Trade and Business-Cycle Comovement: Evidence from the EU

Luigi Bocola*
Università degli Studi di Torino

This paper is an empirical study of the determinants of business-cycle comovement. Using a panel of European countries (1972-2004) it is found that bilateral trade intensity is a robust determinant of real comovement in Europe, this confirming the seminal study by Frankel and Rose (1998). It is also found that convergence in macroeconomic policies (especially fiscal policies) is associated to high degree of intra-european business-cycle correlation. Moreover, having controlled for policy convergence, the effect of bilateral trade on business cycle comovement weakens on average by a factor of 36%-33% with respect to that estimated according to Frankel and Rose's econometric specification, this suggesting the potential endogeneity of the set of instrumental variables adopted by the two authors (Gruben, Koo and Millis, 2002). [JEL Classification: F15, F33, E32]

1. - Introduction

One of the main results of the optimum currency area (OCA) theory is that countries whose frequency of idiosyncratic shocks is high are less suitable to take part in a fixed exchange rate regime[1]. The reason lies in the idea that joining a currency area

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[1] This idea was first put forward by MUNDELL R. (1961) and it was further explored by KENEN P. (1969). The last 20 years have seen a growing body of articles in the area of OCA theory, see TAVLAS G. (1994) for a survey. On the empirical

limits the ability of countries to use national monetary policy to respond to country-specific shocks. This finding has been successively extended into the environment of a monetary union[2] and it has been one of the main analytical tools adopted by economists to gauge the economic suitability of the European monetary unification process. Relying on historical patterns of real comovement, some authors[3] have argued that the adoption of the single currency would create macroeconomic stability problems for the euro zone, especially because of the low degree of intra-European labor mobility and because of the absence of a federal risk sharing system.

This view has been criticized by Frankel and Rose (1998). Applying a Lucas critique to the synchronicity criteria of OCA theory the two authors have argued that a fixed exchange rate regime may dramatically change the historical record of real comovement: indeed, the boost in trade intensity between countries participating in the currency area may cause their cycles to be more and more similar. In other words, a currency area may be *self-validating*[4], so that *ex-ante* valuations of its optimality would be redundant. As an empirical support of their idea, the two scholars have estimated a positive and wide effect of bilateral trade intensity on the correlation of cycles between 21 OECD countries.

Frankel and Rose's study have stimulated a growing body of empirical literature that has further investigated the determinants of business-cycle comovement. Perturbation of the basic model (see in particular Baxter and Kouparitsas, 2005; Imbs, 2004) did not question the positive "trade effect" although its magnitude has been

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implementation of the criteria of optimum currency area theory see BAYOUMI T. - EICHENGREEN B. (1997). On the theoretical side, an excellent formalization is provided by ALESINA A. - BARRO R.J. (2002).

[2] See DE GRAUWE P. (2000) for a two-country model that studies the efficiency of the common monetary policy in presence of asymmetries both in the occurrence of shocks and in the transmission mechanism.

[3] On the historical records of symmetry of shocks between European countries see BAYOUMI T. - EICHENGREEN B. (1993). See OBSTFELD M. - PERI G. (1998) for a discussion of the links between inter-regional labor mobility, asymmetric shocks and risk sharing in the European Monetary Union framework.

[4] This terminology is due to CORSETTI G. - PESANTI P. (2002). In this work, the two authors have furnished a rationale for a Lucas critique to OCA criteria that differs from that of Frankel and Rose.

partially revisited (Imbs, 2003; Kose, Prasad and Terrones, 2003). Gruben, Koo and Millis (2002) have argued that omitted variables bias and endogeneity of the instruments adopted by Frankel and Rose may produce an overestimation of around 50%. Of course, a strong downward revision of the effect of bilateral trade on real comovement may cast some doubts on the economic relevance of this Lucas critique to the "synchronicity" criteria of OCA theory.

This paper extends the econometric specification of Frankel and Rose by considering the role of fiscal and monetary policy convergence on determining the international co-fluctuation of cycles. Using a panel of 14 European countries observed between 1972 and 2004 we find that countries having similar fiscal policies and similar real rates of return are likely to have more synchronized business cycles. Moreover, it is found that the impact of bilateral trade on real comovement, although positive and highly statistically significant, is lower (between 36% and 33%) than that estimated according to Frankel and Rose's procedure.

The paper is organized as follows: Sections 2 and 3 deal with Frankel and Rose's endogeneity hypothesis and with a discussion of their estimation strategy. Section 4 presents the results of our estimation and some sensitivity analysis. Section 5 concludes, discussing the main policy implication for the European Monetary Union.

2. - Frankel and Rose's Endogeneity Hypothesis

A symmetric distribution of shocks has been identified as a crucial prerequisite for countries to form an optimal currency area. Frankel and Rose's idea (1998) consists of considering this criterion endogenous to the creation of the currency area itself, making its *ex-ante* compliance less relevant. Their idea is based on two main conjectures:

— Fixed exchange rate should promote trade linkages between countries sharing the agreement;

— A greater extent of trade interdependence should result in more synchronized business cycles.

Recent empirical studies based on the gravity model of international trade have pointed out that the effect of monetary unification on trade is positive and statistically significant, although its magnitude tends to be linked to the econometric procedure adopted[5]. Nevertheless, as far as the direction of the effect is concerned, there is little empirical ambiguity that a common currency increases trade linkages between the countries that adopt it.

The relation between trade intensity and business-cycle synchronization seems to be more controversial from the point of view of economic theory. To formally present the channels through which trade may affect the comovement of real variables we follow Frankel ad Rose (1998) assuming that the real growth rate of output of a country may be expressed as:

$$\Delta y_t = \sum_{i=1}^{n} a_i * u_{it} + v_t + g \tag{1}$$

Where $\{u_i\}$ is the deviation of sector i's growth rate of output from the average growth rate of output v at time t, a_i is the share of sector i over total output and g is the trend growth rate of GDP. All variables are in real terms. By definition, shocks to $\{u_i\}$ are orthogonal to the average growth rate of output v_t. For convenience, it is assumed that that $\{u_{it}\}$ is distributed independently across both sectors[6] and time and that $\{v_t\}$ is distributed independently over time.

[5] In particular, cross-sectional studies tend to estimate a "currency union effect" more sizable than time-series studies do, see GLICK R. - ROSE A. (2002): given the underlying question ("how much does trade increase when two countries adopt a common currency?") the time series approach is more suitable. Another econometric problem that can heavily affect estimation of the trade effect is that of reverse causality, as PERSSON T. (2001) highlights. A time series approach which deals explicitly with the problem of reverse causality is MICCO A. - ORDONEZ G. - STEIN E. (2003), who have estimated a positive impact of the euro on intra-EMU trade between 4% and 23%.

[6] In reality, sectors are interdependent so that this assumption appears restrictive. However, this assumption is not strictly necessary for the empirical specification of this study: in fact, the regression analysis is based on the estimation of a reduced form of equation *(2)*.

Denoting by $\tilde{y}_t$ the de-trended growth rate of output at time t of a country we have:

$$COV(\tilde{y}, \tilde{y}^*) = \sum_{i=1}^{n} a_i a_i{}^* \sigma_i^2 + \sigma_{v,v^*} \qquad (2)$$

Equation *(2)* tells us that the comovement of business cycles between two countries, "Home" and "Foreign" (denoted by an asterisk) depends on the specialization pattern of the two economies, the variability of sector *i*'s cycle, and the covariance between the two countries' aggregate shocks. Imagining the two polar cases, equation *(2)* collapses to[7]

$$COV(\tilde{y}, \tilde{y}^*) = \sigma_{v,v^*} \qquad (3)$$

when two countries are fully specialized in the production of two different goods. In the case of two highly diversified economies the equation *(2)* becomes[8] instead:

$$COV(\tilde{y}, \tilde{y}^*) = c^2 n \sigma^2 + \sigma_{v,v} \qquad (4)$$

c being the share that each sector has out of the total output and *n* the number of industries in the two economies.

As can be inferred from equations *(3)* and *(4)*, two countries with a positive relation between $\{a_i\}$ and $\{a^*_i\}$ (similar sectoral composition of output) tend to have more synchronous business cycles than those with large differences between $\{a_i\}$ and $\{a^*_i\}$[9] (polarized production structure). Moreover, business-cycle correlation is likely to be positively related to the extent of country-specific aggregate shocks.

[7] The case of two highly specialized economies amounts to assume $a_i = 1$ $a{-}i = 0$ $a_j^* = 1$ $a{-}j^* = 0$ for $i \neq j$.

[8] In the case of two highly diversified economies we are assuming that $a_i = a_j^* = c$ $\forall i,j$. For simplicity we also assume that $\sigma^2_i = \sigma^2$ $\forall i$.

[9] When $\{u_i\}$ are not independent of one another, this may no longer be true. In fact, it could be in principle possible that countries fully specialized in the production of two different goods (negative relation between ai and a^*i) may have a large degree of business-cycle comovement if there is positive dependence between sector specific shocks.

There are different channels through which bilateral trade intensity may affect the extent of business-cycle synchronization.

A first channel concerns the cyclical behaviour of the trade balance. When an economy is affected by a negative aggregate shock it will tend to import less and to export more with its trading partners: the economic importance of such spill-over effects depends on the extent of bilateral trade[10]. Hence, there is little ambiguity that trade intensity positively affects the covariance of country specific aggregate shocks σ_{v,v^*}.

A second channel concerns the effect that trade has on a country's sectoral composition of output. On the one hand, reduction of transaction costs may induce countries to specialize in sectors where they have comparative advantages: in this case, international trade should induce a divergence of $\{a_i\}$ between partners. As a result, the extent of similarities in production structure between countries could in principle be reduced by the adoption of the single currency, thus leading to more asynchronous business cycles; this idea has been supported by Krugman (1993) for the case of the European Monetary Unification process. On the other hand, the adoption of the common currency may enhance the similarity of $\{a_i\}$ between trading partners whenever the boost in trade takes place more *within* industries rather than between them.

Therefore, bilateral trade intensity has ambiguous effects on the extent of comovement of business cycles, given that its overall outcome depends on the relative strength of the mechanisms highlighted above. Frankel and Rose (1998) have tackled empirically the theoretically ambiguous effect of trade over business-cycle comovement. The two scholars, using a panel of 21 OECD countries observed between 1959-1993, have estimated the following equation:

$$\text{Corr}\,(\tilde{y}, \tilde{y}^*) = \alpha + \beta\ \text{TRADE}_t + \varepsilon_t \qquad (5)$$

[10] In particular PRASAD E. - GABLE J. (1997) have pointed out the importance of "export led" recoveries in the group of industrialized countries, the magnitude of this effect being proportional to the degree of openness of the economies.

Where Corr $(\tilde{y}, \tilde{y}^*)_t$ is the correlation coefficient between a measure of de-trended real activity[11] for a given country pair (e.g. France-Germany) in sub-period[12] t and TRADE_t measures the intensity of trade linkages for the same country pair in the same sub-period. Frankel and Rose's hypothesis of endogeneity should be read as a hypothesis on the coefficient β. A positive β implies that the boost in bilateral trade amplifies the extent of comovement of real variables between countries taking part in the monetary union. The size of the coefficient identifies the economic relevance of this phenomenon: a *large* value of β implies that real comovement sharply increases after the adoption of the single currency so that Frankel and Rose's hypothesis would be confirmed. Instrumental variable (IV) estimation of equation *(5)* produces a positive and sizable effect of bilateral trade intensity over business-cycle synchronization. This "trade effect" appears robust with regard to the macroeconomic series used to measure cross-country real comovement and to the methods adopted to isolate the trend component of GDP. In one of the benchmark estimations of the model, an increase in trade intensity by 1 standard deviation increases average real comovement from 0.17 to 0.2775.

3. - A Note on Frankel and Rose's Hypothesis

The positive and large value of the coefficient β estimated by Frankel and Rose appears to confirm the endogeneity hypothesis on the synchronicity criteria of OCA theory. Countries joining a currency union are likely to experience an increase in trade *vis-à-vis* their partners and thus a (strong) increase in the degree of comovement of their business cycles: as a result, the common

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[11] The authors use four different measures of real economic activity: *i)* real GDP; *ii)* total Employment; *iii)* unemployment rate; *iv)* index of industrial production. Moreover, the authors adopt four different de-trending methodologies: *a)* fourth difference; *b)* H-P filter; *c)* linear and quadratic de-trending; *d)* H-P filtered residuals from a regression of the measure of real economic activity on a constant and quarterly dummies.

[12] The authors split their sample into four equally sized parts.

monetary policy will prove to be more efficient over time. This finding is not only important *per se*, but it brings about a number of implications regarding the economic policy institutional design of a currency area. As far as EMU is concerned, for instance, this finding indicates that fiscal policy rules may not be as costly as static OCA criteria based analysis may suggest. It is therefore crucial to settle the robustness of Frankel and Rose's findings. This section discusses the main econometric problems that may arise in the estimation of equation *(5)*.

3.1 *Inconsistency of OLS*

The first doubt regards the inconsistency of Ordinary Least Square (OLS) estimation of equation *(5)*, a point that has been made by the two authors. There are two broad explanations for that. First, measurement error in the independent variable would enable the OLS estimation of the coefficient β to be biased towards 0. Secondly, bilateral trade intensity can be considered as *endogenous* to equation *(5)*. As noted by the authors, the poor fit of the regression would suggest a "true" model of the kind:

$$(6) \qquad \text{Corr}\ (\tilde{y}, \tilde{y}^*)_t = \alpha + \beta\ \text{TRADE}_t + I_t'\gamma + \varepsilon_t$$

Where I_t is a set of relevant variables that have been excluded from the analysis. A first reason for endogeneity arises because "TRADE" can act as a proxy for variables belonging to the set I_t. A potential candidate for being excluded from the regression and being correlated with "TRADE" is monetary policy coordination. *On the one hand*, in fact, both theoretical and empirical (an exception is Clark and Van Wincoop, 2001) analyses have shown that coordination of monetary policy may have a positive impact on business-cycle comovement[13]. *On the*

[13] Checchi D. (1989) has supported the idea that business-cycle comovement between G-7 countries has increased since the mid 70's because of the process of convergence in real rates of return, due both to market development and to more

other hand, it is reasonable to expect a positive relation between trade intensity and monetary policy coordination, especially in Europe where incentives to coordinate macroeconomic policies were more intense within the European Monetary System (EMS). In fact, as OCA theory suggests, the benefits to participate in a fixed exchange rate regime are positively linked to the extent of bilateral trade so that we should expect participation in the EMS (and consequently coordination of monetary policies) being a positive function of bilateral trade linkages. Moreover, the EMS itself should have contributed to enhance trade linkages between countries that are part of the agreement. As a result, in the sample period considered, one should expect convergence in monetary policy to occur more intensively between countries with strong trade relations: as section 4.2 shows, such a claim is confirmed in the data. Hence, the exclusion of a measure of monetary policy coordination from equation *(5)* would bias upward β. A second concern for endogeneity has been made by Imbs (2003) who has noted that countries with asynchronous cycles are likely to trade more than countries with a greater extent of real comovement. Simultaneity would thus underestimate the "trade effect".

Table 1 summarizes the main sources of inconsistency of OLS and the "likely" direction of the bias of β when estimating equation *(5)* with OLS. In view of these concerns, Frankel and Rose proposed IV estimation of equation *(5)* relying on a set of instruments borrowed from gravity models of international trade: geographic and cultural proximity[14].

similar monetary policies; a second reason, built on a "discipline argument", has been sustained by ARTIS M.J. - ZHANG W. (1995), who have suggested that monetary policy may itself be a source of shocks. Participation in a fixed exchange rate regime, producing more coordination, should reduce idiosyncratic behaviour and should enhance real comovement. As an evidence, the two authors reported the emergence of a "European business cycle" during the period of the European Monetary System (EMS).

[14] Geographic proximity is measured by a dummy variable that takes the value of 1 if two countries share the same border and by the distance (in miles) between the biggest cities between country pairs, while cultural proximity by a "common language" dummy variable. See the Appendix.

TABLE 1

SOURCES OF INCONSISTENCY OF OLS ESTIMATION OF β IN EQ. *(5)*

Source of inconsistency	Direction of the bias
Measurement errors in the independent variable "TRADE"	Underestimation of β
Endogeneity of "TRADE"	Underestimation of β (Simultaneity) Overestimation of β (Omitted variable bias)

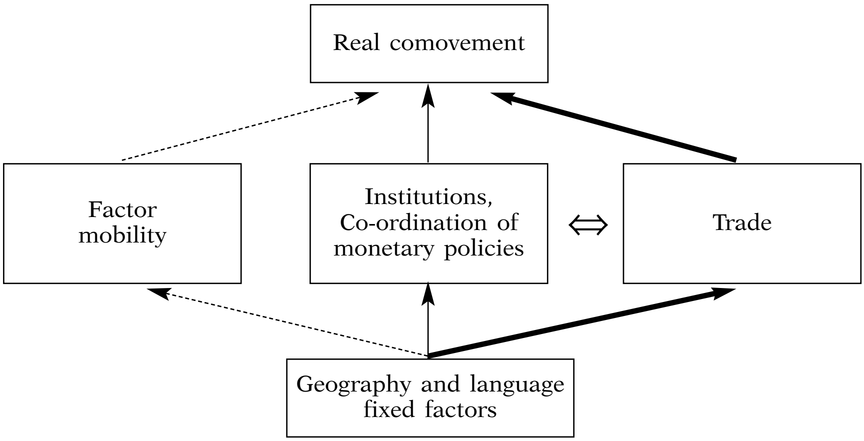
3.2 *Are Geography and Language Good Instruments?*

The issue raised here concerns the validity of the set of instruments adopted in Frankel and Rose's estimation of *(5)*. Although there is no doubt that bilateral trade is highly correlated with geographic and cultural factors, it is reasonable as well to assume that these factors affect the international correlation of business cycles not only through the trade channel. Gruben, Koo and Millis (2002), whose argument is summarized in Graph 1, were the first to make this point. The idea is as follows: geography and language proximity may affect the synchronization of business cycles also through other institutional and economic channels. Countries sharing the same border and the same culture (language) are likely to have a greater extent of factor mobility, more similar institutions and more coordinated policies, all things that, *ceteris paribus*, should have a positive influence on business-cycle comovement. If such claims were confirmed in the data, it would cast doubts on the exogeneity assumption of the set of instruments proposed by Frankel and Rose.

Gruben, Koo and Millis did not test this hypothesis explicitly. Rather, they estimated a variant of equation *(6)* by OLS, including in I_t the set of instruments adopted by Frankel and Rose. Their idea was to use geography and language to proxy for factor mobility and monetary policy coordination. The results of their estimation

Graph 1

THE GRUBEN, KOO AND MILLIS HYPOTHESIS"



Source: Gruben W - Koo J. - Millis E. (2003, page 21). A thick line indicates Frankel and Rose's hypothesis.

suggest a much lower "trade effect" (circa 50% smaller). In part, this finding is consistent with the recent literature that has further investigated the relation between trade integration and the international comovement of business cycles, (e.g. Imbs, 2003, 2004; Kose, Prasad and Terrones, 2003). Nonetheless, their econometric procedure may be questioned on two grounds: OLS estimation of equation *(6)* does not consider measurement errors and simultaneity that, as argued in the previous section, works to bias β towards 0; moreover, the use of proxies does not make it possible to distinguish among the determinants of real comovement that have been made explicit in Graph 1.

4. - Empirical Analysis

This section deals with the relation between trade intensity and business-cycle comovement for a group of European countries. The aim is to provide an estimation of the "trade effect" which is robust with regard to the econometric problems highlighted in the previous section. We extend the work of Frankel and Rose and Gruben, Koo and Millis by explicitly considering: *i)* the process of monetary and financial convergence that has taken place in Europe from the mid 70's onwards; *ii)* the role of fiscal policy coordination. Fiscal policy co-ordination, as well as monetary policy, may indeed have positive effects on business-cycle comovement: first, fiscal policy may be itself a source of shock for an economy so that its coordination, preventing idiosyncratic behaviour, should increase real comovement; secondly, coordination of fiscal policies may increase business-cycle correlation between countries when the geographic distribution of shocks is symmetric.

The benchmark framework is that of equation *(6)*. The first subsection describes the data set used. Then the focus is on the descriptive analysis of the main variable involved. Finally, estimation of the parameters and a sensitivity analysis are performed.

4.1 *Description of the Data Set*

The analysis deals with countries belonging to the EU-14[15] observed between 1972 and 2004, the time span being split into five sub-samples[16]. The choice of dividing the set of observations into sub-samples rather than relying on a cross-sectional specification is dictated by two main concerns:

— A panel structure makes it possible to increase the sample size and enhance the precision by which parameters are estimated.

— A panel structure allows controlling for sub-sample stability of the parameters, this appearing important given the large time span considered in the analysis.

As far as the splitting rule is concerned, it has been decided to consider sub-samples made up of seven years[17]. Such criterion

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[15] The countries are: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain, Sweden, UK.

[16] The sub-periods are: 1) 1972-1978; 2) 1979-1985; 3) 1986-1992; 4) 1993-1999; 5) 2000-2004.

has the main advantage of linking each sub-period to a particular phase in the construction of the European Monetary Union[18], enhancing the ability to interpret the time series behaviour of the variables. Section 4.4 checks for the robustness of our results with respect to this arbitrary division in the set of observations.

Real comovement is measured by the cross-country correlation coefficient of cyclical GDP growth rate. For each country pair (e.g. France-Germany) one measure of real comovement is observed in each sub-period, so that the maximum sample size is made up of 455 observations. Business cycle is measured as the difference between actual and trend GDP growth rates. The trend component of real GDP growth rate is isolated using two alternative filters: *i)* Baxter and King (1999) band pass filter (CICL1); *ii)* Hodrik and Prescott filter (CICL2). Although the Baxter and King band-pass filter has become a standard in this literature, the adoption of the Hodrick and Prescott's filter makes the analysis more comparable with that of Frankel and Rose. Moreover, an alternative measurement of the business cycle may furnish some robustness checks to the results.

Frankel and Rose's measure of bilateral trade intensity has become a benchmark in literature. Denoting by TRADE_{ijt} bilateral trade intensity between country i and country j in sub-period t we have:

$$\text{TRADE}_{ijt} = \frac{\left(X_{ijt} + M_{ijt}\right)}{\left(X_{it} + M_{it} + X_{jt} + M_{jt}\right)}$$

[17] Except the last sub-sample, which is made up of five years due to lack of data.

[18] See TSOUKALIS L. (1997) for this classification. The sub-samples approximately coincide with the following phases of the process of monetary unification: 1) 1972-1978: failure of the first attempt to construct a currency area in Europe (Werner Report) and period of high exchange rate's instability in Europe; 2) 1979-1985: First and second phase of the European Monetary System (EMS); 3) 1986-1992: Third phase of the EMS described by high stability of exchange rate and by the enlargement of the agreement to new Mediterranean members of the European Community; 4) 1993-1999: Crises of the European Monetary System and Maastricht convergence policies; 5) 2000-2004 Birth of the European Monetary Union.

X_{ijt} (M_{ijt}) being the total exportation (importation) from country j to country i and vice-versa. This indicator measures the relevance of bilateral trade between country i and country j over their total trading activities in sub-period t, so that it is reasonable to expect TRADE_{ijt} to be large whenever the two countries have strong trade linkages. However, it may prove to be inadequate in situations where the countries considered differ remarkably as regard to their relative size. In particular, whenever country i and country j are characterized by dissimilar levels of economic activity, TRADE_{ijt} may register low numerical values[19] even in situations where the small economy highly depends on the large one. Section 4.4 explores the effects of this issue in more detail.

As it has been discussed above, fiscal and monetary policy convergence should positively influence real comovement. At this point the problem relates to their quantitative measurement. As far as monetary coordination is concerned, the following distance indicator is proposed:

$$\text{INT}_{ijt} = \sum_k |r_{ik} - r_{jk}| / k$$

r being the short term real rate of return and k being the number of years in each sub-period. This indicator captures both the tendency towards financial integration and monetary policy convergence. The greater this indicator, the greater is the spread of short term real rate of return between country i and country j in sub-period t.

The measurement of fiscal policy coordination poses more problems. Studies that sought to gauge the impact of fiscal policy convergence on real comovement (e.g. Clark and Van Wincoop; 2001, Bergman, 2004; Darvas, Rose and Szapàry, 2005) have all used various indicators based on the fiscal balance. Incidentally, there are very good reasons to think that those measures may induce reverse causality[20]. Policy makers, in fact, partly target real variables when deciding policies, and the output gap is without doubt one

[19] This appear to be the case whenever the small economy highly depends on its trade relations with the large country and the large economy is an open and well diversified (as far as trade partners are concerned) economy.

[20] DARVAS Z. - ROSE A.K. - SZAPÀRY G. (2005) have made this concern explicit by using instrumental variables to correct the reverse causality problem.

of the variables they look at. As it is explained below, this may render problematic the determination of the direction of causality between business-cycle synchronization and policy convergence. Let us consider two countries (e.g. France and Germany) facing a positive phase of their cycle. If policy makers adopt countercyclical fiscal policies, an improvement in the fiscal balance should be observed in both countries. As a result, for the country pair considered, the econometrician registers a positive association between comovement of business cycles and convergence in fiscal positions although this correlation would not reflect causality of the latter event. Moreover, such a problem appears even worse when considering the role of automatic stabilizers: even when the discretionary component of fiscal policy is not sensitive to the output gap, a convergence of fiscal balances is going to be observed whenever the two countries are facing similar phase of the cycle. The identification problem discussed above may be attenuated by two considerations. First, existing empirical works on European countries point out the low sensitivity of the discretionary component of fiscal policy to the output gap, especially in the first part of the sample period considered (e.g. Gali and Perotti, 2003). Moreover, indicators of fiscal policy convergence can be constructed using the cyclically adjusted[21] fiscal balance, this contributing to reduce the effects of reverse casuality. Consistently, we propose three different measures of fiscal policy co-ordination[22]:

$$\text{FISC}(z)ijt = \textstyle\sum_k |D_{izk} - D_{jzk}| \,/\, k$$

where D_{izk} is:

— *fiscal balance*, defined as the difference between government revenues and expenditures (FISC1_{ijt});

— *cyclically adjusted balance*, defined as the fiscal balance net of its component attributable to the effect of the cycle (FISC2_{ijt});

[21] In order to compute the cyclical component of net lending we adopt the procedure in GIORNO C. *et* AL. (1995): the elasticity of the fiscal balance to the output gap is multiplied by the output gap. Elasticities are from VAN DER NOORD P. (2000) (sub-periods 1-4) and from GIRUARD N. - ANDRÈ C. (2005) (sub-period: 5). The output gap is the difference between actual and trend (H-P filtered) GDP growth rate.

[22] All variables are intended to be in percentage of GDP.

— *fiscal impulse*, defined as the yearly change in the cyclically adjusted primary[23] balance ($FISC3_{ijt}$)[24].

A detailed description of data sources is in Table 9 in the Appendix.

4.2 *Descriptive Analysis*

Summary statistics concerning the variables involved in the analysis are presented in the Appendix. The first point worth mentioning concerns the evidence of an upward trend in business-cycle comovement. Table 10 shows a moderate evidence for an increase in intra-European synchronization of cycles, although this path seems to depend on the variable we adopt: the CICL2 median tends to increase monotonically over time, while the same does not happen for CICL1. To investigate this aspect further, Graph 2 presents kernel density estimates (for CICL1) in three sub-periods (1972-1978; 1986-1992; 2000-2004). The graph displays evidence of bi-modal distribution, consistent with a core-periphery pattern already highlighted in literature (Bayoumi and Eichengreen, 1993; De Cecco and Perri, 1996). The evolution of the distribution over time also displays an interesting course: while the two modes seem to converge during the 80s there is a tendency for polarization in the late 90s, although the high comovement mode displays a much greater fraction of the observations.

As far as the other variables involved are concerned, their behaviour over time is somewhat expected: integration of financial markets and coordination of monetary policy starting from the second oil price shock has led to an increase in the similarity of short term real rates of return between European countries, although this tendency has stopped in the last sub-period[25]; fiscal

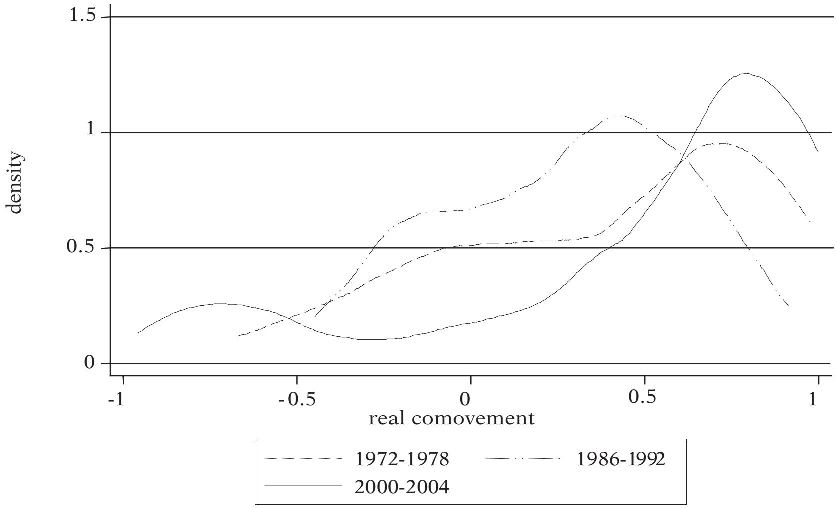
[23] The primary balance is the fiscal balance net of public debt's interest expenditure.

[24] For a description of this indicator see BLANCHARD O. (1990).

[25] This is entirely due to the increase in inflation differential which has characterized the Euro area in the last years. For possible explanations of this phenomenon, see HONOHAN P. - LANE P. (2003).

Graph 2

CROSS-COUNTRY DISTRIBUTION OF REAL COMOVEMENT
(BY SUBPERIODS)



Source: elaboration based on OECD.

policy has converged greatly starting from the fourth sub-period (which coincides with the convergence process induced by the Maastricht rules)[26]; bilateral trade intensity has on average increased throughout the sample.

Tables 11 and 12 in the Appendix report simple correlation coefficients between the dependent variable and the regressors. For almost all the sub-periods, with the exception of the "Maastricht convergence process" (fourth sub-period), the signs are those expected: a greater extent of divergence in macroeconomic policies and lower trade intensity are associated with lower real comovement. Moreover, it is interesting to notice (Table 13) that countries with higher bilateral trade linkages are

[26] This is true when measuring fiscal policy coordination with FISC1 (fiscal balance) and FISC2 (cyclically adjusted fiscal balance). When we adopt FISC3 (fiscal impulse), there is no evidence for convergence in sub-periods 4 and 5. This is somewhat intuitively clear since the compliance with European fiscal policy rules concerns the public budget rather than the primary balance.

characterized by more intense similarities of their real rates of returns, confirming the conjectures of section 3.1. It is possible to observe also a (weaker) relation between trade intensity and fiscal policy convergence[27].

4.3 *Estimation*

Before presenting the baseline model, Frankel and Rose's specification is first estimated using our dataset. This test may help to understand if the different time horizon and the different sample of countries covered may alter the results.

Table 2 compares the estimation of β in equation *(5)* across the two samples: they look very similar, although both OLS and IV estimations adopting our dataset display a wider effect of trade intensity on business-cycle comovement. So the quantitative prediction of the model does not change much when considering only European countries: rising bilateral trade by one standard deviation leads to an increase of 0.1 in the average cross-country correlation coefficient of business cycles.

Our baseline equation is[28]:

$$(7) \qquad \text{Corr}\,(\tilde{y}, \tilde{y}^*)_t = \alpha + \beta\ \text{TRADE}_{ijt} + \gamma \text{INT}_{ijt} + \delta \text{FISC}(z)_{ijt} + \varepsilon_t$$

Two alternative estimation techniques are proposed: OLS and IV[29]. Tables 3 and 4 present the benchmark results respectively for CICL1 and CICL2. Moreover, the robustness of benchmark estimates is checked with respect to: *i)* sample period stability of the estimated parameters; *ii)* alternative sub-sample's

[27] This may be the result of a spurious relation. In particular, political economy literature has pointed out that institutions matter in the conduction of national fiscal policy, see PERSSON T. - TABELLINI G. (1999). As a result, it is reasonable to expect countries with similar institutions having more similar fiscal policies. On the other hand, similar institutions may be the reflection of cultural linkages which positively affect bilateral trade. This may in principle explain the weak relation we have found between trade intensity and our indicator of fiscal policy convergence.

[28] We use the natural logarithm of TRADE as in the Frankel and Rose specification.

[29] Instruments are those employed by Frankel and Rose. See Table 9 in the statistical annex.

TABLE 2

MEASURING THE TRADE EFFECT ADOPTING FRANKEL AND ROSE'S METHODOLOGY

	Our data set		Frankel and Rose data set
	CICL1	CICL2	
OLS estimation of β	0.09	0.061	0.057
	(0.00)	0.01	(0.00)
IV estimation of β	0.104	0.095	0.087
	(0.00)	(0.00)	(0.00)
R^2	0.04	0.03	0.04
Obs.	438	438	840

Note: Estimation of Equation *(5)*. In the Frankel and Rose equation the dependent variable is the cross country correlation of GDP growth rate (de-trended adopting the H-P filter). *P-value* (robust) on a two tailed test that the coefficients equal 0.

division of the set of observations; *iii)* presence of outliers; *iv)* country size.

Results of benchmark estimations suggest three main points:

First, fiscal policy convergence is associated with a greater extent of business-cycle comovement. This result applies irrespectively of the variable used to measure real comovement (CICL1 and CICL2), and irrespectively of the variable adopted to measure convergence in fiscal policies. Despite the identification problem highlighted in the previous section, the coefficients associated to the variable FISC2 are very similar to those associated to the variable FISC1; a possible explanation relies in the fact that our fiscal convergence variables are averaged over seven years, this contributing to cancel the effects of the cycle. The effect of fiscal policy convergence on business-cycle comovement appears quantitatively relevant: a shift from the third to the first quartile in the distribution of FISC2 is associated to a 26% increase in the cross-country correlation coefficient of cyclical GDP (CICL1) from its mean value. Similar results have been independently reached by Bergman (2004) and Darvas, Rose and Szapàry (2005).

TABLE 3

BENCHMARK ESTIMATION

	CICL1 (OLS)	CICL1 (OLS)	CICL1 (OLS)	CICL1 (IV)	CICL1 (IV)	CICL1 (IV)
TRADE	0.074	0.09	0.09	0.062	0.059	0.073
	(0.00)	(0.00)	(0.00)	(0.04)	(0.04)	(0.01)
FISC1	–0.026			–0.026		
	(0.01)			(0.01)		
FISC2		–0.025			–0.028	
		(0.01)			(0.00)	
FISC3			–0.029			–0.032
			(0.02)			(0.01)
INT	–0.013	–0.005	–0.01	–0.016	–0.011	–0.016
	(0.27)	(0.39)	(0.39)	(0.33)	(0.33)	(0.22)
R^2	0.09	0.10	0.10	0.10	0.10	0.103
F test	13.83	15.93	14.08	9.57	12.59	7.87
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Hansen J test				3.6	3.9	3.1
				(0.16)	(0.13)	(0.20)
Obs.	432	418	418	432	418	418
TEST				0.09	0.07	0.15

Note: Based on estimation of equation *(7)*. CICL1 as dependent variable. Constant omitted. F test on the joint significance of the coefficients. *Hansen J test* on the exogeneity of instruments to eq. *(7)* (*P-value* in parentheses). *P-value* (robust) on two tailed tests in parentheses. The row "TEST" report p-values of the null hypothesis that the coefficient β estimated in equation *(7)* is not significantly lower (one tailed test) than that estimated in equation *(5)*.

Second, a lower spread in real rates of return is associated with greater extent of business-cycle correlation. Despite being negative in all our estimates, the coefficient γ in equation *(7)* is most of the time not significantly different from zero at conventional statistical level. Incidentally, it is comforting that the p-values (two tailed) are stable across the twelve specification ranging from 39% to 6%.

Finally, more intense trade linkages are associated with greater synchronization of cycles. This result holds for both CICL1 and CICL2 and for both the estimation procedures, suggesting the robustness of Frankel and Rose's findings. Incidentally, the size of this effect is lower than that estimated without controlling for policy coordination. The last column in Tables 3 and 4 reports the p-values of the null hypothesis that the coefficient on TRADE in the IV estimation of *(7)* is not significantly lower than that resulting from IV estimation of equation *(5)*. The null hypothesis is rejected at conventional level most of the times.

We now consider some sensitivity checks for our benchmark estimates.

TABLE 4

BENCHMARK ESTIMATION

	CICL2 (OLS)	CICL2 (OLS)	CICL2 (OLS)	CICL2 (IV)	CICL2 (IV)	CICL2 (IV)
TRADE	0.043	0.05	0.059	0.069	0.06	0.071
	(0.02)	(0.00)	(0.00)	(0.01)	(0.04)	(0.01)
FISC1	–0.013			–0.01		
	(0.06)			(0.17)		
FISC2		–0.014			–0.013	
		(0.05)			(0.07)	
FISC3			–0.021			–0.02
			(0.07)			(0.09)
INT	–0.021	–0.017	–0.02	–0.016	–0.015	–0.017
	(0.06)	(0.15)	(0.09)	(0.20)	(0.23)	(0.16)
R^2	0.05	0.05	0.5	0.5	0.05	0.5
F test	6.81	7.65	7.43	6.75	7	6.75
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Hansen J test				4.06	3.9	3.7
				(0.13)	(0.13)	(0.15)
Obs.	432	418	418	432	418	418
TEST				0.12	0.09	0.12

Note: Based on estimation of equation *(7)*. CICL2 as dependent variable. Constant omitted. F test on the joint significance of the coefficients. *Hansen J test* on the exogeneity of instruments to eq. *(7)* (*P-value* in parentheses). *P-value* (robust) on two tailed tests in parentheses. The row "TEST" reports p-values of the null hypothesis that the coefficient β estimated in equation *(7)* is not significantly lower (one tailed test) than that estimated in equation *(5)*.

4.4 *Sensitivity Analysis*[30]

A first check is to consider how the partial correlations change when varying the sample period adopted. We do this in the benchmark specification by excluding one sub-period at a time (Table 14 in the appendix). As can be seen, the coefficient on FISC2 is surprisingly stable across sub-samples: when the dependent variable is CICL1, it fluctuates around –0.025, which is also the benchmark result. Moreover, its statistical significance is robust with regard to this exercise. Also TRADE, to a lesser extent, displays a good degree of stability in both the point estimates and its statistical significance. The worst performer is INT, which displays a high degree of variability in both the point estimates and the 95% confidence bands.

As discussed in section 4.1, the division of the set of observations into five sub-samples is rather arbitrary. To check whether our results are robust with respect to alternative splitting rules we estimate equation *(7)* on:

— The entire sample (cross-sectional specification)
— Two sub-periods (1972-1987; 1988-2004) sample.

Results are presented in Table 5. As can be seen, the main findings of the analysis seem to be preserved. The signs of the coefficients are those expected; their statistical significance appears low in the cross-sectional specification (columns 1 and 2), but increases substantially when dividing the sample into two sub-periods. This pattern may be due to the low number of observations in the cross-sectional specification.

Further, it is checked whether results are robust with respect to the presence of outliers. Outliers are identified as observations (for both the dependent and independent variable) whose numerical value lies below (above) the first (third) quartile by a factor equal to three times the interquartile range (q3-q1). This

[30] Results are presented just for the correlation coefficient of cyclical GDP detrended using the Baxter and King band-pass filter. Figures are very similar when using the Hodrick-Prescott filter and can be made available by the author upon request.

TABLE 5

CONTROLLING FOR ALTERNATIVE SUB-SAMPLE DIVISIONS

	Cross-sectional Specification		Panel (two sub-periods) Specification	
	CICL1 (IV)	CICL1 (IV)	CICL1 (IV)	CICL1 (IV)
TRADE	0.11	0.068	0.11	0.076
	(0.00)	(0.2)	(0.00)	(0.01)
FISC2		–0.025		–0.02
		(0.06)		(0.04)
INT		–0.032		–0.029
		(0.3)		(0.05)
R^2	0.25	0.26	0.18	0.21
Obs.	90	90	180	180
TEST		0.23		0.13

Note: Based on estimation of equation *(7)*. CICL1 as dependent variable. Constant omitted. *P-values* (robust) on two tailed tests in parentheses. The row "TEST" reports *P-values* of the null hypothesis that the coefficient β estimated in equation *(7)* is not significantly lower (one tailed test) than that estimated in equation *(5)*.

amounts to excluding twelve observations[31] in the sample. Results are reported in Table 6. Controlling for outliers does not affect the robustness of the coefficients for bilateral trade and fiscal policy convergence. Rather, the coefficient on monetary policy convergence now becomes significantly different from zero at conventional levels.

As discussed in section 4.1, the trade intensity indicator may prove to be inadequate in capturing phenomena of economic dependence of a small economy on a large one. This may be problematic in the estimation of equation *(7)* because, whenever a small economy is remarkably dependent on its trading relation with a large one, one is likely to observe a large extent of correlation of their indicators of economic performance[32]. As a

[31] All these observations relate to the variable INT. Moreover, nine of these observations come from the Maastricht convergence sub-period while the other three excluded observations belong to the second sub-period.

[32] In this particular situation it would be probably more accurate to speak of causation rather than comovement: a shock hitting the large economy is likely to affect heavily the small one, while the opposite would not be true. I am indebt to an anonymous *referee* for calling attention to this issue.

TABLE 6

CONTROLLING FOR OUTLIERS

	CICL1 (OLS)	CICL1 (OLS)	CICL1 (OLS)	CICL1 (IV)	CICL1 (IV)	CICL1 (IV)
TRADE	0.077	0.10	0.10	0.067	0.06	0.078
	(0.02)	(0.00)	(0.00)	(0.01)	(0.03)	(0.01)
FISC1	–0.025			–0.024		
	(0.01)			(0.02)		
FISC2		–0.023			–0.026	
		(0.00)			(0.00)	
FISC3			–0.028			–0.03
			(–0.02)			(0.01)
INT	–0.025	–0.017	–0.02	–0.023	–0.022	–0.027
	(0.05)	(0.19)	(0.09)	(0.08)	(0.09)	(0.04)
R^2	0.11	0.11	0.10	0.10	0.11	0.10
F test	15.82	17.77	16.21	11.41	11.92	9.78
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Hansen J test				3.537	3.48	3.1
				(0.17)	(0.17)	(0.21)
Obs.	420	406	406	420	406	406
TEST				0.12	0.10	0.20

Note: Based on estimation of equation *(7)*. CICL1 as dependent variable. Constant omitted. F test on the joint significance of the coefficients. *Hansen J test* on the exogeneity of instruments to eq. *(7)* (robust *P-value* in parentheses). *P-value* (robust) on two tailed tests in parentheses. The row "TEST" reports p-values of the null hypothesis that the coefficient β estimated in equation *(7)* is not significantly lower (one tailed test) than that estimated in equation *(5)*.

result, in a cross-section of countries, it could be in principle possible to associate very high degree of business-cycle correlation to low values of the trade intensity indicator, this attenuating the estimated β.

In order to detect whether the problem of country size is relevant in the European sample we propose a test. To show how it works, let us consider two country pairs, the first including economies of approximately the same size (e.g. France-Italy) while the second consisting of countries of different size (e.g. France-Belgium). Let us assume that the two country pairs have approximately the same numerical level of the variable "$TRADE_t$". If the problem highlighted above is relevant in the sample, then the econometrician should observe a greater degree of business-cycle comovement between France and Belgium than that between France and Italy. Repeating the same reasoning for all the country pairs, the marginal effect of TRADE_{ijt} should be expected to be wider whenever the countries considered differ in their size. Accordingly, the following equation is estimated :

$$\text{(8)} \qquad \text{Corr}\,(\tilde{y}, \tilde{y}^*)_t = \alpha + \beta\,\text{TRADE}_{jit} + \phi\,\text{TRADE}_{jit}\times\text{SIZE}_{ijt} + \\ + \gamma\text{INT}_{ijt} + \delta\text{FISC}_{ijt} + \varepsilon_t$$

Where SIZE_{ijt} is a binary variable so defined[33]:

$$\text{SIZE}_{ijt} \begin{cases} 1 & \text{if country } i(j) \text{ is small and country } j(i) \text{ is large in subperiod } t \\ 0 & \text{otherwise} \end{cases}$$

If the sample suffers from the problem highlighted above, ϕ should be significantly greater than zero. Table 7 presents the

TABLE 7

ESTIMATION OF EQ. *(8)*

	CICL1 (IV)	CICL1 (IV)	CICL1 (IV)
TRADE*SIZE	0.003 (0.97)	0.003 (0.97)	0.002 (0.80)
TRADE	0.061 (0.03)	0.059 (0.04)	0.072 (0.01)
FISC1	–0.025 (0.00)		
FISC2		–0.028 (0.00)	
FISC3			–0.032 (0.01)
INT	–0.012 (0.30)	–0.01 (0.30)	–0.016 (0.20)
R^2	0.10	0.10	0.09
Obs.	418	418	418
TEST	0.09	0.07	0.15

Note: Based on estimation of equation of eq. *(8)*. CICL1 as dependent variable. Constant omitted. *P-values* (robust) on two tailed tests in parentheses. The row "TEST" reports *P-values* of the null hypothesis that the coefficient β estimated in equation *(8)* is not significantly lower (one tailed test) than that estimated in equation *(5)*.

[33] A country is defined to be "small" ("large") if its real GDP is lower (larger) than the 25th (75th) percentile in sub-period *t*.

results. As can be seen, the null hypothesis that ϕ equals zero is never rejected at conventional level.

4.5 *Summary of Results*

The benchmark analysis and the robustness checks point out three main results:

First, countries with similar fiscal policies tend to have more similar cycles. This result is robust to variation in the sample period considered. Moreover, the discussion in section 4.1 and the adoption of cyclically adjusted measure of the fiscal balance seems to suggest that the estimated δ in equation *(7)* reflects the causal impact of fiscal policy convergence on real comovement.

Second, a low spread of short term real rates of return is associated to a greater extent of business-cycle synchronization. The benchmark estimates display a low statistical significance and a low extent of sub-sample stability of the parameter γ in equation *(7)*, this casting doubts on the robustness of these results. Nonetheless, sensitivity analysis has shown that the statistical significance of this coefficient increases when cheeking for the presence of outliers in the sample.

Third, greater trade intensity is associated to high degree of cross-country correlation coefficient of business cycles, confirming the claim by Frankel and Rose and other studies testing their hypothesis. Such an effect appears robust even when controlling in the econometric specification for coordination of macroeconomic policies. In this latter case, by the way, the size of the effect significantly weakens. Graph 3 presents weighted average of our benchmark estimates of the trade effect with those estimated according to Frankel and Rose's procedure.

The striking result concerns the fact that benchmark IV estimates are on average 33% lower than those estimated with Frankel and Rose's procedure. To show why this is the case, one may consider the analytic expression for the IV coefficient of β from equation *(5)*.

$$\tilde{\beta}_{IV} = \frac{COV(y, \text{TR}\hat{\text{A}}\text{DE})}{VAR(\text{TR}\hat{\text{A}}\text{DE})} \tag{9}$$

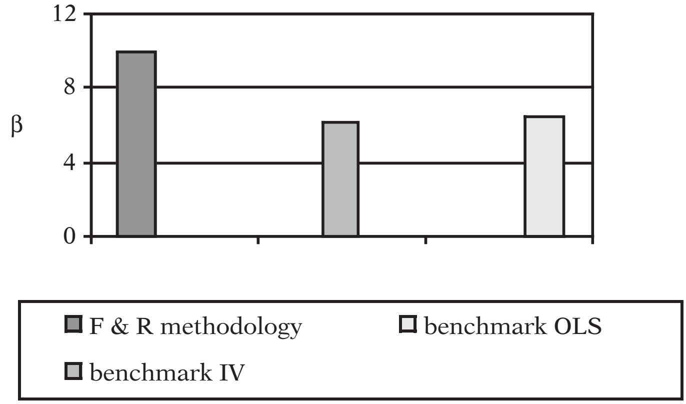
TRÂDE being the vector of fitted values from the first stage regression, and y being our dependent variable. Assuming that the true model is, instead, represented by *(7)* and assuming the orthogonality between TRÂDE and the disturbance term in equation *(7)*, we can express *(a)* as:

$$\tilde{\beta}_{IV} = \beta\frac{COV(\text{TRADE},\text{TR}\hat{\text{A}}\text{DE})}{VAR(\text{TR}\hat{\text{A}}\text{DE})} + \delta\frac{COV(FISC, \text{TR}\hat{\text{A}}\text{DE})}{VAR(\text{TR}\hat{\text{A}}\text{DE})} + \gamma\frac{COV(INT, \text{TR}\hat{\text{A}}\text{DE})}{VAR(\text{TR}\hat{\text{A}}\text{DE})} \tag{10}$$

The sample correlation of TRÂDE and the policy variables is given in Table 8. As can be seen, the sample moments seem to

GRAPH 3

COMPARISON BETWEEN OUR BENCHMARK RESULTS AND FRANKEL AND ROSE'S METHODOLOGY



Note: Weighted average (robust *t-statistic* as weights) of β estimated according to the different specifications.

confirm the idea put forth by Gruben, Koo and Millis that countries sharing the same border and the same culture (language) are characterized by a more intense coordination of their policies. The exclusion of the policy variables from *(5)* is, thus, expected to produce an overestimation of β in Frankel and Rose's specification, since:

$$\tilde{\beta}_{IV} \xrightarrow[p]{} \beta + k \qquad k > 0 \tag{11}$$

TABLE 8

CORRELATION BETWEEN REGRESSORS AND $\hat{\text{TRADE}}$

	$\hat{\text{TRADE}}$ (fitted value from 1st stage regression)
FISC1	–0.22
FISC2	–0.225
FISC3	–0.10
INT	–0.19

5. - Conclusions: Policy Implication for EMU

A symmetric distribution of shocks is one of the criteria that the theory developed by Mundell (1961) identifies as being a determinant for the economic success of a currency area. Building a common monetary policy may be inefficient and may produce macroeconomic stability problems when countries taking part in the arrangement are characterized by a low extent of business-cycle synchronization. This view has been criticized by Frankel and Rose (1998), who have argued that this criterion should be considered as endogenous to the construction of the currency area: in fact, fixed exchange rates should stimulate intra-union trade intensity thus contributing to more synchronous cycles and to a better functioning of the common monetary policy. As a support of their thesis, the two authors have estimated a positive and sizable effect of bilateral trade intensity on business-cycle comovement using a panel of OECD countries. This paper has extended Frankel and Rose's econometric approach. Building on the considerations presented by Gruben, Koo and Millis (2002) we have explicitly taken into account the role of policy convergence on real comovement. Our empirical analysis suggest three main points:

i. Coordination of macroeconomic policies seems to help business-cycle comovement: countries that have similar fiscal policy seem to co-move more than countries with idiosyncratic policies, this result being robust with regard to a number of sensitivity checks. Moreover, we have also found a weak evidence that convergence in short term real interest rates has been beneficial for the intra-European comovement of business cycles, especially when checking for the presence of outliers;

ii. Bilateral trade intensity positively affects business-cycle correlation, thus confirming Frankel and Rose's claim. The effect remains statistically significant even when including in the analysis the policy coordination variables, this suggesting the robustness of Frankel and Rose's study.

iii. European data appear to support the idea put forth by Gruben, Koo and Millis on the endogeneity of the set of instruments proposed by Frankel and Rose. Countries being geographically close and sharing the same language tend to have more similar fiscal and monetary policies. This furnishes an explanation of the discrepancies between our estimates of the "trade effect" and the one implied by Frankel and Rose's procedure. The downward revision is sizable, ranging on average between 36%-33%.

These findings suggest a number of implications for the functioning of the European Monetary Union (EMU).

As a first point, the results are consistent with the view that patterns of business-cycle comovement in Europe are not going to change much in the short run as a result of the adoption of the Euro. On the one hand, recent empirical estimations of the impact of the Euro on intra-EMU trade have not identified a dramatic boost after 1998 (see De Nardis and Vicarelli, 2003; Micco, Ordonez and Stein, 2003). On the other hand, our estimation of the effect of trade intensity on business-cycle comovement suggests that a move from the median to the 75th percentile of TRADE (which correspond to an increase in bilateral trade intensity by 150%) increases business-cycle correlation by 15% of the mean value. Together, these two factors may lead us conjecture that Frankel and Rose's hypothesis is likely to operate in Europe in the long run.

Another implication of the analysis is that fiscal policy rule may be to some extent beneficial for the macroeconomic policy framework of the EMU. Early studies on the *rationales* of the Maastricht treaty and on the Stability and Growth Pact (SGP) (Lamfalussy, 1989; Bovemberg, Kremer and Masson, 1991; Buiter, Corsetti and Roubini, 1993; Artis and Winckler, 1999) have identified financial stability and ECB credibility as the main benefits for the adoption of fiscal policy rules. The results in this paper suggest that fiscal policy rules may also help the stabilizing effort of the European Central Bank: to the extent that the SGP prevents idiosyncratic behaviour in the carrying out of domestic fiscal policies, we should observe more synchronous cycles between European countries and thus a more efficient common monetary policy. This conclusion, nonetheless, needs to be weighted with the fact that domestic fiscal policies in Europe have become a more important buffer for negative shocks than in the past: the adoption of the single monetary policy, the absence of a federal risk sharing arrangement (e.g. Farina and Tamborini, 2002) and the low extent of intra-European labor mobility are all factors suggesting that national fiscal policies are one of the few instruments that European countries have to counteract idiosyncratic movements in real variables. The solution of this *trade-off* between "European rules" and "national discretion" represents one of the crucial points Europe should face to improve its short run macroeconomic policy frame and to increase the economic success of the Euro.

APPENDIX

Table 9

THE DATA SET

Variable	Description	Source	Missing Observations	Note
CICL1	Cross country correlation coefficient of cyclical GDP growth rate (H-P filtered)	OECD (Retrospective Statistics, 2002; OECD Economic Outlook)	//	//
CICL2	Cross country correlation coefficient of cyclical GDP growth rate (Baxter and King band pass filtered)		//	//
FISC1		OECD (Retrospective statistics, 2002 and OECD Economic Outlook	//	//
FISC2		Our elaboration on OECD (Retrospective statistics, 2002; OECD Economic Outlook, various years)	Portugal (1972-1978)	See note 16 in the text for the computation of the cyclical component of fiscal balance
FISC3	See section 4.1			
INT		AMECO database (EU-Commission)	Greece and Sweden (1972-1978)	Private consumption deflator in computing short term real rate of return

(continued on next page)

(continued) TABLE 9

THE DATA SET

Variable	Description	Source	Missing Observations	Note
TRADE		ITCS database, OECD		data on Austria are from of Trade IMF direction of Trade statistics
	Ling	Dummy variable that takes the value of 1 if the two countries share the same language	Frankel and Rose data-set	
Instruments	Lndist	Natural logarithm of the distance (in miles) between the two "economic" capital of the country pair	Frankel and Rose data-set	
	Adjacent	Dummy variable that takes the value of 1 if the two countries share the same language	Frankel and Rose data-set	

Table 10

SUMMARY STATISTICS, BY SUBPERIODS

		1972-1978	1979-1985	1986-1992	1993-1999	2000-2004	1972-2004
CICL1	P25	0.013	–0.013	–0.0262	0.21	0.33	0.093
	P50	0.48	0.29	0.31	0.63	0.71	0.47
	P75	0.76	0.61	0.54	0.77	0.88	0.74
	Mean	0.38	0.25	0.27	0.47	0.47	0.37
	Cv	1.15	1.50	1.23	0.92	1.16	1.19
CICL2	P25	0.012	0.15	0.13	0.31	-0.03	0.11
	P50	0.42	0.45	0.48	0.62	0.67	0.51
	P75	0.68	0.60	0.67	0.79	0.865	0.74
	Mean	0.34	0.38	0.41	0.49	0.41	0.40
	Cv	1.18	0.87	0.82	0.77	1.35	1.01
FISC1	P25	2.03	2.54	2.60	1.65	1.9	2.07
	P50	3.75	4.10	4.47	2.48	2.72	3.33
	P75	5.9	7.04	7.47	4.1	4.2	5.72
	Mean	4.27	4.97	5.38	3.19	3.1	4.15
	Cv	0.64	0.61	0.61	0.62	0.57	0.67
FISC2	P25	2.14	2.57	2.79	1.62	1.89	2.08
	P50	3.8	4.13	4.72	2.32	2.71	3.2
	P75	6.08	6.9	7.33	3.54	4.25	5.57
	Mean	4.48	4.9	5.37	2.62	3.1	4.11
	Cv	0.63	0.59	0.6	0.5	0.59	0.67
FISC3	P25	1.64	2.16	1.99	1.89	1.38	1.75
	P50	2.94	3.14	2.85	2.95	1.98	2.82
	P75	4.50	4.02	4.05	4.19	2.9	4.05
	Mean	3.27	3.22	3.24	3.2	2.34	3.05
	Cv	0.62	0.42	0.47	0.51	0.55	0.53

(continued on next page)

continued Table 10

UMMARY STATISTICS, BY SUBPERIODS

		1972-1978	1979-1985	1986-1992	1993-1999	2000-2004	1972-2004
TRADE	P25	0.003	0.003	0.0046	0.0048	0.0039	0.0042
	P50	0.007	0.006	0.0076	0.008	0.006	0.007
	P75	0.016	0.017	0.021	0.02	0.017	0.018
	Mean	0.013	0.014	0.016	0.0164	0.015	0.0152
	Cv	1.25	1.23	1.10	1.05	1.14	1.146
INT	P25	2.9	2.3	1.5	1.15	0.76	1.35
	P50	3.92	2.9	2.21	1.55	1.182	2
	P75	5.4	4	3.07	2.12	1.8	3.27
	Mean	4.01	3.53	2.4	1.4	2	2.58
	Cv	0.42	0.56	0.49	0.74	0.56	0.67

Note: P25, P50 and P75 are, respectively, the 25^{th}, the median and the 75^{th} percentile; Cv is the coefficient of variation.

Table 11

CORRELATION AMONG REAL COMOVEMENT (CICL1)
AND REGRESSORS

	1972-1978	1979-1985	1986-1992	1993-1999	2000-2004	1972-2004
TRADE	0.24	0.06	0.20	–0.03	0.39	0.20
FISC1	–0.61	–0.19	0.02	0.05	–0.28	–0.29
FISC2	–0.67	–0.22	0.005	0.05	–0.25	–0.29
FISC3	–0.67	–0.20	–0.14	–0.18	–0.04	–0.25
INT	0.14	–0.17	–0.14	0.13	–0.10	–0.16

Table 12

CORRELATION AMONG REAL COMOVEMENT (CICL2)
AND REGRESSORS

	1972-1978	1979-1985	1986-1992	1993-1999	2000-2004	1972-2004
<						
TRADE	0.37	0.02	0.17	–0.03	0.24	0.19
FISC1	–0.21	–0.12	–0.06	0.19	–0.13	–0.13
FISC2	–0.20	–0.16	–0.09	0.12	–0.15	–0.16
FISC3	–0.32	–0.15	–0.33	0.04	–0.04	–0.12
INT	–0.002	–0.33	–0.1	0.13	–0.24	–0.15

Table 13

CORRELATION BETWEEN REGRESSORS

	TRADE	FISC1	FISC2	FISC3	INT
TRADE	*	*	*	*	*
FISC1	–0.19	*	*	*	*
FISC2	–0.20	0.95	*	*	*
FISC3	–0.12	0.50	0.52	*	*
INT	–0.25	0.24	0.25	0.1320	*

TABLE 14

ESTIMATION OF EQUATION *(7)* BY SUBPERIOD (CICL1)

	No 1st sub-period		No 2nd sub-period		No 3rd sub-period		No 4th sub-period		No 5th sub-period	
	OLS	IV	OLS	IV	OLS	IV	OLS	IV	OLS	IV
TRADE	0.079	0.058	0.11	0.08	0.1	0.05	0.12	0.09	0.06	0.014
	(0.01)	(0.07)	(0.00)	(0.02)	(0.00)	(0.1)	(0.00)	(0.01)	(0.03)	(0.65)
FISC2	–0.02	–0.02	–0.03	–0.03	–0.03	–0.03	–0.01	–0.02	–0.022	–0.025
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.04)	(0.02)	(0.00)	(0.00)
INT	–0.02	–0.02	0.021	0.01	–0.05	–0.01	–0.005	–0.009	–0.008	–0.018
	(0.1)	(0.09)	(0.2)	(0.3)	(0.66)	(0.36)	(0.7)	(0.5)	(0.5)	(0.17)
OBS.	363		327		328		327		327	

Note: This table reports the estimated coefficients of regression *(7)* excluding one sub-period at time. The dependent variable is CICL1. Constant not reported. Robust p-value (two-tailed) in parentheses.

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