TRADE AND BUSINESS CYCLE CO-MOVEMENT: EVIDENCE FROM THE EU

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Abstract

This paper studies empirically the determinants of business cycle co-movement using a panel of European countries (1972-2004). We find that both policy convergence (in particular fiscal policy) and bilateral trade intensity are robust determinants of real co-movement in Europe, this confirming the seminal study by Frankel and Rose (1998), and more recent finding by Bergman (2004) and Darvas, Rose and Svapary (2005). Moreover ,once controlling for policy convergence, the effect of bilateral trade on business cycle co-movement weakens by a factor of 36%-33%. This finding is interpreted as being evidence in favour of the recent claim by Gruben, Koo and Millis (2002) that Frankel and Rose econometric procedure suffers from omitted variables bias and endogeneity in the set of instruments.
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1 Introduction

One of the main results of optimum currency area (OCA) theory is that countries whose frequency of idiosyncratic shocks is high are less suitable to take part to a fixed exchange rate regime[1]. The reason lies in the idea that joining a currency area limits the ability of countries to use national monetary policy to respond to country-specific shocks. This finding has been successively extended into

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[1] This idea has been proposed first by Mundell (1961) and has been further specified by Kenen (1969). The last 20 years have seen a growing body of articles in the area of OCA theory (see Tavlas, 1994). On the empirical implementation of the criteria of optimum currency area theory see Bayoumi and Eichengreen (1999). On the theoretical side, an excellent formalization is provided by Alesina and Barro (2002).

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 co-movement, 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 change dramatically the historical record of real co-movement: indeed, the boost in trade intensity between countries participating to 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 large effect of bilateral trade intensity on the co-movement of cycles between 21 OECD countries.

Frankel and Rose study have stimulated a growing body of empirical literature that has further investigated the determinants of business cycle co-movement. Perturbation of the basic model (see in particular Baxter and Kouparitsas, 2005) did not question the positive “trade effect” although its magnitude has been 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 co-movement may cast some doubts on the economic relevance of this Lucas critique to the “synchronicity” criteria of OCA theory.

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[2] See De Grauwe (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 and Eichengreen (1993). See Obstfeld and Peri (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 and Pesanti (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.

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 between 1972 and 2004 we find that countries having similar fiscal policies and similar real rate of return are likely to have more synchronized business cycles. Moreover, it is found that the impact of bilateral trade on real co-movement, although positive and highly statistically significant, is lower (between 48% and 36%) than that estimated according to Frankel and Rose procedure. These results can be interpreted as being evidence in favour of Gruben, Koo and Millis claim. The paper is organized as follows: Sections 2 and 3 deal with Frankel and Rose 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 in considering this criterion endogenous to the constitution to the currency area itself, this rendering less relevant its ex-ante compliance. Their idea is based on two main conjectures:

- Fixed exchange rate should promote trade between countries sharing the agreement;
- Higher bilateral trade 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 sign of the effect is concerned, there is little ambiguity empirically that a common currency increases trade linkages between countries adopting it.

More controversial from the point of view of economic theory seems the relation between trade intensity and business cycle synchronization. To formally present the channels trough which trade may affect real co-movement it is assumed, as in Frankel ad Rose (1998), that the growth rate of output of a country may be expressed as:

$$\Delta y_t = \sum_{i=1}^{n} a_i u_{it} + v_t + g \qquad [1]$$

Where $\{u_i\}$ is the deviation of sector i 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 trend growth rate of GDP. Denoting by $\tilde{y}_t$ the de-trended growth rate of output at time t of a country we have[6]

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

Equation [2] tells us that the co-movement 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 cycle, and the covariance between the two countries' aggregate shocks. Imagining the two polar cases, equation [2] collapses to[7]

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[5] In particular, cross-sectional studies tend to estimate a "currency union effect" much higher than time-series studies (see Glick and Rose, 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* (see Personn, 2001). A time series approach which deals explicitly with the problem of reverse causality is Micco, Ordonez and Stein (2003), who have estimated a positive impact of the euro on intra-EMU trade between 4% and 23%.

[6] It is assumed that $\{u_i\}$ is distributed independently across both sectors and time and that $\{v_i\}$ is distributed independently of the sector specific shocks. This last assumption may be somewhat restrictive for countries characterized by a low extent of diversification, since industry specific shocks will have a strong effect on the aggregate state of the economy.

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

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

When two countries are fully specialized in the production of two different goods while, in the case of two highly diversified economies, it becomes[8]:

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

c being the share that each industry has got on total output and *n* the number of industries in the two economies.

As it is possible to see from equation [3] and [4] two countries with similar industrial structure tends to have more synchronous business cycles. Moreover, business cycle correlation is likely to be positively related to the extent of co-movement of country-specific aggregate shocks.

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

A first channel regards the cyclical behaviour of the trade balance. When a country is affected by a negative aggregate shock it will tend to import less and to export more with its trading partners: the economic relevance of such spill-over effects depends on the extent of bilateral trade[9]. 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 got on the industrial structure of a country. On the one hand, reduction of transaction costs may induce countries to specialize in sectors in which they have comparative advantages so that international trade should induce a divergence of $\{a_i\}$ between partners. Hence, the adoption of a single currency may in principle reduce the extent of similarity of the industrial structure between countries thus leading to more asynchronous business cycles, an idea that have been supported by Krugman (1993) for the case of the

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

[9] In particular, Prasad and Gable (1997) have pointed out the relevance 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.

European Monetary Unification process. On the other hand, the adoption of the common currency may enhance the similarity of $\{a_i\}$ between trading partners if the increase in trade takes place more *within* industries rather than *between* them. As a result, it is not possible a-priori to say if the adoption of a single currency induces more synchronous cycles: if inter-industry trade dominates over intra-industry trade, then specialization effects may produce less synchronous business cycles between countries taking part to a currency area.

Frankel and Rose have tackled empirically the theoretically ambiguous effect of trade over business cycles co-movement. The two scholars have estimated, using a panel of 21 OECD countries, the following equation:

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

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 should amplify the extent of co-movement of real variable between countries taking part to the monetary union. The size of the coefficient will then identify the economic relevance of this phenomenon: a *large* value of β will imply that real co-movement will sharply increase after the adoption of the single currency so that Frankel and Rose's hypothesis will 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" seems to be robust to the macroeconomic series used to measure cross-country real co-movement 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 co-movement 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 confirms 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 synchronicity of cycles which will render *ex post* the common monetary policy more efficient. This result 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, thus, crucial to settle the robustness of Frankel and Rose's findings. This section discuss 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 let OLS estimation of the coefficient β to be bias 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:

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

Where I_t is a set of relevant variables that have been excluded from the analysis. A first reason for endogeneity arises here 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 cycles' synchronization[10]. *On the other hand*, it is reasonable to expect a positive relation between trade intensity and

[10] Checchi (1989) has supported the idea that business cycle co-movement between G-7 countries have 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 similar monetary policies; a second reason, build on a "discipline argument", has been sustained by Artis and Zhang (1995), who have suggested that monetary policy may itself be a source of shocks. Participation to a fixed exchange rate regime, producing more coordination, should reduce idiosyncratic behaviours and should enhance real co-movement. As an evidence, the two authors reported the emergence of an "European business cycle" during the period of the European Monetary System (EMS).

monetary policy coordination, especially in the context of EU where incentives to coordinate policies during the 80s arose within the fixed exchange rate agreement of the European Monetary System. In fact, as OCA theory suggests, the benefits to participate to a fixed exchange rate regime are positively linked to the extent of bilateral trade so that we should expect participation to 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 intensity between countries part of the agreement. As a result, in the sample period considered, we should expect convergence in monetary policy to occur more between countries with intense trade relations: as section 4 will show, such claim will be 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 higher real co-movement. Simultaneity would thus underestimate the "trade effect".

Table 1: Sources of inconsistency of OLS estimation of β in eq. [5]

Source of inconsistency	Direction of the bias
Measurement errors	Underestimation of β
Endogeneity	Underestimation of β (Simultaneity) Overestimation of β (exclusion of monetary policy coordination from equation [5])

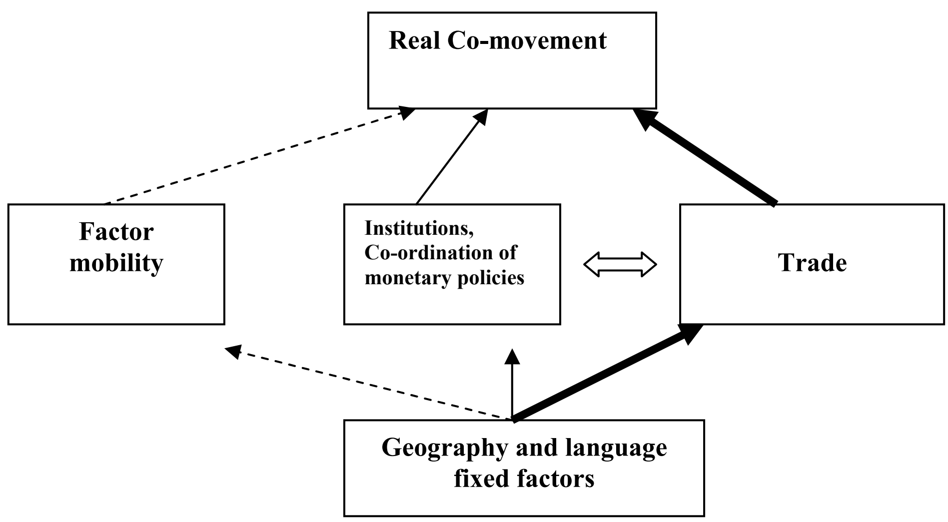
Table 1 summarizes the main sources of inconsistency of OLS and the "likely" direction of the bias of the coefficient β when estimating equation [5] with OLS. Accordingly with these concerns, Frankel and Rose proposed IV estimation of [5], relying on a set of instruments borrowed from gravity models of international trade: geographic and cultural proximity[11].

[11] Geographic proximity is measured by a dummy variable that take 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.

3.2 Are geography and language good instruments?

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

Figure 1: "Gruben,Koo and Millis hypothesis"



Source: Gruben, Koo and Millis (2003), pag. 21. A thick line indicates Frankel and Rose's hypothesis.

Gruben, Koo and Millis did not test explicitly this hypothesis. Rather, they have 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. Result of their estimation suggests a much lower "trade effect" (circa 50% smaller). In part, this finding is consistent with recent literature that has further investigated the relation between trade intensity and cycle synchronization[12]. 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 allow to distinguish among the determinants of real co-movement that has been made explicit in figure 1.

4. Empirical Analysis

This section deals with the relation between trade intensity and business cycle co-movement for a group of European countries. The aim is to provide an estimation of the "trade effect" robust to both the problems highlighted in the previous section. The work by Frankel and Rose and by Gruben, Koo and Millis is extended 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 co-movement: first, fiscal policy may be itself a source of shock for an economy so that its coordination, preventing idiosyncratic behaviours, should increase real co-movement; secondly, coordination of fiscal policies may increase business cycles correlation between countries when the geographic distribution of shocks is symmetric.

The basic framework is that of equation [6]. The first subsection describe the data set used. Then the focus is on descriptive analysis of the main variable involved. Finally, estimation of parameters and sensitivity analysis are performed.

[12] On cross-sectional studies see Baxter and Kouparitsas (2005) and Imbs (2003, 2004). In those studies the "trade effect" seems to extend also for developing countries, although Imbs finds that including a measure of cross-country similarity in the industrial structure lowers the estimates of the trade effect, this suggesting omitted variable bias in Frankel and Rose specification. In a panel data framework, Kose, Prasad and Terrones (2003) have found mixed evidence for a positive effect of trade on business cycle correlation and only for industrialized countries.

4.1 Description of the data set

The analysis deals with countries belonging to the EU-14[13] observed between 1972 and 2004, the time span being split in five sub-samples[14].

Real co-movement is measured by cross-country correlation coefficient of cyclical GDP growth rate. For each country pair (i.e. France-Germany) it is observed one measure of real co-movement at each sub-period so that the maximum sample size is made by 455 observations. Business cycle is measured as the difference between actual and trend GDP growth rates, using two alternative filters: i) Baxter and King (1999) band pass filter (CICL1); ii) Hodrik and Prescott filter (CICL2). Although Baxter and King band-pass filter has become a standard in this literature, the adoption of the Hodrick and Prescott filter makes the analysis more comparable with those of Frankel and Rose. Moreover, 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 the literature. Denoting by $TRADE_{ijt}$ bilateral trade intensity between country *i* and country *j* in sub-period *t* we have:

$$TRADE_{ijt} = (X_{ijt} + M_{ijt}) / (X_{it} + M_{it} + X_{jt} + M_{jt})$$

X_{ijt} (M_{ijt}) being total exportation (importation) from country *j* to country *i* and vice-versa. The higher the value of $TRADE_{ijt}$ the higher will be trade intensity between country *i* and *j* at sub-period *t*.

As it was argued before, fiscal and monetary policy similarities should positively influence real co-movement between country pairs. The problem, now, relates to their quantitative measurement. As far as monetary co-ordination is concerned, it is proposed the distance indicator:

$$INT_{ijt} = \sum_k |r_{ik} - r_{jk}| / k$$

[13] Countries are: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain, Sweden, UK.
[14] The sub-periods are: 1) 1972-1978; 2) 1979-1985; 3) 1986-1992; 4) 1993-1999; 5) 2000-2004

r being 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 the spread of short term real rate of return between country *i* and country *j*.

Measurement of fiscal policy co-ordination poses more problems. Studies that have seek to gauge the impact of fiscal policy convergence on real co-movement (i.e Clark and Van Wincoop; 2001, Bergman, 2004; Darvas, Rose and Swapàry, 2005) have all used various indicators based on elaboration of the fiscal balance. By the way, there are very good reasons to think that those measure may induce reverse causality[15]. Assume for instance that France and Germany are experiencing a boom of their economies. Holding fixed discretionary fiscal policies, the two countries will both face a reduction of their fiscal deficits due to the working of automatic stabilizer. What the econometrician will observe in such cases is that cycle synchronization between France and Germany takes place in presence of a co-movement of fiscal deficits, this being the effect of the cycle itself. As a result of this reverse causality problem, OLS will tend to overestimate the impact of fiscal policy coordination on business cycles co-movement. Cyclically adjusted balances[16] should solve this problem. Consistently, we propose three different measures of fiscal policy co-ordination:

$$FISC(k)_{ijt} = \sum_k |D_{ik} - D_{jk}| / k$$

where D_{ik} is: a) fiscal balance ($FISC1_{ijt}$); b) cyclically adjusted fiscal balance ($FISC2_{ijt}$); c) *fiscal impulse* ($FISC3_{ijt}$)[17].

A detailed description of data sources is in table 2 in the appendix.

[15] Darvas, Rose and Svapàry have made explicit this fact by instrumenting their fiscal policy convergence variable.

[16] To compute the cyclical component of net lending it is adopted the procedure in Giorno et al (1995): the elasticity of the fiscal balance is multiplied by the output gap by the output gap. The elasticities are from Van den Noord (2000) (subperiods 1-4) and from Giruard and André (2005) (sub-period: 5). The output gap is the difference between actual and trend (H-P filtered) GDP growth rate.

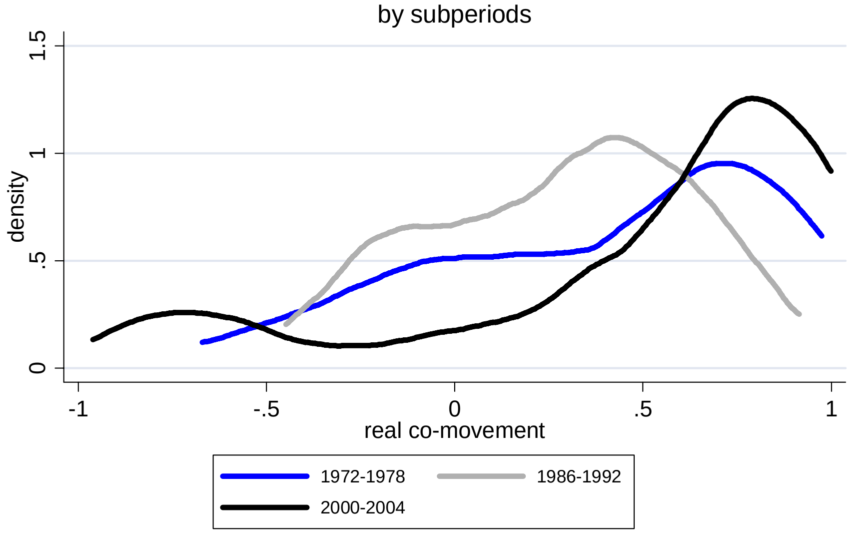
[17] The "fiscal impulse" is defined to be the variation of the cyclically adjusted primary balance. For a description of this indicator see Blanchard (1990).

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 co-movement. Table 3 shows a soft evidence for an increase in intra-European synchronization of cycles, although this path seems to depend on the variable we adopt: CICL2 median tends to increase monotonically over time while the same does not happen for CICL1. To investigate further this aspect figure 2 presents kernel density estimates 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 the literature (Bayoumi and Eichengreen, 1993; De Cecco and Perri, 1996). The evolution over time of the distribution 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 co-movement mode displays a much greater fraction of the population.

Figure 2: Cross-country distribution of real co-movement

by subperiods



source: our elaboration on OECD

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 rate of return between European countries, although this tendency has stopped in the last sub-period[18]; fiscal policy has converged much since the fourth sub-periods (which coincide with the convergence process induced by the Maastricht rules)[19]; bilateral trade intensity has on average increased throughout the sample.

Table 4 and table 5 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: higher divergence of policies and lower trade intensity are associated with lower real co-movement. Moreover, it is interesting to notice (Table 6) that countries with higher bilateral trade linkage are characterized by more intense similarities of their real rate of returns, this confirming the conjectures of section 3.1. It is possible to notice also a (weaker) relation between trade and fiscal policy convergence[20].

4.3 Estimation

Before presenting the baseline model, it is first run Frankel and Rose's specification using our data. This test may help understanding if the different time horizon and the different sample of countries covered may alter results. Table 7 compare estimation of β across the two sample: they look very similar, although both OLS and IV estimation adopting our data set display a greater trade effect on business cycle co-movement. So, the quantitative prediction of the model do not change much when considering only European countries: rising bilateral trade by one

[18] This is entirely due to the increase in inflation differential which has characterized the euro area in the last years. On possible explanations of this phenomenon see Honohan and Lane (2003).

[19] This is true when measuring fiscal policy coordination with d1 (budget balance) and d2 (cyclically adjusted budget balance). When we adopt d3 (fiscal impulse), there is not evidence for convergence in sub-period 4 and 5. This is somewhat intuitive since the compliance with European fiscal policy rules concerns the public budget, rather than the primary balance.

[20] This may be the result of a spurious relation. In particular, the political economy literature (see Persson and Tabellini, 1999) have pointed out that institutions matter in the conduction of national fiscal policy so that it is reasonable to expect countries with similar institution having more similar fiscal policies. On the other hand, similar institution 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.

standard deviation leads to an increase of 0.1 in the average cross country business cycles correlation coefficient.

Table 7: 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 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.

Our baseline equation is[21]:

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

Two alternative estimation techniques are proposed: OLS and IV[22]. Table 8 and Table 9 present the benchmark result respectively for CICL1 and CICL2. Moreover, the robustness of benchmark estimates is controlled for: i) endogeneity of the policy variables; ii) sample period stability of the estimated parameters; iii) presence of outliers.

4.3.1 Benchmark Estimates

Results of benchmark estimation (Table 8; for CICL2 see Table 9 in the appendix) suggest three main points:

First, fiscal policy convergence is associated with a greater degree of business cycle co-movement. This result applies irrespectively of the variable adopted to measure real co-movement (CICL1 and CICL2), and irrespectively of the variable adopted to measure convergence in fiscal policies (FISC1 and FISC3). Moreover, the effect appears also quantitatively relevant: a shift from the third to the first quartile of

[21] We use natural logarithm of TRADE as in Frankel and Rose specification.
[22] Instruments are those employed by Frankel and Rose. See table 2 in the statistical annex.

the distribution of FISC1 should induce a 26% increase in cross-country correlation coefficient of cyclical GDP (CICL1) from its mean value.

Table 8: 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)	(0.65)	(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	432	432	418	418	418

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.

Second, a lower spread in real rate of return is associated with greater co-movement of cycles: despite the fact that the coefficient is negative in all our estimates, most of the time it is not significantly different from zero at conventional statistical level, this casting doubts on the relevance of such effect.

Finally, more intense trade linkages are associated with greater synchronization of cycles. This result holds for both CICL1 and CICL2 and for both estimation procedure, suggesting the robustness of Frankel and Rose findings. By the way, the size of this effect is lower than that estimated without controlling for policy coordination. This last result, if robust, would confirm Gruben et al. (2002) claim.

We now consider some sensitivity checks for our benchmark estimates.

4.3.2 Endogeneity of policies

Policy makers partly target real variables when deciding policies, and the output gap is without doubt one of the variables they look at. Such observation highlights a potential identification problem in our benchmark estimates: that of endogeneity. Consider two countries experiencing a boost of their economy: if policy maker adopt countercyclical fiscal policies, it should be observed an amelioration of the fiscal balance in both countries. As a result, business cycle correlation will be associated with a co-movement of the fiscal balance, although such result would not reflect causality of the latter. These considerations appear particularly relevant in the context of fiscal policies, where the working of automatic stabilizers may render this problem even worst (see Section 4.1). By the way, as far as fiscal policy is concerned, such problem may be attenuated by the following considerations:

i. Existing empirical works on European countries point out the low sensitivity of fiscal policies to the output gap, especially in the first part of the sample period considered (Gali and Perotti, 2003);

ii. It is possible to control for the automatic component of the fiscal balance by using as a fiscal coordination variable FISC2. As can be seen from Table 8, the estimated coefficients look very similar to that of FISC1. A possible explanation lies in the fact that our variables are averaged over seven years, this contributing to cancel out the effects of the cycle on our fiscal coordination variable.

4.3.3 Sensitivity to the time span considered and to outliers[23]

A further check is to consider how the partial correlation change when varying the sample period adopted. We do this in the benchmark specification by excluding one sub-period at time (Table 10 and Table 11 in the appendix). As it is possible to see, the coefficient on FISC1 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 to this exercise. Also TRADE, to a less extent, displays a good degree of stability in both the point estimates and its statistical significance. The worst performer is INT, which displays

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

an high degree of variability in both the point estimates and the 95% confidence bands.

A final check consist to see whether results are robust to the presence of outliers. We identify outliers as observation (for both the dependent and independent variable) whose numerical values lies below (above) the first (third) quartile by a factor equal to three times the interquartile range (q3-q1). This amount to loose 10 observations in our sample. Results are reported in Table 11. 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 become significantly different from zero at conventional levels.

Table 11: 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.	422	408	408	408	408	418

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.

4.4 Summary of Results

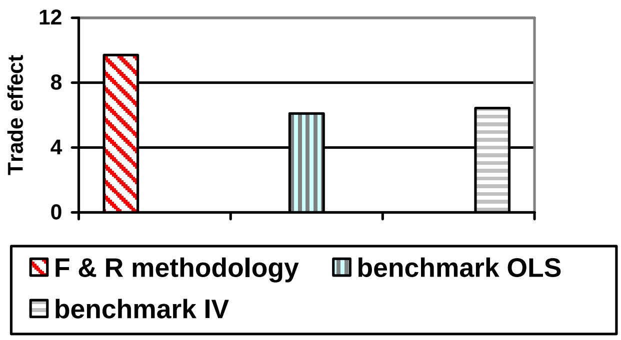
The benchmark analysis and the robustness checks point out three main results:

First, countries with similar fiscal policies tend to have more similar cycles. This result is robust to variation in the sample period considered. Moreover, although this point deserve further work, our discussion seems to suggest that the coefficient is in large part reflecting the causal impact of fiscal policy convergence on real co-movement.

Second, countries with a low spread of short term real rate of return are characterized by a greater degree of synchronization when controlling for the presence of outliers. By the way, the endogeneity problems highlighted in section 4.3.2, the low significance of the coefficients when working with the full sample and the low extent of sub-sample stability of the parameters cast doubts on the robustness of these results.

Third, greater trade intensity boost real co-movement this confirming the claim by Frankel and Rose and further studies testing their hypothesis. Such effect appear robust even when controlling for policy coordination. In this latter case, by the way, the size of the effect weakens, this confirming Gruben et al (2002) claim. Figure 4 presents weighted average of our benchmark estimates of the trade effect with those estimated with Frankel and Rose procedure.

Figure 4: Comparison between our benchmark results and Frankel and Rose methodology



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

The striking result concern the fact that benchmark IV estimates is 33% lower than that estimated with Frankel and Rose procedure. To show why this is the case, consider the analytic expression for the IV coefficient of β from equation [5].

$$\tilde{\beta}_{IV} = \frac{COV(y, TR\hat{A}DE)}{VAR(TR\hat{A}DE)} \qquad [a]$$

$T\hat{R}ADE$ being the vector of fitted values from the first stage regression, and y being our dependent variable. Assuming the true model is, instead, represented by [7] and assuming orthogonality between TRADE and the disturbance in equation [7], we can express [a] as:

$$\tilde{\beta}_{IV} = \beta \frac{COV(TRADE, T\hat{R}ADE)}{VAR(T\hat{R}ADE)} + \delta \frac{COV(FISC, T\hat{R}ADE)}{VAR(T\hat{R}ADE)} + \gamma \frac{COV(INT, T\hat{R}ADE)}{VAR(T\hat{R}ADE)} \qquad [b]$$

The sample correlation of TRÂDE and our policy variables is given in table 12. As can be seen, the sample moments seems to 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:

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

Table 12: Correlation between regressors and TRÂDE

	TRÂDE (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 produce macroeconomic stability problems when countries taking part to the arrangement are characterized by a low extent of cycle synchronization. This view has been criticized by Frankel and Rose, who have argued that this criterion should be considered as endogenous to the construction of the currency area: fixed exchange rate should, in fact, stimulate intra-union trade intensity thus contributing to more synchronous cycles and to a better functioning of the common monetary policy. In this paper it has been extended the econometric approach adopted by Frankel and Rose (1998) to support their hypothesis. Building on the considerations presented by Gruben, Koo and Millis (2002) we have explicitly taken into account the role of policies' convergence on real co-movement. Our empirical results suggest three main points:

- **i.** Coordination of policies seems to help cycle synchronization: countries that have got similar fiscal policy seem to co-move more than countries with idiosyncratic policies, this result being robust 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 intra-European co-movement of business cycles, especially when controlling for the presence of outliers. Further research should address the problem of endogeneity of policies in this framework before the robustness of our point estimates can be settled;
- **ii.** Bilateral trade intensity positively affects business cycles co-movement, this confirming Frankel and Rose's claim. In terms of the discussion in section 2, it implies that "specialization" induced by inter-industry trade is not so strong in Europe to overcome intra-industry trade and spill-over effects that positively influence cycle synchronization. Nonetheless, the size of this "trade effect" appears to be smaller than that estimated adopting Frankel and Rose's methodology (on average between 36% and 33%);
- **iii.** European data 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 tends 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.

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

As a first point, it can be supported the idea that the pattern of cycle co-movement between European countries will not change much trough the trade channel as a result of the adoption of the euro. Although we do not provide a formal proof of that, there are two clues concerning this outcome. 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. Micco, Ordonez and Stein (2003) have quantified this effect to be 8,9% while De Nardis and Vicarelli (2003) have estimated a positive impact of the euro of the order of 6%. Those results are far from the 65% estimated by Glick and Rose (2002) for other experiments of monetary unification (including developing countries as well): these discrepancies may be caused by the fact that Europe is already a very integrated area in the goods market, so that the adoption of the euro had a minor impact on intra-EMU trade. On the other hand, our estimation of the impact of trade on business cycle co-movement suggests that a move from the median to the 75th percentile in bilateral trade intensity (which correspond to an increase in bilateral trade intensity by 150%) increase (average) cycle's synchronization by only 15%, far from the 50% implied by Frankel and Rose estimation. Together, these two factors may let us conjecture that Frankel and Rose hypothesis is likely to operate in Europe only in the very long run.

Another implication of the analysis is that fiscal policy rule may be to some extent beneficial for the macroeconomic policy framework of EMU. Early studies on the *rationales* of the Maastricht treaty and of 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. We argue that fiscal policy rules may help also the stabilizing effort of the European Central Bank: to the extent that the SGP prevents idiosyncratic behaviour in the conduction of domestic fiscal policies, we should observe more synchronous cycles between European countries and thus a more efficient common monetary policy. This conclusion, nonetheless, need 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, especially for countries whose internal needs will not be represented by the ECB policy: the adoption of the single currency, the absence of a federal risk sharing arrangement (e.g. Farina and Tamborini, 2002) and the lack of intra-European labor mobility all point in the direction of national fiscal policies to be one of the few instruments in the hand of European countries to counteract idiosyncratic movements in real variables. In absence of a central political authority, the solution of this *trade-off* between "European rules" and "national discretion" represents one of the crucial point Europe should face to improve its short run macroeconomic policy frame and to increase the economic success of the euro.

Statistical Annex

Table 2: 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) and OECD Economic outlook	//	//
CICL2		Cross country correlation coefficient of cyclical GDP growth rate (Baxter and King band pass filtered)		//	//
FISC1		See section 4.1	OECD (Retrospective statistics, 2002 and OECD economic outlook	//	//
FISC2			Our elaboration on OECD (Retrospective statistics, 2002 and 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					
INT			AMECO database (EU-Commission)	Greece and Sweden (1972-1978)	Private consumption deflator in computing short term real rate of return
TRADE			ITCS database, OECD		Data on Austria are from IMF direction of Trade statistics
Instruments	Ling	Dummy variable that takes the value of 1 if the two countries share the same language	Frankel and Rose data-set		
	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 3: Summary statistics, by sub-period							
		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
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 25th, the median and the 75th percentile; cv is the coefficient of variation.

Table 4: Correlation among real co-movement (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 5: Correlation among real co-movement (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 6: 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 9: 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)	(0.18)	(0.21)	(0.15)
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	432	432	418	418	418

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.

Table 10: Estimation of equation [7] by sub-periods (CICL1)										
	No 1st subperiod		No 2nd subperiod		No 3rd subperiod		No 4th subperiod		No 5th subperiod	
	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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