and inflation.

VII. Conclusion

My purpose in this paper has been to investigate the transmission of international business cycles shocks during the classical gold standard period, 1861-1913, and its effects for Scandinavia and the Southern European countries of Italy, Portugal, and Spain. I analyze the behavior of real GDP and changes in price level over time using cointegration tests and vector autoregressions VARs. Specifically, I emphasize the importance of exchange rate systems in the transmission of business cycles disturbances from abroad. The results are consistent, for the most part, with the traditional view that flexible exchange rates insulate an economy from foreign cycle disturbances whereas fixed exchange rates do not.

I have shown that there is no causality between the real outputs and inflation rates between the periphery of Europe, and Great Britain and the United States during the last years of the nineteenth and the beginning of the twentieth centuries. Nevertheless, when I study the dynamic behavior of inflation rates for the Scandinavian countries, I found that British inflation has a contemporaneous influence on the Scandinavian rates. I interpret this empirical finding as consistent with the idea of fixed exchange rates acting as an instrument for the transmission of international business cycles. I believe that the close trade relationship between Scandinavia and Great Britain can explain, in part, this short-run relation between output and inflation rates.

Moreover, the cointegration tests' results reveal the presence of a linear relationship among the variables. The series do not wander apart and they share a long-run relationship. I interpret this result as consistent with Craig and Fisher's view that there was a real business cycle among the European periphery and the economic powers of the time.

On the whole, these findings appear to explain some of the differences among these two groups of countries, Scandinavia and Southern Europe. Countries on the gold standard, and thus with a fixed exchange rate system are not totally insulated from shocks and reflect the transmission of both real and nominal business cycles disturbances. In the case of Scandinavia, nominal shocks are transmitted among all the countries while real shocks are not. On the contrary, Southern European countries are able to absorb both real and nominal shocks due to the flexible exchange rate regime. I do then conclude that flexible exchange rates do prevent the transmission of international business cycles against what it has been traditionally claimed by some work in macroeconomics.

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Table 1

Unit Root Test for Real GDP

	GDP DF	**GDP** PP
Denmark		
Differences [3]	2.325	2.098
Norway		
Differences [1]	0.374	1.553
Sweden		
Differences [2]	1.424	1.357
Italy		
Differences [1]	1.279	0.874
Portugal		
Differences [1]	1.893	- 2.879
Spain		
Differences [1]	- 3.099	- 3.099
U.K.		
Differences [4]	- 0.975	- 1.303
U.S.		
Differences [1]	0.816	0.816

Note: Critical values at five per cent for the Dickey-Fuller (DF) and Phillips-Perron (PP) tests are based on estimated OLS t-statistics. Critical values for both tests are from Hamilton (1994). Number of lags in brackets.

Table 2

Unit Root Test for Inflation

	Inflation DF	Inflation PP
Denmark		
Differences [3]	- 5.362	- 4.404
Norway		
Differences [3]	- 4. 247	- 4.279
Sweden		
Differences [0]	- 4.881	- 4.881
Italy		
Differences [1]	-6.463	- 6.693
Portugal		
Differences [0]	- 6.353	- 6.353
Spain		
Differences [0]	- 7.932	- 7.933
U.K.		
Differences [0]	- 5.833	- 5.833
U.S.		
Differences [0]	- 4.694	- 4.694

Note: Critical values at five per cent for the Dickey-Fuller (DF) and Phillips-Perron (PP) tests are based on estimated OLS t-statistics. Critical values for both tests are from Hamilton (1994). Number of lags in brackets.

Table 3

Cointegration Test for Real GDP

	T-Statistic	Std. Error
Denmark	-4.019	0.125
Norway	-3.103	0.098
Sweden	-4.083	0.123
Italy	-5.957	0.152
Portugal	-3.196	0.129
Spain	-4.186	0.112

Table 4

Cointegration Test for Inflation

	T-Statistic	Std. Error
Denmark	-6.207	0.150
Norway	-6.285	0.157
Sweden	-7.597	0.141
Italy	-8.108	0.137
Portugal	-6.927	0.142
Spain	-8.602	0.138

Note: Critical values at five per cent for the Dickey-Fuller test are based on estimated OLS t-statistics. Critical values are from Hamilton (1994). All t-statistics are significant at the one per cent level (Tables 3 and 4).

Table 5

Johansen Procedure – Real GDP

	Calculated Value	Critical Value
Denmark		
λ_{trace}	53.81	32.093
λ_{max}	34.28	19.796
Norway		
λ_{trace}	42.14	32.093
λ_{max}	23.69	19.796
Sweden		
λ_{trace}	40.68	32.093
λ_{max}	20.63	19.796
Italy		
λ_{trace}	40.97	32.093
λ_{max}	25.23	19.796
Portugal		
λ_{trace}	37.79	32.093
λ_{max}	21.07	19.796
Spain		
λ_{trace}	37.62	32.093
λ_{max}	20.92	19.796

Note: Distribution of the λ statistics with a constant in the cointegrating vector. 90% critical values.

Source: Distribution of the λ statistics in Walter Enders, 1995.

Table 6

Johansen Procedure – Inflation

	Calculated Value	Critical Value
Denmark		
λ_{trace}	85.80	32.093
λ_{max}	41.08	19.796
Norway		
λ_{trace}	59.28	32.093
λ_{max}	24.87	19.796
Sweden		
λ_{trace}	68.43	32.093
λ_{max}	28.22	19.796
Italy		
λ_{trace}	78.57	32.093
λ_{max}	41.61	19.796
Portugal		
λ_{trace}	62.28	32.093
λ_{max}	26.77	19.796
Spain		
λ_{trace}	76.95	32.093
λ_{max}	42.93	19.796

Note: Distribution of the λ statistics with a constant in the cointegrating vector. 90% critical values.

Source: Distribution of the λ statistics in Walter Enders, 1995.

Table 7

Granger Causality Test for Real GDP

	U. K.	U. S.
Denmark		
F-statistic [4]	0.4383	0.8739
Significance	(0.7798)	(0.4929)
Norway		
F-statistic [4]	1.5455	0.8901
Significance	(0.2185)	(0.4838)
Sweden		
F-statistic [4]	3.6663	2.9779
Significance	(0.0169)	(0.0377)
Italy		
F-statistic [4]	2.0636	0.2905
Significance	(0.1147)	(0.8814)
Portugal		
F-statistic [4]	1.1772	1.7233
Significance	(0.3439)	(0.1751)
Spain		
F-statistic [4]	0.5067	0.8572
Significance	(0.7312)	(0.5024)

Note: The number in brackets refers to the particular lags.

Table 8

Granger Causality Test for Inflation

	U. K.	**U. S.**
Denmark		
F-statistic [1]	9.9842	0.000
Significance	(0.0036)	(0.9955)
Norway		
F-statistic [1]	0.3150	0.0083
Significance	(0.5929)	(0.9288)
Sweden		
F-statistic [2]	1.8244	0.8094
Significance	(0.1741)	(0.4521)
Italy		
F-statistic [2]	1.8169	0.0187
Significance	(0.1754)	(0.9815)
Portugal		
F-statistic [2]	0.8282	0.4706
Significance	(0.4440)	(0.6280)
Spain		
F-statistic [2]	0.0550	2.2027
Significance	(0.9465)	(0.1234)

Note: The number in brackets refers to the particular lags.

Table 9
VAR Estimation for GDP

	Variable	Std. Error	T-Statistic	
Denmark [1]	0.0166	0.2136	0.0779	
Denmark [2]	0.0582	0.2248	0.2587	**R^2** = 0.5669
U.K. [1]	0.0526	0.1458	0.3608	**DW** = 1.9342
U.K. [2]	0.0591	0.1266	0.4672	**N** = 40
U.S. [1]	0.0196	0.0982	0.1998	
U.S. [2]	0.0308	0.0951	0.3243	
Constant	0.4239	0.4669	0.9078	
Resids [1]	-0.0302	0.1716	-0.1758	
Norway [1]	0.7057	0.2126	3.3199	
Norway [2]	0.0380	0.2745	0.1385	**R^2** = 0.6437
U.K. [1]	0.0860	0.3087	0.2785	**DW** = 2.1250
U.K. [2]	0.0463	0.2872	0.1614	**N** = 40
U.S. [1]	0.1363	0.1906	0.7150	
U.S. [2]	-0.2459	0.1767	-1.3921	
Constant	0.7505	0.7754	0.9679	
Resids [1]	-0.2284	0.1196	-1.9093	
Sweden [1]	0.0298	0.1965	0.1518	
Sweden [2]	0.4089	0.1956	2.0902	**R^2** = 0.6937
U.K. [1]	0.0910	0.1981	0.4594	**DW** = 1.7173
U.K. [2]	0.1200	0.1840	0.6523	**N** = 40
U.S. [1]	0.1571	0.1387	1.1327	
U.S. [2]	-0.2867	0.1229	-2.3321	
Constant	0.6504	0.4930	1.3193	
Resids [1]	-0.2926	0.1818	-1.6095	
Italy [1]	0.0540	0.3633	0.1487	
Italy [2]	0.2943	0.3051	0.9647	**R^2** = 0.6408
U.K. [1]	-0.4403	0.2767	-1.5909	**DW** = 2.049
U.K. [2]	0.5858	0.2737	2.1399	**N** = 40
U.S. [1]	-0.0098	0.2681	0.0365	
U.S. [2]	0.0251	0.2897	0.08676	
Constant	0.6039	0.7178	0.8412	
Resids [1]	-0.5656	0.3893	-1.4528	
Portugal [1]	0.3712	0.2021	1.8367	
Portugal [2]	-0.0661	0.2088	-0.3167	**R^2** = 0.4149
U.K. [1]	0.0439	0.2585	0.1699	**DW** = 1.9303
U.K. [2]	-0.3502	0.2238	-1.5645	**N** = 40
U.S. [1]	-0.1489	0.1430	-1.0409	
U.S. [2]	-0.0946	0.1338	-0.7066	
Constant	1.7477	0.7976	2.1912	
Resids [1]	-0.3013	0.1532	-1.9665	
Spain [1]	-0.2464	0.2669	-0.9231	
Spain [2]	-0.2773	0.2146	-1.2922	**R^2** = 0.4374
U.K. [1]	-0.2207	0.3819	-0.5779	**DW** = 1.9905
U.K. [2]	0.1691	0.3554	0.4757	**N** = 40
U.S. [1]	0.2162	0.2651	0.8156	
U.S. [2]	0.3574	0.2299	1.5541	
Constant	0.5508	1.1921	0.4621	
Resids [1]	-0.2828	0.2379	-1.1884	

Note: The number in brackets refers to the particular lags.

Table 10

VAR Estimation for Inflation

	Variable	Std. Error	T-Statistic	
Denmark [1]	0.3589	0.3045	1.1785	
U.K. [1]	-0.7766	0.2458	-3.1598	**R^2** = 0.3598
U.S. [1]	-0.0009	0.1516	-0.0056	**DW** = 2.0178
Constant	0.4637	0.7943	0.5838	**N** = 35
Resids [1]	-1.0149	0.4267	-2.3788	
Norway [1]	-0.1441	0.3843	-0.375	
U.K. [1]	-0.3228	0.5751	-0.5613	**R^2** = 0.5177
U.S. [1]	-0.0194	0.2139	-0.0909	**DW** = 1.8159
Constant	0.4159	1.2486	0.3331	**N** = 20
Resids [1]	-1.2602	0.5543	-2.2734	
Sweden [1]	0.5094	0.3873	1.3154	
Sweden [2]	0.2558	0.2845	0.8994	**R^2** = 0.2176
U.K. [1]	-0.6944	0.3675	-1.8897	**DW** = 2.0228
U.K. [2]	-0.2574	0.2698	-0.9539	**N** = 49
U.S. [1]	-0.0046	0.0962	-0.0479	
U.S. [2]	-0.1154	0.0930	-1.2406	
Constant	-0.0406	0.6689	-0.0607	
Resids [1]	-0.8581	0.4958	-1.7306	
Italy [1]	0.5473	0.2858	1.9150	
Italy [2]	0.1041	0.1947	0.5348	**R^2** = 0.5200
U.K. [1]	-0.5667	0.3131	-1.8101	**DW** = 1.9887
U.K. [2]	-0.1091	0.2470	-0.4416	**N** = 49
U.S. [1]	-0.0201	0.1188	0.1691	
U.S. [2]	0.0138	0.1071	0.1288	
Constant	0.2764	0.7680	0.3599	
Resids [1]	-1.6813	0.3982	-4.2219	
Portugal [1]	0.1454	0.2085	0.6975	
Portugal [2]	0.0144	0.1498	0.0961	**R^2** = 0.5360
U.K. [1]	0.1924	0.1668	1.1534	**DW** = 1.9893
U.K. [2]	-0.0040	0.1674	-0.0240	**N** = 49
U.S. [1]	-0.0747	0.0782	-0.9557	
U.S. [2]	-0.0060	0.0776	-0.0775	
Constant	-0.0718	0.5525	-0.1299	
Resids [1]	-1.0587	0.2633	-4.0205	
Spain [1]	0.0049	0.2493	-0.0196	
Spain [2]	-0.2135	0.1566	-1.3634	**R^2** = 0.6596
U.K. [1]	-0.0800	0.2437	-0.3282	**DW** = 1.9309
U.K. [2]	-0.0434	0.2442	-0.1778	**N** = 49
U.S. [1]	-0.1699	0.1159	-1.4661	
U.S. [2]	0.1312	0.1189	1.1039	
Constant	-0.3806	0.8177	-0.4654	
Resids [1]	-1.1292	0.3389	-3.3323	

Note: The number in brackets refers to the particular lags.

Statistical Appendix

A first-order autoregresion VAR (the longest lag length is unity) in a two variables case, can be written as follows (see Enders, 1995):

$$y_t = b_{10} - b_{12}\, z_t + \psi_{11}\, y_{t-1} + \psi_{12}\, z_{t-1} + \varepsilon_{yt} \qquad (3)$$

$$z_t = b_{20} - b_{21}\, y_t + \psi_{21}\, y_{t-1} + \psi_{22}\, z_{t-1} + \varepsilon_{zt} \qquad (4)$$

where it is assumed (i) that both y_t and z_t are stationary; (ii) ε_{yt} and ε_{zt} are white-noise disturbances with standard deviations of σ_y and σ_z, respectively; and (iii) $\{\varepsilon_{yt}\}$ and $\{\varepsilon_{zt}\}$ are uncorrelated white-noise disturbances.

Equations (3) and (4) are not reduced-form equations since y_t has a contemporaneous effect on z_t and z_t has a contemporaneous effect on $y_{t.}$ However, it is possible to transform this system of equations in a more useful form. A matrix transformation of these two equations leads to the vector autoregressive (VAR) model in standard form. Therefore an equivalent form of writing this vector is:

$$y_t = a_{10} + a_{11}\, y_{t-1} + a_{12}\, z_{t-1} + \varepsilon_{1t} \qquad (5)$$

$$z_t = a_{20} + a_{21}\, y_{t-1} + a_{22}\, z_{t-1} + \varepsilon_{2t} \qquad (6)$$

The first system of equations is called VAR or primitive system while the second is called a VAR in standard form.[1]

A good way to examine the relationship between cointegration and error correction is to study the properties of the simple VAR model in its standard form, equations (5) and (6). This model can be estimated by OLS. Since there are no unlagged endogenous variables on the right-hand side, and since the right-hand side variables are the same in

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[1] For more details, see Enders, 1995, pp. 294-296.

each equation, OLS is a consistent and efficient estimator. Although the errors are correlated across equations, estimating using seemingly unrelated regressions does not add to the efficiency of the estimation procedure since both regressions have identical right-hand side variables (Pindyck and Rubinfeld, p. 354).

Moreover, we can add more information into the interpretation between two given time series by performing causality tests. If $\{y_t\}$ does not improve the forecasting performance of $\{z_t\}$, then $\{y_t\}$ does not Granger-cause $\{z_t\}$. We can say then that a variable y is said to Granger-cause z, if prediction of the current value of z is corrected by using past values of y. This definition is implemented for empirical testing by regressing z on past, current, and future values of y; if causality runs one way, from y to z, the set of coefficients of the future values of z should test insignificantly different from the zero vector – we use for this purpose an F-test – and the set of coefficients of the past values of y should test significantly different from zero. Any autocorrelation in errors should be eliminated prior to run the regression (Kennedy, p. 68).

As in a traditional VAR, innovation accounting can be used to obtain information about the interactions among the variables. The impulse response function shows how shocks to any one variable filter through the model to affect every other variable, and eventually feed back to the original variable itself. The variance decomposition analysis breaks down the variance of the forecast error for each variable into components that can be attributed to each of the endogenous variables (Pindyck and Rubinfeld, pp. 385-389). In both analyses, results can be sensitive to the order of the variables.

Impulse response analysis can quantify and graphically depict the time path of the effects of one variable on the other. This different approach determines how each endogenous variable responds over time to a shock in that variable and in every other endogenous variable. Thus the impulse response function traces the response of the endogenous variables to such shocks. Ideally, we would like to identify shocks with specific endogenous variables, so that we can determine how an unexpected change in one variable affects all variables over time.

Another way of characterizing the dynamic behavior of the model is through variance decomposition. Looking at the system of two equations previously defined, the forecast error variance decomposition tells us the proportion of the movements in a sequence due to its own shocks versus shocks to the other variable. If ε_{zt} shocks cannot explain the forecast error variance of $\{y_t\}$ at all forecast horizons, we can say that the $\{y_t\}$ sequence is exogenous. In such a case, the $\{y_t\}$ sequence is independent of the ε_{zt} shocks and of the other sequence, $\{z_t\}$. In the contrary case, if ε_{zt} shocks could explain all of the forecast error variance in the $\{y_t\}$ sequence at all forecast horizons, then the $\{y_t\}$ sequence would be totally endogenous. Usually, it is normal for a variable to explain almost all of its forecast error variance at short horizons and smaller proportions at longer horizons.