Synchronisation of Business Cycles in the Enlarged European Union

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Uwe Böwer

Referent:	Professor Frank Westermann, PhD
Korreferent:	Professor Dr. Jarko Fidrmuc
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[1] Most of this chapter was produced in cooperation with Catherine Guillemineau at the European Central Bank.

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List of Abbreviations

ADF	Augmented Dickey-Fuller
AIC	Akaike Information Criterion
BIS	Bank of International Settlements
BK	Baxter-King
BTT	Bilateral Trade to Total Trade
BTY	Bilateral Trade to GDP
CD(X)	Codependence of Order X
CD-CHEM	Cross-Country Sectoral Differences, Chemicals
CD-CNT	Cross-Country Sectoral Differences, Construction
CD-FIN	Cross-Country Sectoral Differences, Financial Sector
CD-FUEL	Cross-Country Sectoral Differences, Fuels
CD-IND	Cross-Country Sectoral Differences, Industry
CD-MACH	Cross-Country Sectoral Differences, Machinery
CD-MANU	Cross-Country Sectoral Differences, Manufacturing
CD-TRA	Cross-Country Sectoral Differences, Wholesale and Retail Trade
CEECs	Central and Eastern European Countries
CYSERDIFF	Cyclical Services Difference
DEFDIFF	Deficit Differential
DF-GLS	Dickey-Fuller General Least Squares
EA	Euro Area
EBA	Extreme-Bounds Analysis
ECB	European Central Bank
ECOPAT	Economic Patterns
EMU	Economic and Monetary Union
EPADIFF	Employment Protection (Average) Differential
EU	European Union
GDP	Gross Domestic Product
GEODIST	Geographic Distance X
GMM	Generalised Method of Moments
HP	Hodrick-Prescott
IMF	International Monetary Fund
IP	Industrial Production
IRSDIFF	Differential of Short-Term Interest Rates
ISIC	International Standard Industrial Classification
LAD	Least Absolute Deviation
LBFA	Log Bank Flows, Assets
LBFL	Log Bank Flows, Liabilites
NCIDIFF	National Competitiveness Indicator Differential
NMS	New Member State

OCA	Optimum Currency Area
OECD	Organisation of Economic Cooperation and Development
OLS	Ordinary Least Squares
POPDIFF	Population Differential
SD-NERE	Standard Deviation of Nominal Exchange Rates
SECM	Seasonal Error Correction Model
SIC	Schwarz Information Criterion
SITC	Standard International Trade Classification
SVAR	Structural Vector Auto-Regression
TOTMKDIFF	Total Market Differences
TRADEPAT	Trade Patterns
TTY	Total Trade to GDP
TUDDIFF	Trade Union Density Differential
UK	United Kingdom
US	United States
VAR	Vector Auto-Regression

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Chapter 1

Introduction

The creation of the single European currency in 1999 is an unprecedented experiment in modern international finance. In 2004, the European Union admitted ten new member states and is in negotiation with additional countries to join at a later stage. The inclusion of these new and prospective members in the euro area will be one of the greatest EU challenges ahead.

This dissertation addresses a number of resulting fundamental policy questions of European monetary integration. Which of the new member states are ready to adopt the euro, in the sense that their business cycles show sufficient synchronisation with the euro area? What has been driving cycle synchronisation among the existing euro area countries and what can we infer about potential endogenous effects of the euro? What is the role of risk-sharing and financial integration in the context of monetary union, and may beneficial risk-sharing effects make up for lacking cycle synchronisation of the euro adopters?

In the following, we contextualises these questions by highlighting the political environment of monetary integration in Europe. We then discuss the concepts and results of the three major dissertation chapters, each presented in the light of the relevant theoretical and empirical literature.

1.1 The political context

This section provides background information on the historical and present challenges of European monetary integration. We briefly review the landmarks on the road to the euro and highlight the current political questions arising from EU enlargement and the new member states' transition towards the euro.

1.1.1 A short history of European monetary integration

First attempts towards monetary unification were made in the late 1960s when the international monetary environment started to show signs of instability. The 1970 Werner Report proposed a stepwise procedure to create an economic and monetary union by 1980. This included narrowing fluctuation margins of the European currencies which became known as the "snake". During the 1970s, the break-down of the Bretton Woods system of fixed exchange rates as well as the oil crises created substantial tension in the world economy. The divergent policy responses by the EU member states to these economic shocks slowed down the process of monetary integration.[1] It was only in 1979 that this process regained momentum with the establishment of the European Monetary System with its parity grid and the European Currency Unit (ECU) which served as basket currency. The idea of a single currency resurfaced after the Single European Act was launched in 1987. The common market, it was argued, would remain incomplete without a common currency. The Delors Report, published in 1989, became the blueprint for the chapters on Economic and Monetary Union (EMU) of the 1992 Maastricht Treaty on the European Union. The road to the euro was to be taken in three stages, the third of which involved the establishment of the European Central Bank and the introduction of the euro as a unit of account in 1999. Three years later, the changeover to the euro was completed when euro banknotes and coins were brought into circulation.

[1] The European Union emerged from the European Community of Coal and Steel, the European Atomic Community and the European Economic Community which became the European Community. For the sake of simplicity, we use the term "European Union" throughout.

Eleven EU member states qualified initially for the single currency while Greece joined the euro area in 2001.[2] With the enlargement of the EU by ten countries in 2004, the new member states entered the European System of Central Banks and are committed to prepare for eventual adoption of the euro.

1.1.2 Euro area enlargement

Having fulfilled the *acquis communautaire*, the new members are formally part of EMU already although being in the preparatory phase for the adoption of the euro. Since no "opt-out clauses" as for Denmark and the UK were granted during the accession talks, all new members are required to take the necessary steps towards full monetary integration as specified by the Maastricht convergence criteria. Besides convergence in inflation rates, budget and debt positions, the criteria include a compulsory two-year membership in the second generation of the exchange rate mechanism (ERM II) without major disturbance. Subsequently, the final stage of EMU involves introducing the euro as the sole legal tender. To date, six out of the ten new member states have entered ERM II already so that, provided that the remaining convergence criteria will be fulfilled, the euro area may be enlarged as soon as 2006.[3]

1.2 Literature review and summary of results

Given this political framework, the question of optimal timing arises from an economic point of view. What does economic theory say on currency union membership? What are the answers that empirical analysis can provide on the question of the optimal timing towards euro adoption? In the following, we highlight the major three strands in

[2] The eleven original euro area members were Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Spain and Portugal.

[3] Estonia, Lithuania, and Slovenia joined ERM II in June 2004, Cyprus, Latvia, and Malta followed in April 2005. As a matter of fact, Slovenia will join the euro area in 2007 whereas Lithuania's application was rejected by the European Commission, on the grounds of the excessive Lithuanian inflation rate.

currency union economics: classical optimum currency area theory, the endogeneity of optimum currency areas, and the role of financial markets and risk sharing in monetary unions. These strands correspond to the three subsequent chapters of analysis. In each of the following sub-sections we first elucidate the theoretical background and review the existing empirical research before we summarise the conceptual frameworks and the key results of the economic contributions that constitute the remainder of this dissertation.

1.2.1 The initial optimum currency area approach

This section goes back to the roots of currency union research and reviews the original ideas of the optimum currency area (OCA) approach from the 1960s. We then highlight the empirical evidence which builds on this basic framework which, eventually, brings us to the analysis of common trends and cycles of the new member states with the euro area.

The theoretical framework of Mundell I

The theory of optimum currency areas has been at the heart of currency union research. Although it is no fully-fledged theory, the initial OCA framework provides helpful guidelines for the investigation whether or not certain countries would be good candidates for a currency union.[4] We first describe the basic OCA ideas before we outline various attempts to model OCA theory more formally. This early OCA approach has been known as Mundell I, distinguishing it from Mundell's later work to which we come below.

In his seminal contribution, Mundell (1961) highlights that regions with similar economic characteristics may benefit from a common currency even if they do not belong to the same national federation, i.e. if "the national currency area does not coincide with the optimum currency area" (Mundell 1961: 657). The efficiency benefits of a common currency need, however, to be weighted against the costs of renouncing independent

[4] See Mongelli (2002) for a review on the extensive OCA literature.

monetary policy and exchange rate adjustments.

Over the subsequent years, a number of criteria have evolved which typically characterise an OCA. First, the flexibility and mobility of production factors is regarded as a key prerequisite, see Mundell (1961). If wages can adjust freely and capital or labour can re-allocate without restrictions, the need for exchange rate adjustments in response to economic disturbances is reduced. Second, the more open a country is to international trade, the more is the domestic economy influenced by international price changes. McKinnon (1963) argues that, hence, the scope of national monetary policy and exchange rate adjustments is naturally low. Third, Kenen (1969) suggests that a more diversified economy is favourable because it is less threatened by idiosyncratic shocks and hence not so much in need of domestic monetary or exchange rate response. Furthermore, interregional compensation schemes and the political will for integration have been cited as additional aspects of OCAs, see Krugman (1993) and Mongelli (2002).

Some attempts have been undertaken to formalise the Mundell I catalogue of verbal arguments and translate them into mathematical models. Bayoumi (1994) presents a micro-founded general equilibrium model of trade which integrates and compares the criteria of the early OCA framework. The two-country model assumes full specialisation, labour immobility and downward wage rigidity. In case of individual currencies, the nominal exchange rate adjusts for relative price changes due to asymmetric shocks. In a currency union, however, this adjustment mechanism is absent. Since prices and wages cannot decrease, the adversely affected country will suffer unemployment. The costs of the common currency depend on the size of the asymmetric shock and their correlation. Benefits arise in this model mainly from saved transaction costs. In consequence, the net benefit of a currency union with asymmetric shocks is greater the larger the transaction costs, the higher the trade volume, the smaller the asymmetric shocks and the larger the correlation of disturbances. Hence, the model integrates McKinnon's openness criterion as well as Kenen's argument on diversification - via the size and correlation of shocks.

Also, Mundell's initial labour mobility criterion is included, in that it would provide an alternative adjustment tool to alleviate regional unemployment in the first place.

Other currency union models analyse a common currency as a commitment device. This literature builds on models of credibility in monetary policy of the Kydland-Prescott/Barro-Gordon-type.[5] Typically, a loss function describes the goals of monetary policy, in that the central bank would minimise the deviations from output and inflation targets, in the presence of supply shocks. Then discretionary monetary policy is compared with various commitment designs under dynamic consistency aspects. By adopting a credible anchor currency, a client country can benefit by "importing" credibility and a reputation of stability. However, if the supply shocks of the client country deviate substantially from those of the anchor country's economy, the client may incur considerable costs. Alesina and Barro (2002) present a version of such a model and demonstrate that countries with a high-inflation record and closely correlated business cycles benefit most from currency union. In addition, they show that small countries with a large trade share have the hightest potential for reduction in transaction costs in a currency union.

Empirical evidence based on Mundell I

It has proved difficult to test the OCA criteria empirically in a systematic and consistent manner. For instance, labour market flexibility is notoriously difficult to quantify. Also, similarity indices of diversification or capital mobility tend to involve a significant degree of subjectivity. Instead, it has become customary to analyse the symmetry in the stochastic experience of countries' economic performance, i.e. the symmetry of shocks or the synchronisation of business cycles. This approach has been known as the "meta property" of OCA theory because most of the individual criteria translate into the probability of asymmetric shocks and the economy's ability to respond to these shocks, see

[5] See, for example, Obstfeld and Rogoff (1996) or Illing (1997).

Masson and Taylor (1993) and Mongelli (2002). For example, the more diversified the economic structure, the less likely is the occurance of idiosyncratic shocks in the first place. Moreover, if the countries are very trade-integrated, the probability of being hit by symmetric shocks tends to be larger. In case of mobile production factors and fiscal federalism, adjustments in these areas can cushion the adverse impacts of asymmetric shocks. Thus, the more symmetric the shocks, or the more synchronised the business cycle behaviour of two countries, the more likely it is that the major OCA criteria are satisfied.

Two alternative ways of measuring the stochastic experience stand out. One part of the literature attempts to measure the similarity of shocks directly. Based on the structural vector autoregressive (SVAR) approach of Blanchard and Quah (1989) these scholars distinguish demand and supply shocks by imposing the assumption that only supply shocks exert a permanent effect on output, while the long-term impact of demand shocks is restricted to zero. Bayoumi and Eichengreen (1993) apply this methodology to Western Europe. They argue that the more similar the incidence of shocks across countries, the better are the OCA criteria fulfilled and the more likely a country would benefit from currency union. Comparing European countries to US regions, they establish a core-periphery distinction and assert that only a few core EU countries would be suited for EMU.

Another branch of the literature adopts a more general approach and explores the observed comovement of short-run stochastic output behaviour, i.e. the synchronisation of business cycles. Mostly, real output data have been de-trended using the Hodrick-Prescott filter or the Baxter-King band-pass filter. The correlation coefficients of the resulting cyclical output components are then interpreted as synchronisation indicators across countries. Also, Markov-switching VARs have been employed to identify a common European cycle, see Artis et al. (2004). Furthermore, Engle and Kozicki (1993) formulate the common features approach which investigates business cycle synchronisation by identifying common serial correlation features on the basis of cointegration. Vahid and Engle (1997) develop the advanced codependence technique. Rubin and Thygesen (1996) apply an early version of the common features test and find some evidence of common cycles among Western European countries in the run-up to EMU.

A number of studies with a focus on various Central and Eastern European Countries (CEECs) and the euro area have been conducted recently. These papers typically employ either SVAR approaches or determine simple correlations of cyclical output components. Fidrmuc and Korhonen (2006) provide a comprehensive literature overview and perform a meta-analysis of business cycle correlation. They identify substantial differences between the studies reviewed and highlight the difficulties of empirical analysis in the context of transition countries regarding data availability and methodological validity. On the whole, their survey concludes that the cycles of several CEECs are relatively highly correlated with the euro area cycle, in particular those of Slovenia, Hungary, Poland and, to a lesser extent, Estonia.

However, little attention has so far been paid to the combined analysis of long-run trends and short-run cycles, as incorporated in the common features/codependence technique. While most of the reviewed studies adopt the SVAR technique, only Buch and Döpke (2000) apply the common features framework on the CEECs. They find little evidence of common cycles, which may, however, be due to the limited data period at the time the study was conducted.

Common trends and cycles of Central and Eastern Europe and the euro area

Chapter 2 of this dissertation tests the meta-property of the OCA theory for selected CEECs in relation to the euro area. We follow the approach of cycle synchronisation and not the strategy of shock incidence. The SVAR methodology of imposing identifying restrictions by labelling shocks "demand" and "supply" has not been undisputed.[6]

[6] Juselius (2004) and Rubin and Thygesen (1996) criticise the arbitrariness of imposing restrictions that are based on theoretical grounds instead of allowing the data to determine the model.

The strategy of applying zero restrictions for relatively short time series may lead to uncertainties in the results. Moreover, observing output comovements instead of shocks to catch the OCA "meta property" incorporates not only the external shocks themselves but also the capability of the economies to respond to such shocks.

Based on this rationale, our study makes two contributions. First, we apply an integrated cointegration/codependence approach to investigate long-run output comovement and short-run synchronisation of business cycles between eight CEECs and the euro area. The codependence technique does not only focus on the short-run properties of the time series but considers the comovement of both long-run trends and short-run cycles. To investigate the common cycle properties correctly, the results of the preceding cointegration analysis feed into the calculation. Second, the approach takes seasonal effects explicitly into account. Instead of applying up-front seasonal adjustment procedures, we resort to non-adjusted data and employ seasonal cointegration and codependence techniques which incorporate the seasonality into the statistical model. This strategy draws on the seasonal version of the codependence analysis based on Vahid and Engle (1997) and Cubadda (1999).

We analyse trend and cycle comovement successively. The trend analysis is the foundation of the subsequent cycle tests and estimates catching-up and steady-state convergence. A simple cross-section regression confirms the catching-up convergence hypothesis in that it indicates significantly higher average growth rates for those countries with lower initial income levels. Apparently, most CEECs under investigation are still in the process of transition towards the steady-state equilibrium. This assertion is confirmed by the cointegration analysis. Using quarterly, not seasonally adjusted industrial production data for the aggregate euro area plus the eight individual countries, we perform bivariate seasonal cointegration tests for each country vis-à-vis the euro area. We find no cointegration relations at standard frequency and conclude that the CEECs are still largely in transition towards the steady-state equilibrium. The negative cointegration result is consistent with the existence of catching-up convergence because the two convergence concepts can be regarded as mutually exclusive.

The tests for common business cycles divide into the categories of synchronised and non-synchronised common cycles. We find only one case of synchronised common cycles when testing for contemporaneous common serial correlation features, Slovenia. This indicates that the remaining countries do not share common cycles with the euro area. Regarding common but non-synchronised cycles, we test for codependence as we allow for a delay in the response to shocks. Due to inefficiencies in the propagation mechanism of shocks, the CEECs may respond to shocks but not in the initial period. In fact, we find evidence of first-order codependence for Hungary, Slovakia and, as a borderline case, the Czech Republic. Estonia shows signs of second-order codependence. These countries can, therefore, be considered as having an intermediate degree of cyclical comovement with the euro area. For Poland, as well as for the candidate countries Croatia and Turkey, we do not find any codependence. Their cycles do not even align to that of the euro area after a certain delay.

In sum, these results seem to suggest that real integration of the CEECs with the euro area is still limited. In the framework of Mundell I, only Slovenia appears well-equipped for joining the euro soon. For most of the other CEECs, however, giving up individual monetary policies at a too early stage may entail the risk of incurring major costs. For these countries, there is still some way to go to achieve business cycle synchronisation.

1.2.2 Endogeneity of optimum currency areas

This section examines the second major strand in currency union analysis dealing with the potential endogenous effects which a common currency may unfold. It was chiefly the seminal contribution of Frankel and Rose (1998) which enkindled a number of studies in this area. This sub-section investigates the theoretical rationale, outlines the empirical evidence and summarises the related study of Chapter 3.

Theoretical models of OCA endogeneity

The endogeneity argument is as follows. While the traditional OCA criteria are formulated as prerequisites for currency union, they may in fact evolve as a very consequence of the introduction of a common currency. The currency union itself may increase trade and synchronise business cycles so that, even if a country group would not have qualified as an OCA *ex ante*, it may turn into an OCA *ex post*.

A major discussion has revolved around the question whether or not increased trade would in fact lead to more closely synchronised business cycles. The European Commission (1990) expects more trade to have a positive influence on cycle comovement. This view would be substantiated if economic shocks were predominantly demand-driven and hence spill over more easily across countries via the trade channel. Also, a large degree of intra-industry trade would suggest the rising importance of common shocks, as opposed to idiosyncratic shocks. Krugman (1993), however, suggests that a rise in trade would facilitate industry specialisation across countries and hence trade would become increasingly inter-industry. In this case, we would expect the synchronisation of business cycles to go down as a result of currency union.[7]

The endogeneity debate has produced a few formal models. Ricci (2006) develops a monetary model with firm-location choice and finds that irrevocably fixed exchange rates can reduce shock asymmetry. This is mainly due to the firms' assumed preferences for exchange rate stability, so that, in case of floating rates, firms of a certain industry would agglomerate in the country where macro shocks coincide with the shocks faced by the specific industry of the firm. In the presence of a currency area, variability-adverse firms can therefore afford to spread more evenly across countries. As a result, diversification and intra-industry increases in a currency union, generating more synchronous cycles.

[7] While Krugman (1993) argues that the euro area may follow the U.S. example of increased specialisation, Clark and van Wincoop (2001: 71) point out that U.S. census regions have actually become *less* specialised during the post-war period and that, in 1986, the degree of specialisation in the U.S. and in Europe were about the same.

Corsetti and Pesenti (2002) present a model of "self-validating optimum currency areas" which is independent of real integration and intra-industry trade effects. Allowing for imperfect pass-through of exchange rate onto export prices, they show that, in the presence of a credible commitment to currency union, an equilibrium may emerge in which firms preset prices not in domestic but in consumer currency. In this case of zero pass-through, monetary policies are symmetric across countries, so there is no cost of giving up monetary sovereignty and the currency union becomes self-validating. As a result, output correlation is higher under currency union than under the alternative, floating regime where monetary policy reacts to shocks.

Endogeneity empirics

The endogeneity debate started with a series of empirical papers by Andrew Rose and co-authors in the late 1990s. Frankel and Rose (1998) refute Krugman's (1993) proposition and find a positive effect of trade on business cycle synchronisation. They interprete their result as an indication of the endogeneity of optimum currency areas. Rose (2000) conducts a gravity analysis, regressing bilateral trade on relative country size, geographical distance and numerous control variables. To isolate the effect of a common currency on trade, he introduces a currency union dummy. His results claim that the mere fact of having a common currency is associated with trade volumes higher by a factor of up to three, in relation to those countries that were not part of the same currency union. Engel and Rose (2002) extend the analysis and find a significantly positive effect of currency unions on the correlation of business cycles. Frankel and Rose (2002) link the currency union effect on trade to an increase in output.

This approach has evoked much criticism. It has been noted, for example, that the "Rose effect" of currency unions could only be substantiated when using a vast dataset of diverse countries which involves tiny island states and in which many currency unions are in fact overseas dependencies. Only the East Carribean Currency Area stands out as a currency union of today's understanding wheras EMU is not considered.

Baxter and Kouparitsas (2004) and Imbs (2004) analyse large samples of both developing and industrialised countries and find trade flows, specialisation and financial integration to be important factors for business cycle synchronisation. Their results are, however, not unequivocal and seem to depend on the country samples and time periods chosen.

In the wake of the studies pursued by Rose and co-authors, several scholars undertook the attempt to extend the analysis on the euro area. For instance, Micco et al. (2003) replicate the Rose-type regression approach, employing the most recent data and incorporating various EU-specific variables to control for other aspects of European integration. They tend to find moderate trade efffects of the euro but they have to conclude that it may still be too early to detect a significant and robust effect.

Analysing the degree of business cycle synchronisation over time, various studies indicate increasing synchronicity as monetary integration in Europe intensified, see for example Artis and Zhang (1997, 1999) or Massmann and Mitchell (2004). Applying Markov Switching VAR models, Artis and al. (2004) find evidence of a distinct European business cycle.

Determinants of business cycle synchronisation across euro area countries

The contribution of Chapter 3 of this dissertation deals with the underlying factors of cycle synchronisation in the euro area and hence addresses the endogeneity argument for those countries which have adopted the euro already. We do not try to calculate the currency union effect of EMU explicitly because we believe the results of previous attempts have been rather tentative due to the short time period and problems in truely isolating the effect of the euro from other influences. Rather, we ask which factors are significantly, and robustly, associated with business cycle synchronisation across euro area countries.

We investigate a number of potential determinants of cycle synchronisation in the context of European monetary integration. Our intention is to find out why, inside the euro area, the business cycles of some country pairs are more synchronised than others and whether the importance of these mechanisms have increased or declined over time. We test some standard determinants and, in addition, consider a number of EMU-specific policy and structural indicators which, to our knowledge, have not been tested in this context. We check robustness by applying the extreme-bounds analysis framework as suggested by Leamer (1983) and further developed by Levine and Renelt (1992) and by Sala-i-Martin (1997). Also, we divide our 25-year sample period into sub-samples in order to capture changing effects throughout the different stages of European integration.

Our main results are as follows. We find that bilateral trade have indeed been a robust, positive determinant of business cycle synchronisation. Hence, we see the endogeneity hypothesis of Frankel and Rose (1998) confirmed for the euro area: countries with larger trade volumes tend to have more closely synchronised business cycles. Although we observe this phenomenon over the whole sample, its explanatory power seems to be driven mainly by the earlier sub-sample, 1980-1996. During the period of preparation for EMU and actual currency union, since 1997, we find that the differences in trade structure emerge as robust determinants of cycle synchronisation. In other words, the degree of intra-industry trade plays an increasingly important role in binding euro area business cycles together. In combination with our descriptive finding of rising intra-industry trade among euro area countries, this result seems to point at the potential *ex post* optimality of the euro area.

Regarding our policy and structural indicators, fiscal deficit differentials appear to have driven differences between business cycles until the preparation for EMU. With the implementation of the Stability and Growth Pact, fiscal policy seems to have become less pro-active and fiscal deficit differentials have lost some of their explanatory power. In contrast, similarities in monetary policies, measured by interest rate differentials, have emerged as a robust determinant of business cycle synchronisation. Also, differences in the size of industrial sectors, stock market comovement and similar competitiveness situations appear to have good explanatory power. On the other hand, we could not detect any robust impact of nominal exchange rate variability, bilateral bank capital flows or differences in labour market flexibility on cycle synchronisation. The missing effect of mere exchange rate stabilisation on the synchronisation of business cycles is in line with the endogeneity hypothesis which predicts that only irrevocably fixed exchange rates, i.e. currency union, would unleash synchronisation dynamics.

Taken together, these findings seem to support Frankel and Rose's prediction that EMU would go hand in hand with trade expansion and the development of intra-industry trade which in turn would result in more highly correlated business cycles. Although more time is needed to make definite statements on the impact of the euro, we are cautiously optimistic on the endongeneity of a European OCA.

1.2.3 Mundell II: Risk sharing, financial integration and the insurance mechanism of currency areas

Although the seminal Mundell (1961) paper and its follow-ups have been the starting point for most currency union researchers, Mundell delivered a second influential contribution which has, however, received only recent attention. This 1973 article, called "Uncommon arguments for common currencies", has been known as Mundell II. By adding the role of financial markets and risk sharing to the OCA debate, Mundell (1973) specifies and revises some of his initial arguments on business cycle synchronisation, generating interesting implications for the analysis of EMU.

The idea of Mundell II and theoretical contributions

It was McKinnon (2002) who drew attention to the seminal Mundell II paper, Mundell (1973). The classical framework of Mundell I concentrates on the potential *costs* of currency union incurred by the loss of independent monetary policy and nominal exchange rate adjustments and asserts the importance of economic similarity, notably in terms of business cycles, trade openness, diversification and labour mobility. Mundell II, in contrast, revises this cost argument and turns the attention more towards the *benefits* of a common currency.

Regarding the cost of currency union, Mundell II argues that national monetary policies may not be as effective an adjustment tool to asymmetric shocks as the Keynesian beliefs of the 1960s would have suggested. This period was shaped by the static Mundell-Fleming framework of the open economy with its assumption of stationary expectations regarding prices, interest rates and exchange rates. Also, the Brettow Woods system of fixed exchange rates was functioning reasonably well and most countries had captial controls in place. These circumstances of what has been called the "fine-tuning fallacy"[8] led Mundell I to emphasise the costs associated with the loss of renouncing inidividual monetary policy - over-emphasise, in the eyes of Mundell II.

Moreover, Mundell II no longer considers exchange rates to be an adjustment mechanism only but, to a substantial degree, a source of shocks in itself. In a world with little capital controls, McKinnon (2002) argues, exchange rate movements "are likely to be erratic at best" so that the notion of smooth adjustment under flexible exchange rates, one of Mundell I's key assumptions, turns out to be an illusion. Both aspects, the reduced effectiveness of national monetary policy and the ambiguous role of exchange rate, downsize the role of the costs of currency union as they were pointed out by Mundell I.

The third and probably most interesting point of Mundell II, however, refers to the benefits of currency union due to enhanced risk sharing. In a currency union, financial market integration may develop into a powerful risk-sharing mechanism by providing income insurance across the union. Due to portfolio diversification, adverse shocks to one country can be shared across the union. Trade and financial integration may act as

[8] Buiter (1999: 49).

income insurance since individuals across countries hold claims on each other's output. As a result of this ownership diversification, consumption streams become smoother and more highly correlated across countries, even and particularly in the presence of idiosycratic shocks to production.

Alternatively, imagine a positive productivity shock in one country. Under separate currencies, GDP and consumption rise by the same amount and falling prices lead to increased real balances. With a common currency, however, the union-wide price level goes down less than proportionally to the productivity shock in the respective country. To increase real balance holdings, that country could run a balance of payments surplus, for instance through trade in nominal bonds. The increase in consumption is less than the rise in GDP so that the other countries of the union participate in the positive shock by enjoying higher consumption as well.[9]

While financially integrated countries make good candidates for currency union against this background, Mundell II suggests that a common currency can be expected to deliver risk sharing benefits even for countries with hitherto little financial integration. Exchange rate uncertainty and interest rate risk premia inhibit international portfolio diversification and constitute a major reason behind the home bias puzzle in international finance. A common currency, it is argued, would convince financial intermediaries to diversify their portfolios so that the currency union in itself may develop into a boost for financial market integration.

Against this background, Mundell II challenges a central argument of Mundel I. While the initial OCA framework warns countries with asynchronous business cycles about joining a currency union due to the resulting loss of national monetary policy and exchange rate adjustments, Mundell II suggests that it is exactly those countries with asymmetric shocks which may benefit most from adopting a common currency and the resulting risk-sharing and income insurance mechanism. In other words, a country

[9] See Ching and Devereux (2003).

that considers joining a currency union, such as the new EU member states, may not want to wait until business cycles are perfectly synchronised but rather benefit from the insurance mechanism of a financially-integrated currency union as long as cycles are asynchronous.

The proponents of Mundell II apply a similar logic to the ex post experience in a currency union. While Krugman (1993) predicts problems for EMU due to increased specialisation in a currency union, McKinnon (2002) holds that the case for a common currency grows even stronger as the union members become more specialised and concludes that "the productivity gain from greater regional specialisation is one of the major benefits of having and economic cum monetary union in the first place." (McKinnon 2002: 217)

Buildung on Mundell II, Ching and Devereux (2003) develop a general-equilibrium model to examine the cost and benefit of currency union. They incorporate both Mundell arguments by allowing for the costs of a common currency due to losing the adjustment property of the exchange rate (Mundell I) as well as the benefits arising from consumption risk-sharing in a currency union (Mundell II). By taking both effects into account, the presence of asymmetric shocks does not automatically make flexible exchange rates more desirable, in contrast to what much of the empirical literature has been suggesting. If a country can benefit from the risk-sharing mechanisms of a currency union to a large degree, the presence of shock asymmetry may make the common currency more and not less attractive. Ultimately, the authors find that net losses from adopting a common currency are more likely the more dominant are nominal rigidities. However, according to their model, the welfare difference tends to be small. In consequence, shock asymmetry can be used as argument both in favour and against currency union, depending on the relative importance of the exchange rate adjustment and the role of risk-sharing in the face of nominal rigidities.

Empirical evidence on Mundell II

Empirical research on the arguments of Mundell II remains fragmentary. Farrant and Peersman (2005) analyse whether the exchange rate is a shock absorber or a source of shocks in itself. Althoug they do not consider the context of currency union but focus on the U.S. dollar exchange rate vis-à-vis a number of other currencies, we may suspect similar effects between the euro area and potential euro adopters. The authors employ an SVAR approach using sign restrictions instead of the traditional zero restrictions. In a subsequent variance-decomposition exercise they find that nominal shocks exert a considerable permanent effect on variations in the nominal exchange rate. They interprete their result as a strong indication that the exchange rate does not only act as shock absorber but also gives rise to shocks. These disturbances, they argue, could be reduced by joining a currency union.

A number of empirical studies have been conducted on the areas of financial integration and risk sharing although rarely linked explicitly to the Mundell II argument. Genereally, financial integration and risk sharing are notoriously difficult to measure. Baele et al. (2004) provide a survey on price-based and quantity-based indicators of financial integration. Price-based indicators rely on the idea of purchasing power parity (PPP) and imply converging interest rates across countries. Quantity-based measures include cross-country capital flows although data on bilateral flows tend to be scarce. In a series of papers, Lane and Milesi-Feretti (2001, 2005) analyse the dynamics of international financial integration on the basis of foreign assets and liabilities. Their findings suggest an increasing degree of financial integration among a selection of industrialised countries. Another quantity-based indicator is captured by the consumption-correlation puzzle which is one of the "Six major puzzles in international macroeconomics" as pointed out by Obstfeld and Rogoff (2000). A large degree of financial integration should be reflected, it is argued, by the correlation of private consumption across countries because consumers can smooth their consumption flows by benefiting from the risk-sharing effect of international portfolio diversification. Notably, consumption correlation would, from a theoretical viewpoint, be exptected to exceed output correlation. However, poor empirical evidence on consumption correlation has been puzzling. For instance, Darvas and Szapáry (2005) find that consumption correlation among the European Union countries remains below GDP correlation. Demyanyk and Volosovych (2005) come to similar results, applying the utility-based risk-sharing model by Kalemli-Ozcan et al (2001). They interprete their results along the lines of Mundell II, arguing that those countries with little risk sharing, namely the Czech Republic, Slovakia and the Baltic states, would be expected to reap the largest potential gain from joining monetary union.

In a next step, we would be interested in the interaction of financial integration, risk sharing and business cycle synchronisation in the context of currency union. Although many studies do not make explicit reference to currency union, they do touch on related topics. Kalemli-Ozcan et al. (2003) argue that financially integrated regions can afford to exploit increasing returns to scale by specialisation because capital markets make up for the insurance function otherwise exerted by geographical diversification. In an empirical excercise, they find evidence for their hypothesis that regions with well-integrated financial markets, such as U.S. states, tend to be more specialised than European countries. This is interpreted as supporting the Krugman (1993) argument which predicts increasingly asynchronous business cycles due to integration-induced specialisation. Imbs (2004, 2006), on the other hand, finds a positive impact of financial integration on cycle synchronisation. He employs various financial integration indicators in a simultaneous equations model and argues that a direct spill-over channel from financial integration to cycle sychronisation prevails over potential indirect effects via specialisation. But none of these studies considers the beneficial impact of risk-sharing via consumption insurance which may, according to Mundell II, compensate the adverse effects of asynchronous cycles.

As for other potential endogenous effects of currency union, more time is needed to make reliable statements about the impact of the euro on financial integration. First indications are, however, encouraging. Cappiello et al. (2005) find evidence on a positive effect of the euro on capital markets. On the micro levels, conditional correlations between euro area equity returns tend to move up at around 1999 and the volatility of bond markets has been reduced. Concerning macro aspects, the variability of yield premia has decreased with EMU, related to a reduction in macroeconomic volatility. Hence, the unfolding impact of currency union on financial integration seems to lend support to parts of Mundell II.

Risk sharing, financial integration and Mundell II in the enlarged European Union

Chapter 4 of this dissertation takes the Mundell II framework to the European data. From the codependence analysis of Chapter 2 we know that the degree of business cycle synchronisation between the CEECs and the euro area is still poor. Following the logic of Mundell II, this asymmetry may not be a reason against early euro adoption but rather highlight the potential gain for the prospective entrants, given that countries with asymmetric shocks typically benefit most from risk-sharing in a currency union.

We explore the past degree and future potential of risk sharing and financial integration in the context of euro area enlargement. Considering eight new member states (NMS) vis-à-vis the aggregate euro area, we investigate risk sharing by looking at correlation and codependence measures of private consumption and GDP. For comparison, we conduct the same analysis for the "old" EU member states. In a second step, we examine real interest rate comovement measures to proxy the degree of financial integration. Again, we pair the NMS with the euro area and compare their development with the experience of the EU-15 countries in the 1980s and 1990s vis-à-vis Germany in preparation to EMU. We employ correlation, dispersion and variability measures as well as the codependence technique. Taken together, we draw a threefold conclusion from our analysis.

First, we find that consumption correlations of the NMS with the euro area tend to be lower than GDP correlations. This result is confirmed by the codependence analyis and in line with the consumption correlation puzzle. One reason behind this result may be the relatively low degree of financial integration. Both correlation and codependence measures of real interest rate comovement between the NMS and the euro area indicate low values over the past decade. According to Mundell II, these countries would enjoy the largest potential gain from euro adoption. Second, the although GDP correlation still exceeds consumption correlation for the EU-15 countries, they are both at much higher levels and with a more narrow gap than those of the NMS. Also, financial integration as increased markedly in the run-up to EMU. In the face of the long common history of economic integration, we may expect a similar pattern for the NMS as they further integrate with the EU economy. Third, both consumption and GDP correlations tend to increase over time, for the NMS as well as for the EU-15 countries. We note that GDP correlations tend to rise even faster than consumption correlations. Also, interest rate comovement goes up as time proceeds. These observations seem in line with the hypothesis of Imbs (2006) who finds that financial integration does not only increase consumption correlation but also, at an even faster rate, output comovement. He argues that the consumption correlation puzzle may not originate in too little risk sharing in the first place but is rather due to the often neglected effect on output synchronisation.

The question that remains open at present is whether the introduction of the euro will speed up financial integration with the CEECs. If that will be the case, and we see indications for such an effect in the existing euro area, Mundell II would eventually prove correct. In combination with the OCA endogeneity argument, this would be good news in a twofold sense for the new EU member states. First, the euro may result in consumption and income insurance based on a risk-sharing effect if financial markets integrate quickly after joining the euro area. This effect would make up for some of the present shock asymmetry. Second, business cycles may synchronise endogenously, following further integration in trade and financial markets. Given the limited data situation to date, further research is required to shed more light on the effects of Mundell I and Mundell II on the enlarged euro area but indications so far seem to imply cautious optimism.

Chapter 2

Common trends and cycles of Central and Eastern Europe and the euro area

Since the enlargement of the European Union by ten countries in May 2004, most new members have expressed the goal to adopt the euro in due time. Does EU accession imply the end of the transition phase, or is more real integration required to pave the way to the euro? Given the compulsory two-year membership in the second generation of the exchange rate mechanism, the most advanced new members may enter the euro area soon, among them Slovenia which will introduce the euro in January 2007. In addition, more countries are expected to join the EU in the near future, implying their adoption of the euro at a later stage as well. For the former planned economies, preparing for the final stage of EMU has been a demanding task. Therefore, euro area enlargement is a prevailing policy question both for the existing Union and for the entrants, and demands new answers from empirical economics. In response, this study investigates trend and cycle comovement of six new EU member states as well as Croatia and Turkey with the euro area, employing seasonal cointegration and codependence approaches.[1]

We analyse trend and cycle comovement as follows. The first part constitutes the

[1] The new EU members considered are the Czech Republic, Estonia, Hungary, Poland, Slovakia, and Slovenia. The remaining four new member states Cyprus, Latvia, Lithuania, and Malta, as well as Bulgaria and Romania, were not included due to data constraints.

trend analysis and estimates alternative measures of convergence between the CEECs and the euro area. Following Bernard and Durlauf (1996), we distinguish and test two concepts: catching-up and steady-state convergence. The former is also known as beta convergence and investigates the catching-up process of countries in transition to a new steady state. We find preliminary evidence that since 1994, lower initial income levels led to significantly higher economic growth across European countries. In other words, the catching-up economies seem to have experienced a decade of transition towards a new steady state. The alternative understanding of convergence applies to those countries that have reached a steady state already. In this case, convergence is the process of mean-reversion to the steady state level after a shock to the system. We employ the seasonal version of cointegration analysis to avoid spurious results that may arise from using up-front seasonal adjustment. In bivariate seasonal cointegration tests, we examine output series of six new EU member states plus Croatia and Turkey against the euro area and find no cointegration relations at frequencies zero or 1/4. At frequency 1/2, we find Croatia and the Czech Republic to be cointegrated with the euro area. Hence, the CEECs are still largely in transition towards the steady-state equilibrium.

The second part of the paper deals with cyclical comovement. Distinguishing between synchronised and non-synchronised common cycles, we use the common feature and codependence approaches. The former concept was developed by Engle and Kozicki (1993) and serves to detect contemporaneous comovement among business cycles, using difference-stationary series. However, given that the common feature analysis is a measure of simultaneous comovement, it does not capture any delays in response by the other series. To discover cycles that are common but not synchronised, we use the codependence approach of Vahid and Engle (1997). In this case, the common response to a shock may not materialise in the first period but at some later stage. In other words, the codependence framework tests for common features in the qth period, allowing for different initial responses to a given shock. Building on the seasonal cointegration framework, the common feature/codependence analysis takes different seasonal frequencies into account. We find that only Slovenia reveals a common serial correlation feature while Hungary, Slovakia, Estonia and, as a borderline case, the Czech Republic exhibit signs of collinear cycles after one or two periods. For Poland, Croatia and Turkey, we find no evidence of codependence of any order.

The remainder of this chapter is structured as follows. Section 2.1 divides the trend analysis into catching-up and steady-state convergence and performs tests of these concepts. Section 2.2 focuses on common cycles and conducts common feature and codependence tests. Section 2.3 concludes.

2.1 Trend analysis

This section deals with the long-run convergence among the CEECs and the euro area. The understanding of convergence is, however, not clear-cut. Following Bernard and Durlauf (1996), we distinguish between two, mutually exclusive concepts of convergence: catching-up and steady-state convergence. Both concepts will be applied to the countries under consideration, using cross-section regression and seasonal cointegration analysis.

2.1.1 Catching-up convergence

The concept of catching-up convergence stems from the well-known convergence hypothesis of the neoclassical growth literature. A Solow-type production function with non-increasing returns to scale typically implies that the long-term behaviour of the economy will be independent of the initial conditions. Due to the concavity of the production function in the capital stock, capital-poor countries will grow sufficiently faster, i.e. catch up to the capital-rich countries to offset the initial differences. Hence, we would expect to find catching-up convergence primarily among developing and transition countries that are converging towards a steady-state which they have not yet reached.

The data in this sub-section are mostly taken from Eurostat and the World Bank's World Development Indicators. Data on years of schooling are extracted from the Barro and Lee (2000) database. The individual variables are explained below.

To obtain a preliminary indication of whether catching-up convergence exists in Europe, we test for beta convergence in a simple cross-section setting. Figure 2.1 plots the average 1994 to 2004 annual growth rates of real GDP against the logs of the 1994 initial real per-capita GDP levels of 25 European countries, i.e. 14 EU countries (except Luxembourg) plus eleven CEECs. For catching-up convergence, we expect lower initial incomes to be associated with higher growth rates. Graphical inspection suggests an overall negative relationship and divides the countries into two broad categories.

EU-14 countries are characterised by high 1994 income levels and lower growth rates whereas most of the Central and Eastern European states are located in the high-growth/low-initial-income area of the graph. Only two countries do not match this categorisation. First, Slovenia seems to be located closer to the EU-14 group than to the remaining CEECs. Second, Ireland stands out with its high initial income and high average growth rate. On the whole, most CEECs seem to be catching up to Western European income standards.

Our OLS regression analysis is a simplified version of Barro and Sala-i-Martin (1995) and includes the following control variables.[2] First, *school* measures the years of schooling in 1995 to proxy education attainment. Second, *edu* stands for average public spending on education, as a percentage of GDP. Naturally, this variable captures the quality of education beyond the mere years of schooling. Both variables are expected to raise the average growth rate. Third, the variable *invest* represents average gross domestic investment as a percentage of GDP. Since higher investment values increase output per

[2] There is little consensus in the empirical growth literature with regard to the choice of appropriate control variables. Here, we do not enter this discussion but simply adopt the approach of Barro and Sala-i-Martin (1995), with one modification: While they employ pre-sample values of most variables as instruments to avoid endogeneity problems, our data for Central and Eastern Europe does not naturally allow this approach. In particular, pre-1994 values may bias the results since they involve enormous variation due to the breakdown of the command economy systems.

effective worker, the growth rate tends to increase as well. Finally, government consumption is controlled for by the variable *gov*. It is measured as average general government final consumption expenditure as a percentage of GDP. Assuming that higher government consumption tends to distort private decisions, we expect a negative impact on the growth rate.

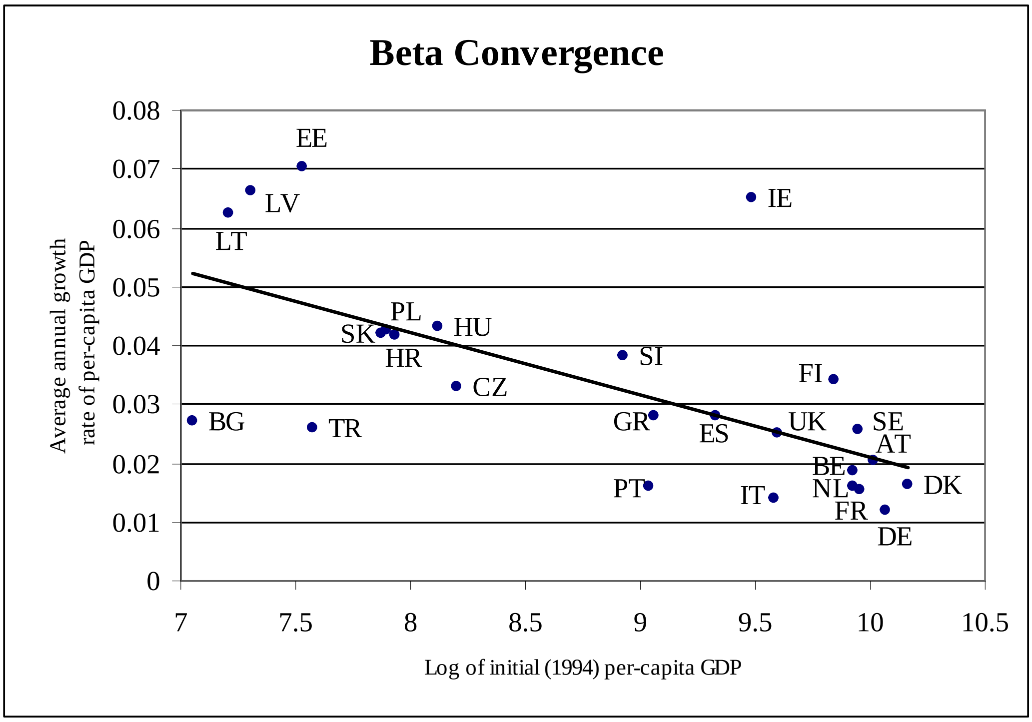


Figure 2.1: Beta convergence: cross-section regression of the avergage 1994-2004 annual per-capita real GDP growth rate of 25 European countries on the logs of the initial (1994) per-capita real GDP levels.

The regression equation can be expressed as

$$g_i = c + \beta x_{i94} + \gamma_1 school_i + \gamma_2 edu_1 + \gamma_3 invest_i + \gamma_4 gov_i + \varepsilon_i. \quad (2.1)$$

We regress g_i, the average annual GDP growth rate between 1994 and 2004 for country $i = 1, ..., 25$, on a constant c, initial per-capita GDP x_{i94}, the control variables as specified above, and an error term ε_i. The coefficient β measures the convergence effect and is expected to be negative.[3]

Table 2.1: Beta convergence

Estimation	(1)	(2)	(3)	(4)
GDP 94	-.0106	-.0105	-.0100	-.0114
	(-4.07)	(-3.81)	(-3.49)	(-4.26)
school		.0022	.0035	
		(1.21)	(1.98)	
edu		.4829		.4984
		(1.67)		(2.28)
invest		.0635	.0820	.0415
		(.89)	(1.11)	(.60)
gov		-.0673	.0025	
		(-.84)	(.03)	
adj. R^2	.39	.48	.44	.48

Note: OLS regression of the average 1994 - 2004 growth rate of per-capita GDP on the log of initial 1994 real per-capita GDP levels, with t-statistics given in parentheses. Constant terms are included but not reported.

The regression results are summarised in table 2.1. Estimation (1) includes only initial income as a determinant of the average growth rate. As expected, the relation is negative and, with a t-statistic of -4.07, clearly significant. *Ceteris paribus*, convergence occurs at the rate of around 1.0 percent per year. Adding the control variables in various combinations hardly affects the coefficient size or t-statistic of the initial income variable. In estimation (2), we include all four controls but none turns out to be significant. This may be due to multicollinearity. Indeed, *edu* is strongly correlated with *school* and with *gov*. Hence, estimation (3) excludes *edu*, while regression (4) is estimated without *school* and *gov*. The results imply a significantly positive effect of education on the growth rate, if only one of the two respective variables is involved. The education quality variable *edu*,

[3] In this indicative exercise, we employ a larger country sample than in the following time-series calculations, for the simple reason that more data are available on an annual basis as compared to the quarterly case. We doublecheck the beta regressions with an alternative, smaller dataset which matches the time-series country sample and find very similar results.

however, appears to have a far larger effect in size than the mere years of schooling. On the other hand, government consumption and investment do not seem to play a major role.

On the whole, we conclude from our simple cross-section framework that catching-up convergence seems to take place and that, besides initial income levels, educational attainments have played a major role. The following sections will investigate the other dimension of convergence more formally.

2.1.2 Steady-state convergence

The second convergence concept we consider is steady-state convergence. This case typically deals with those countries that have reached a common steady state already and move together in the long-run. Convergence denotes the process of mean reversion back to the steady state after the occurrence of a shock to the system. Hence, there is no permanent deviation from the common long-run equilibrium. In the words of Bernard and Durlauf (1996: 165), this process means that "the long-run forecasts of output differences tend to zero as the forecasting horizon increases." Analytically, we test for steady-state convergence by means of cointegration analysis. To detect long-run comovement between two or more non-stationary series, we try to find a linear combination which is then stationary. The cointegrated series follow a common stochastic trend and hence are steady-state converging.

Preliminary analysis

Before conducting the seasonal cointegration tests, we examine some descriptive properties of the series. The data used in this section comprises non-seasonally adjusted quarterly industrial production (IP) index series as a proxy of real economic activity.[4]

[4] Alternatively, we checked quarterly real GDP data which pointed at roughly similar results. However, the econometric analysis using GDP was complicated by the fact that the data were available for only few countries from 1994 onwards and that, in many cases, the apparently more volatile GDP series were

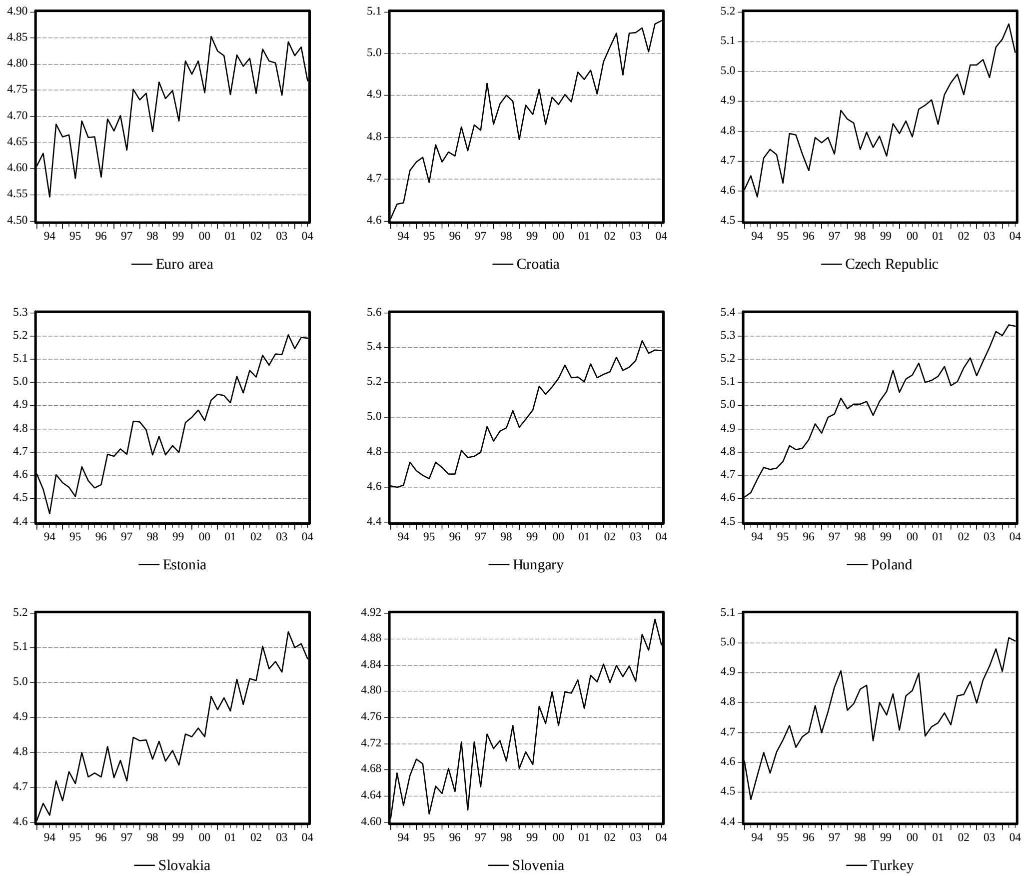


Figure 2.2: Levels of the logs of quarterly, non-seasonally adjusted industrial production indices, 1994Q1-2004Q3.

Nine economies are included in this analysis: the aggregate euro area as well as Croatia, the Czech Republic, Estonia, Hungary, Poland, Slovakia, Slovenia, and Turkey. The series are provided by the OECD's Main Economic Indicators, the International Financial Statistics of the IMF and by national sources. The sample covers the period 1994Q1 to 2004Q3; the starting point was chosen as to exclude the major downturn after the breakdown of the centrally-planned systems between 1990 and 1993.

only borderline difference-stationary. In chapter 4, we use an updated GDP dataset starting in 1995 as measure of comparison for consumption comovement in the framework of risk-sharing analysis.

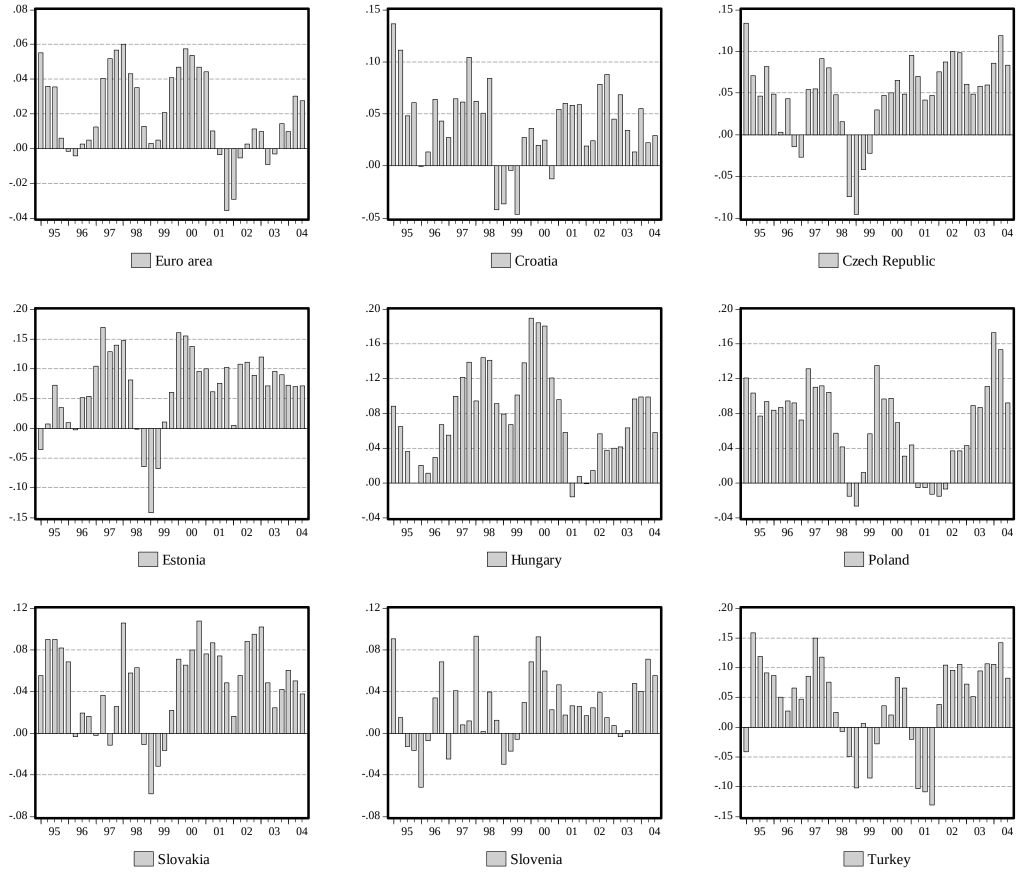


Figure 2.3: Seasonal differences of the logs of quarterly, non-seasonally adjusted industrial production indices, 1995Q1-2004Q3.

Figures 2.2 and 2.3 illustrate the series in levels, $x_{i,t}$, and in seasonal differences, $\Delta_4 x_{it} = x_{i,t} - x_{i,t-4}$. The line graphs of the levels reveal upward-sloping curves with clear seasonal patterns. While the IP indices of the euro area and most of the CEECs exhibit a relatively stable positive trend, only Turkey stands out. Its 2001 financial crisis is reflected in a downturn in output which reaches its 1997 level again only in 2002.

Graphed in seasonal differences, it becomes clear that many countries suffered from temporary set-backs. The bar graphs of Croatia, the Czech Republic, Estonia, Poland and Slovakia show negative spikes around 1998/1999 when the Russian financial crisis unfolded. Slovenia seems to have been less affected while Hungary apparently follows the post-2000 recession which is visible in the euro area series.

To analyse the data more systematically, we apply the Box-Jenkins techniques and test for unit roots. Analysing levels and seasonal differences subsequently, we first inspect the autocorrelation and partial autocorrelation functions. Slowly decaying autocorrelation in combination with abruptly diminishing partial autocorrelation suggests an autoregressive pattern in the data. Minimising the Akaike (AIC) and the Schwartz (SIC) information criteria serves to determine the appropriate lag lengths. However, two restrictions apply. First, the AIC is known to overstate the correct lag length while the SIC typically understates it. Second, the relatively small number of observations may distort the information criteria results. Hence, we examine the autocorrelation properties of the autoregressions' error terms of various lag choices and the related Q statistics.

Table 2.2: ADF test results

Country	Levels		Seasonal differences	
	ADF statistic	Lag	ADF statistic	Lag
Euro area	-1.50	5	-3.21**	3
Croatia	-3.66**	1	-3.58**	3
Czech Republic	-1.03	5	-3.40**	3
Estonia	-3.47**	1	-3.34**	4
Hungary	-2.62*	1	-3.04**	3
Poland	-2.95*	1	-3.28**	3
Slovakia	-1.09	3	-3.43**	3
Slovenia	-3.42**	2	-4.60**	3
Turkey	-1.78	5	-3.24**	3

*Note: Augmented Dickey-Fuller test results, IP data in levels and seasonal differences. We apply the small-sample, lag-adjusted critical values for the ADF test by Cheung and Lai (1995). Significance at the 5 percent level is indicated by "**", at 10 percent by "*". For the levels, we include a deterministic trend.*

Having determined the lag lengths this way, we employ the Augmented Dickey-Fuller (ADF) test for unit roots with intercept and, in the case of levels, with a deterministic trend. Table 2.2 outlines the results of the unit root tests and the corresponding lag lengths. In the case of levels, the null hypothesis of a unit root cannot be rejected in all but three cases on the 5 percent significance level. For the seasonal differences, we reject the null hypothesis of a unit root on the 5 percent level for all countries. Hence, we consider the data stationary in differences.

Seasonal cointegration

In the presence of persistent seasonal behaviour of a time series, one option is to seasonally adjust the data up-front. However, as Lee (1992) points out, seasonal adjustment may result in mistaken inference regarding economic relationships, particularly with finite samples. Moreover, it comes at the cost of losing valuable information if seasonal fluctuations are an important source of variation in the system. Therefore, working with non-adjusted data and incorporating seasonality into the statistical model is the preferable option, implemented by the concepts of seasonal integration and seasonal cointegration.

Seasonal integration implies that a series can have a unit root not only in the standard case of zero frequency but at seasonal frequencies as well and exhibits a spectrum with distinct peaks at these frequencies. Hence, the series exhibits "long memory" in the sense that shocks last forever and may permanently change seasonal patterns. In the case of quarterly data, we may identify an annual cycle of the four seasons, i.e. a quarter-cycle per quarter, and/or a semi-annual pattern with two cycles per year, i.e. a half-cycle per quarter. The seasonal series exhibits a spectrum with distinct peaks at the seasonal frequencies $\theta_s \equiv 2\pi j/s,\ j = 1, ..., s/2$. With $s = 4$ indicating the number of observations per year, we have π, and $\pi/2$ as seasonal frequencies, besides zero frequency for the standard case. Hyllenberg et al. [HEGY] (1990) provide the quarterly-data example

$$(1 - B^4)x_t = \varepsilon_t, \tag{2.2}$$

with B as the backshift operator. This equation can be expressed as

$$\begin{aligned}(1-B)(1+B+B^2+B^3)x_t &= \varepsilon_t \\ (1-B)(1+B)(1+B^2)x_t &= \varepsilon_t \\ (1-B)S(B)x_t &= \varepsilon_{t.}\end{aligned} \tag{2.3}$$

Following Lee (1992), the process x_t is denoted $x_t \sim I_\theta(d)$, with the frequencies $\theta = 0, \pi$, and $\pm\pi/2$ and integration order d. In this exercise, we focus on the standard case of $d = 1$. Thus, the process has four roots with modulus one: one root at the zero frequency ($\omega = 0$), one root at two cycles per year which equals half a cycle per quarter ($\omega = \frac{1}{2}$), and a pair of complex roots at one cycle per year which equals a quarter cycle per quarter ($\omega = \frac{1}{4}$).

Seasonal cointegration designates the case in which two or more series share common stochastic seasonals. In a generalisation of Engle and Granger's (1987) cointegration approach, Lee (1992) states that the components of a seasonally integrated vector $x_t \sim I_\theta(1)$ are seasonally cointegrated at frequency θ, denoted by $x_t \sim CI_\theta(1,1)$ if there exists a vector α ($\neq 0$) so that $z_t = \alpha' x_t \sim I_\theta(0)$. Intuitively, seasonal cointegration has a connotation similar to the standard cointegration approach. It implies that an innovation has only a temporary effect on the seasonal behaviour of $z_t = \alpha' x_t$ but a permanent impact on the seasonals of x_t.

We consider an n-dimensional vector autoregressive process of order p, $VAR(p)$, of the form

$$x_t = \Phi_1 x_{t-1} + \Phi_2 x_{t-2} + ... + \Phi_p x_{t-p} + \varepsilon_t, \tag{2.4}$$

with $t = 1, 2, ..., T$ and $\varepsilon_t \sim i.i.d.N_n(0, \Omega)$. The determinant of the matrix polynomial $\Phi(z) = I - \Phi_1 z - \Phi_2 z^2 - ... - \Phi_p z^p$ has four roots on the unit circle ($z = \pm 1, \pm i$), corresponding to the frequency cases $\omega = 0, \frac{1}{2}, \frac{1}{4}$. The related seasonal error correction model (SECM) can be expressed as

$$\begin{aligned}\Delta_4 x_t &= \mu + \Psi D_t + \delta t + \Pi_1 y_{1,t-1} + \Pi_2 y_{2,t-1} + \Pi_3 y_{3,t-1} + \Pi_4 y_{4,t-1} \\ &\quad + \Gamma_1 \Delta_4 x_{t-1} + ... + \Gamma_{p-4} \Delta_4 x_{t-p+4} + \varepsilon_t, \end{aligned} \tag{2.5}$$

where $\Delta_4 x_t = (1 - B^4)x_t$ is stationary and $h = p - 4$ is the order of the SECM. In addition, $y_{1,t} = S_1(B)x_t = (1 + B + B^2 + B^3)x_t$, $y_{2,t} = S_2(B)x_t = (1 - B + B^2 - B^3)x_t$, $y_{3,t} = S_3(B)x_t = B(1 - B^2)x_t$, and $y_{4,t} = S_4(B)x_t = y_{3,t+1}$. The seasonal filter $S_k(B)$ with $k = 1, ..., 4$ eliminates unit roots at all frequencies other than the one in the filtered series. μ, ΨD_t, and δt represent deterministics, namely a constant, seasonal dummies, and a linear time trend, respectively. If the coefficient matrices Π_k that convey the long-run information are of reduced rank r, with $0 < r < n$, then Π_k can be decomposed as $\Pi_k = \gamma_k \alpha'_k$, where γ_k is an $n \times r$ matrix of short-run coefficients and $\alpha'_k y_{k,t-1}$ is an $r \times 1$ vector of stationary cointegration relations. $\Pi_1 = \gamma_1 \alpha'_1$ corresponds to seasonal cointegration at frequency zero ($\omega = 0$), $\Pi_2 = \gamma_2 \alpha'_2$ refers to biannual frequency ($\omega = \frac{1}{2}$), while a seasonal cointegration test at annual frequency ($\omega = \frac{1}{4}$) can be conducted based on the property of Π_3, given $\Pi_4 = 0$. Thus, the SECM be transformed into

$$\begin{aligned}\Delta_4 x_t &= \mu + \Psi D_t + \delta t + \gamma_1 z_{1,t-1} + \gamma_2 z_{2,t-1} + \gamma_3 z_{3,t-1} + \gamma_4 z_{4,t-1} \\ &\quad + \Gamma_1 \Delta_4 x_{t-1} + ... + \Gamma_{p-4} \Delta_4 x_{t-p+4} + \varepsilon_t, \end{aligned} \tag{2.6}$$

with the error correction terms $z_{k,t} = \alpha'_k y_{k,t}$. The likelihood ratio test for the hypothesis of r cointegrating vectors, i.e. $H_0 : \Pi_k = \gamma_k \alpha'_k$, with $n \times r$ matrices γ_k and α_k (for $k = 1, 2, 3$), can be summed up by the expression

$$TR_k = -T \sum_{j=r+1}^{n} \ln(1 - \lambda_{k,j}). \tag{2.7}$$

The smallest $(n-r)$ squared partial canonical correlations between each $y_{k,t-1}$ and $\Delta_4 x_t$ are represented by $\lambda_{k,r+1}, ..., \lambda_{k,n}$, given lagged $\Delta_4 x_t$'s $\{\Delta_4 x_{t-i}$ for $i = 1, ..., k\}$, other $y_{1,t-1}(1 \neq k)$ and deterministic terms. For further details on the test procedure, see Lee (1992) and Lee and Siklos (1995).

Table 2.3: Seasonal cointegration results

Country	h	rank	Frequency		
			0	1/2	1/4
Croatia	2	$r = 0$	6.77	22.12**	26.43
		$r = 1$	1.74	2.64	8.46
Czech Rep.	1	$r = 0$	6.45	22.75**	19.62
		$r = 1$	1.02	8.94	3.24
Estonia	2	$r = 0$	6.52	15.85	11.32
		$r = 1$	1.31	3.13	3.38
Hungary	3	$r = 0$	12.05	14.54	22.29
		$r = 1$	2.31	5.57	1.27
Poland	1	$r = 0$	10.53	14.97	16.42
		$r = 1$	1.12	6.90	4.58
Slovakia	1	$r = 0$	3.07	20.53	11.66
		$r = 1$	0.73	5.89	4.88
Slovenia	2	$r = 0$	13.57	12.22	28.62
		$r = 1$	0.94	2.80	3.93
Turkey	2	$r = 0$	7.76	7.82	20.81
		$r = 1$	0.42	2.48	7.49

*Note: Table 2.3 reports bivariate seasonal cointegration results of quarterly, non-seasonally adjusted industrial production data of each country vis-à-vis the euro area. The seasonal error correction model (SECM) order is given by h. In the case of cointegration, rejection of the null hypothesis of no cointegration ($r = 0$) at the 5 percent level is marked with "**" and corresponds to acceptance of one cointegration vector ($r = 1$), as in the 1/2 frequency cases of Croatia and the Czech Republic. In all other cases, we find no seasonal cointegration as the null hypothesis of no cointegration ($r = 0$) is widely accepted. The finite-sample critical values are provided by Lee and Siklos (1995).*

In the case of the CEECs, we construct bivariate cointegrated VARs, testing the euro area vis-à-vis each of the eight countries in the data set. We first set up the VARs and determine the optimal lag length p. As in the univariate case, minimising the Akaike information criterion alone may lead to biased results. Therefore, we also consider the sequential modified LR test statistic which tends to deliver shorter lag lengths than AIC. To finally decide upon p, we check the VAR residual serial correlation LM test to ensure that there is no significant autocorrelation left in the residuals of the VAR. The order of the corresponding SECM is then $h = p - 4$. Finally, we compute the seasonal cointegration test statistics for the frequencies zero, $\frac{1}{2}$, and $\frac{1}{4}$. To allow for the limited data series, we employ the finite sample critical values as tabulated in Lee and Siklos (1995). Table 2.3 reports the seasonal cointegration test results including the corresponding SECM orders. In fact, not a single country reveals a cointegration relation with the euro area at zero frequency. At frequency $\frac{1}{2}$, the test rejects the notion of no cointegration ($r = 0$) for Croatia and the Czech Republic. In the case of frequency $\frac{1}{4}$, there is again no evidence of cointegration for any series.

The weak cointegration results are in line with the existence of beta convergence. The CEECs have obviously not yet reached a steady state equilibrium with the euro area but are still in the process of transition.

2.2 Cycle analysis

After having explored the trends and long-run comovement among the CEECs and the euro area, we now turn to the cyclical components of the output series to gain insights regarding short-run synchronisation. Again, we use the non-adjusted, quarterly industrial production series as in the cointegration part of the analysis. To investigate common business cycle behaviour, we first calculate simple correlations of growth rates and cyclical components of the data. Next, we move on to the common feature framework to test for synchronised common cycles and then generalise the approach to non-synchronised cycles, testing for codependence. Given that persistent stochastic seasonality may influence the variables, both trends and cycles must be disentangled from seasonal components. Employing the approach of Cubadda (1999), that part builds on the seasonal cointegration results and thus takes seasonality into account.

2.2.1 Cycle correlations

First, we investigate contemporaneous correlations. Table 2.4 presents the correlation coefficients of each CEEC with the euro area. We correlate the growth rates in form of the seasonal differences before we extract the cyclical component from the data, applying the band-pass filter by Baxter and King (1999).[5]

Table 2.4: Contemporaneous correlation

Country	Seasonal differences	BK cycles $k = 4$	BK cycles $k = 8$
Croatia	0.22	0.49	0.40
Czech Republic	0.23	0.74	0.53
Estonia	0.35	0.68	0.72
Hungary	0.79	0.79	0.93
Poland	0.49	0.57	0.68
Slovakia	0.25	0.70	0.58
Slovenia	0.47	0.63	0.75
Turkey	0.24	0.52	0.57

Note: Contemporaneous correlations with the euro area, data in seasonal differences and as Baxter-King band-pass filtered cyclical components (BK cycles) with alternative lead-lag lengths $k = 4$ and $k = 8$.

We observe that the correlation coefficients tend to be larger in the case of the cycles as compared to the seasonal differences. Except for Croatia, all countries display cycle correlation coefficients larger than 0.5. Apparently, the correlation of the short-term cycles tends to be larger than the comovement of long-run stochastic trends which are

[5] We use the conventional "Burns-Mitchell" settings for the minimum and maximum oscillation period, see Burns and Mitchell (1946). For quarterly data, these translate into a minimum of 6 and a maximum of 32 oscillation periods. The choice of the appropriate lead-lag length, k, however, involves a trade-off because larger ks downsize the already small number of observations. In table 2.3, we present results with alternative lead-lag lengths $k = 4$ and $k = 8$.

still included in the seasonal differences. Also, this may be due to the fact that the Baxter-King filter can lead to amplified cycle frequencies and spurious cycles.[6]

Across countries, we note that Hungary, Slovenia and Poland exhibit large correlation coefficients, followed by the Czech Republic and Estonia. Croatian and Turkey, on the other hand, assume low values in all set-ups. These simple correlations deliver a preliminary indication of the data properties. In the following, we take an econometrically more advanced approach by analysing common business cycles within the framework of common features and codependence.

2.2.2 Synchronised common cycles

The common feature framework is based on Engle and Kozicki (1993) and Vahid and Engle (1993) and investigates the existence of synchronisation, i.e. contemporaneous comovement among business cycles. In analogy to the non-stationary cointegration case, common feature analysis puts series together which exhibit a certain stochastic component each, e.g. autocorrelation. The series are then said to have a common feature, or a common serial correlation cycle, if there is a linear combination which does not have any correlation with the past. It is required, however, that the individual series have the same autoregressive order for the common feature to cancel out in the linear combination. The rank of the common feature space provides the number of common feature vectors. Engle and Kozicki (1993) employ a two-stage least squares approach in which one variable is regressed on the lagged values of all variables, serving as instruments. This way, they test whether the dependence of the variables on the past is only through common channels which would, in turn, hint at the existence of common cycles.

In Vahid and Engle (1993), the analysis is further developed to incorporate results of the preceding cointegration tests. In particular, the differences of cointegrated variables

[6] See Guay and St-Amant (1997).

are related to the past not only through the lagged differences, $\Delta_4 x'_{t-1}, ..., \Delta_4 x'_{t-h+1}$, but also through the lags of the estimated error correction terms, $\widehat{z}_{1,t-1}, ..., \widehat{z}_{4,t-1}$. This set of variables constitutes the "relevant past" and is summed up as $w_t = (\Delta_4 x'_t, ..., \Delta_4 x'_{t-h+1}, \widehat{z}_{1,t}, ..., \widehat{z}_{4,t})'$. Regressing the common feature linear combination on these past variables delivers the TR^2 which serves as a statistical measure of the dependence of the linear combination on the relevant past. The linear combination which minimises the TR^2 points at the potential common feature vector. The minimand will then be the limited information maximum likelihood (LIML) estimator of the regression of one of the elements of $\Delta_4 x_t$ on the others, employing the relevant past as instruments. The LIML estimator is also the canonical covariate corresponding to the smallest canonical correlation between the transposed differences and the relevant past information. Testing for the number of linearly independent common feature vectors is based on the number of zero canonical correlations and goes back to Tiao and Tsay (1989). They test the significance of the smallest m canonical correlations which translates into the null hypothesis of at least m common feature vectors. The understanding is similar to the notation in cointegration. In fact, m is the dimension of the common feature space and $n-m$ indicates the number of common cycles. In the bivariate case, the existence of $m=1$ common feature vector corresponds to one common cycle. The test statistic is given by

$$C(h, m) = -(T-h-4) \sum_{j=1}^{m} \ln\left(1 - \widehat{\lambda}_j\right) \tag{2.8}$$

where $\widehat{\lambda}_j$ is the jth smallest sample squared canonical correlation between $\Delta_4 x_t$ and w_{t-1}. The statistic $C(h, m)$ is χ^2-distributed under the null hypothesis, with $(m(nh+r+m-n))$ degrees of freedom. The SECM order is denoted by h and r is the number of cointegrating vectors in the system.

Table 2.5 reports the results of the common feature tests in its centre column. Slovenia is the only one of the eight relationships tested that reveals a common serial correlation feature. The corresponding test statistic for the hypothesis of one common feature vector ($m = 1$) is accepted while the notion of two common feature vectors ($m = 2$) can be clearly rejected at the 1 percent level. Hence, Slovenia shares one common business cycle with the euro area. For all other countries, we cannot find a common feature vector, indicated by the rejection of the hypothesis of any common feature vectors at the 1 percent level.

Table 2.5: Common feature/codependenc results

Country	rank	Common features	Codependence Order 1	Order 2	Order 3
Croatia	$m = 1$	24.46***	10.52**	9.99**	2.46
	$m = 2$	82.76***	28.17***	18.87**	8.60
Czech Rep.	$m = 1$	35.26***	1.84	4.54	2.08
	$m = 2$	89.03***	11.89*	10.19	10.03
Estonia	$m = 1$	33.72***	9.59**	4.11	2.01
	$m = 2$	83.44***	21.71***	16.01**	10.49
Hungary	$m = 1$	11.46***	3.52*	0.29	0.06
	$m = 2$	65.41***	10.99**	3.76	3.48
Poland	$m = 1$	38.18***	6.95***	1.53	0.79
	$m = 2$	85.04***	16.79***	8.44*	6.15
Slovakia	$m = 1$	23.25***	2.52	0.01	0.37
	$m = 2$	75.90***	14.08***	8.13*	11.40**
Slovenia	$m = 1$	6.16	2.65	4.68	5.71
	$m = 2$	59.87***	14.82*	15.69**	10.24
Turkey	$m = 1$	35.98***	11.71***	7.91**	3.12
	$m = 2$	94.06***	23.08***	16.17**	11.50

*Note: Table 2.5 reports bivariate seasonal common feature and codependence results of quarterly, non-seasonally adjusted industrial production data of each country vis-à-vis the euro area. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector ($m = 1$) and rejecting a second vector ($m = 2$), we conlude the existence of $n - m = 2 - 1 = 1$ common cycle.*

In the case of Estonia, it is unlikely to find common serial correlation features *ex ante* since Estonia is the only country in the sample which does not have the same autoregressive order as the euro area when analysed individually. As shown above, the euro area is modelled by an AR(3) process, as are all other countries except Estonia for which an AR(4) model applies. We acknowledge, however, a certain natural uncertainty in the determination of lag length parameters.

In conclusion, we find only little overall evidence of contemporaneous short-run comovement of the CEECs with the euro area.

2.2.3 Non-synchronised common cycles

The concept of common serial correlation features is very restrictive in that it requires the variables to react contemporaneously to shocks. Consequently, this allows us only to detect perfectly synchronised common cycles or no common cycles at all. Those shocks that require some time to propagate across countries at different speeds are not captured. The codependence framework of Vahid and Engle (1997) relaxes this constraint and formulates the same idea in a more general setting. In particular, it permits the series to respond to shocks with a certain delay. Even if one country does not immediately react to a shock in one country, it may fully react at a later stage. This kind of non-synchronised common cycle can be measured with the codependence test. The system is then said to be codependent if the impulse responses of the variables are collinear beyond q periods. For $q = 0$, the codependence test is equivalent to the common feature test.

The test statistic for the null hypothesis of at least m codependence vectors after the qth period is given by

$$C(h, q, m) = -(T - h - q - 4) \sum_{j=1}^{m} \ln \left(1 - \frac{\widehat{\lambda}_j(q)}{d_j(q)} \right). \tag{2.9}$$

Now, $\widehat{\lambda}_j(q)$ represents the jth smallest sample squared canonical correlation between $\Delta_4 x_t$ and w_{t-q-1}, and $d_j(q)$ is defined as $d_j(q) = 1 + 2 \sum_{i=1}^{q} \widehat{\rho}_i(\widehat{\alpha}^{*\prime} \Delta_4 x_t) \widehat{\rho}_i(\widehat{\gamma}_w^{\prime} w_{t-q-1})$, where $\widehat{\rho}_i(\cdot)$ is the lag-i sample autocorrelation of the series in argument, and $\widehat{\alpha}^*$ and

$\widehat{\gamma}_w$ are the sample canonical variates associated to $\widehat{\lambda}_j(q)$. Under the null hypothesis, the statistic $C(h,q,m)$ is asymptotically distributed as $\chi^2(m(nh+r+m-n))$. The test statistic is based on the following intuition. The required linear combinations $\alpha^{*\prime}\Delta_4 x_t$ and $\gamma_w' w_{t-q-1}$ have non-zero cross correlations up to lag q but zero cross-correlations from lag $q+1$. The $(q+1)$th cross correlation between $\Delta_4 x_t$ and w_{t-q-1} is the smallest canonical correlation. According to the Bartlett formula, the corresponding variance is $d_j(q)/(T-h-q-4)$.

The codependence results for Central and Eastern Europe are presented in the right-hand panel of table 2.5. We accept the notion of one codependence vector ($m=1$) at order one in the cases of Hungary and Slovakia. In the Czech case, we can reject the hypothesis of a second codependence vector only at the 10 percent level. For Hungary, we accept one vector at the 5 percent level although rejection is indicated for the 10 percent level. For Slovakia, the case is more clear-cut: one vector is definitely accepted while a second vector is rejected even at the 1 percent level. Overall, we conclude that these three countries share common but non-synchronised cycles with the euro area and adjust after $q=1$ quarter, with the Czech Republic as a borderline case. For the remaining countries, both $m=1$ and $m=2$ are rejected when testing for first-order codependence. However, we find one codependence vector at order 2 for Estonia. Thus, Estonia seems to respond to a euro area shock after two quarters.

Summing up, the cycle analysis can find a common feature vector only in the case of Slovenia. Hence, the case for contemporaneous comovement, i.e. synchronised common cycles, of most CEECs with the euro area is limited. When delays in the response are permitted, however, we do find common cycles. Estonia, Hungary, Slovakia and, with some uncertainty, the Czech Republic, exhibit one codependence vector after one or two periods. The fact that Croatia and Turkey do not show any common cycles with the euro area is not unexpected since these two countries are only at the beginning of their integration process with the EU and started formal accession talks only recently.

Identifying Poland as the third country without any commonality in its cycles with the euro area is surprising because a number of studies identify Poland as one of the leading countries in terms of business cycle synchronisation.[7] This divergence may partly stem from the fact that the Polish economy has a large agricultural sector which is not captured by our output proxy, industrial production.[8] On the other hand, our finding may be substantiated in that Poland is the largest of the new EU member states. Hence, it would naturally be less integrated than its smaller neighbours. On the whole, the CEECs show an intermediate degree of cycle comovement with the euro area but they do not yet seem to be substantially synchronised.

2.3 Conclusion

Employing a trend/cycle approach, this study investigates the degree of output integration of eight Central and Eastern European countries with the euro area. According to the traditional OCA framework, a higher the degree of business cycle synchronisation impies lower costs of renouncing individual monetary policy when adopting the euro.

The trend analysis is the foundation of the subsequent cycle tests and estimates catching-up and steady-state convergence. A simple cross-section regression confirms the beta convergence hypothesis in that it indicates significantly higher average growth rates for those countries with lower initial income levels. Apparently, most countries under investigation are still in the process of transition towards the steady-state equilibrium. This presumption is confirmed by the cointegration analysis. Using quarterly, non-seasonally adjusted industrial production data for the aggregate euro area plus the eight individual countries, we perform bivariate seasonal cointegration tests for each country vis-à-vis the euro area. We find no cointegration at zero or 1/4 frequency. Croatia and

[7] See, for example, Fidrmuc and Korhonen (2006).

[8] We ran alternative tests on GDP data which include the agricultural sector. Although the GDP dataset is not sufficiently rich to be presented in full, it indicates an intermediate degree of synchronisation for Poland. See chapter 4 for an up-dated GDP dataset including Poland.

the Czech Republic cointegrate with the euro area when tested at frequency 1/2. This widely negative cointegration result is consistent with the existence of beta convergence because the two convergence concepts can be regarded as mutually exclusive.

The tests for common business cycles divide into the categories of synchronised and non-synchronised common cycles. When testing for contemporaneous common serial correlation features, we find only one case of synchronised common cycles, Slovenia. This indicates that the remaining countries do not share common cycles with the euro area. To allow for a delay in the response to shocks, we also test for codependence. Due to inefficiencies in the propagation mechanism of shocks, the CEECs may respond to shocks but not in the initial period. In fact, we find evidence of first-order codependence for Hungary, Slovakia and, as a borderline case, the Czech Republic. Estonia shows signs of second-order codependence. These countries can, therefore, be considered as having an intermediate degree of cyclical comovement with the euro area. For Poland, as well as for the candidate countries Croatia and Turkey, we do not find any codependence. Their cycles do not seem to align to that of the euro area even after a certain delay.

On the whole, our results suggest that real integration of the CEECs with the euro area is still limited. Only Slovenia appears well-equipped for joining the euro soon. In fact, the Slovenian government assured its intention to do so by joining ERM II immediately after EU accession in 2004. In the meantime, the European Commission has approved of the Slovenian request to introduce the euro in January 2007. For most of the other CEECs, however, giving up individual monetary policies at too early a stage may entail the risk of incurring major costs. There is still some way to go to achieve business cycle synchronisation.

However, a number of questions remain open. First, there is still considerable uncertainty concerning the data situation with regard to the CEECs. The ten years of data available represents the lower boundary of statistically meaningful time series analysis. As time proceeds, longer series will allow for more reliable investigation. Second, our approach attempts to capture the "meta-property" of the OCA theory. Finding ways to analyse the individual OCA criteria in a convincing and consistent manner has not yet been achieved. Third, the OCA framework itself is not devoid of ambiguity. It has been argued that business cycle synchronisation may not be an *ex ante* requirement but may evolve endogenously after the adoption of a single currency. Moreover, the proponents of the "Mundell II" framework have argued that currency union membership may be desirable even in the presence of non-synchronised cycles if risk-sharing is facilitated by integrated financial markets.[9] Although more evidence regarding these hypotheses remains to be inferred from the ongoing EMU experiment, the following chapters shed more light on these questions. So far, however, it appears that EU accession has by no means concluded transition of the CEECs and is only one milestone on road to the euro.

[9] See McKinnon (2002).

Chapter 3

Determinants of business cycle synchronisation across euro area countries[1]

Will the euro area countries move together or apart in their business cycle fluctuations? Since the launch of the single currency, researchers and policy makers have sought to learn more about the driving forces of business cycles and the role of the euro. The effects of a common currency on business cycle synchronisation is at the heart of the second major strand of currency union economics, the endogeneity of optimum currency areas. If a common currency promotes trade and if trade increases business cycle synchronisation, Frankel and Rose (1998) argue, then an *ex ante* non-optimum currency area may turn into an OCA *ex post*, due to the unfolding impact of the currency union itself. So is the euro area, arguably not an *ex ante* OCA, going to be optimal *ex post*?

This chapter examines the underlying factors of business cycle synchronisation in the euro area. We do not address the endogeneity question directly because at such an early stage, it proves difficult to isolate a clear effect of the euro. Instead, we follow Frankel and Rose (1998) and approach the topic by asking which factors are signficantly associated with business cycle synchronisation across euro area countries. A positive

[1] Most of this chapter was produced in cooperation with Catherine Guillemineau at the European Central Bank.

association of trade and cycle synchronisation may be interpreted as an indication of OCA endogeneity.

Various studies have shown that European business cycles have become increasingly synchronous, see Artis and Zhang (1997, 1999), Massmann and Mitchell (2004). Applying Markov Switching VAR models, Artis et al. (2004) find evidence of a distinct European business cycle. Few academics have, however, explored the underlying factors behind cycle synchronisation in Europe. Baxter and Kouparitsas (2004) and Imbs (2004) analysed large samples of both developing and industrialised countries and found trade flows, specialisation, and financial integration to be important factors for business cycle synchronisation. Their results are, however, partly conflicting and seem to depend on the country and time samples chosen.

In the following, we investigate a variety of potential determinants of cycle synchronisation in the context of European monetary integration. The purpose of our analysis is to find out why, inside the euro area, the business cycles of different countries may be synchronous or asynchronous and why they may converge or diverge. We test some standard determinants and, in addition, consider a number of EMU-specific policy and structural indicators which, to our knowledge, have not been tested in this context. We check robustness by applying the extreme-bounds analysis framework as suggested by Leamer (1983) and further developed by Levine and Renelt (1992) and by Sala-i-Martin (1997). Also, we divide our 25-year sample period into sub-samples in order to capture changing effects throughout the different stages of European integration.

We find that bilateral trade has indeed been a robust, positive determinant of business cycle synchronisation. Hence, we see the endogeneity hypothesis of Frankel and Rose (1998) confirmed for the euro area: countries with larger trade volumes tend to have more closely synchronised business cycles. Although we observe this phenomenon over the whole sample, its explanatory power seems to be driven mainly by the earlier sub-sample, 1980-1996. During the period of preparation for EMU and actual currency union, since 1997, we find that the differences in trade structure emerge as robust determinants of cycle synchronisation. In other words, the degree of intra-industry trade plays an increasingly important role in binding euro area business cycles together. In combination with our descriptive finding of rising intra-industry trade among euro area countries, this result gives rise to cautious optimism with regard to *ex post* optimality of the euro area.

Regarding our policy and structural indicators, fiscal deficit differentials appear to have driven differences between business cycles until the preparation for EMU. With the implementation of the Stability and Growth Pact, fiscal policy seems to have become less pro-active and fiscal deficit differentials have lost some of their explanatory power. In contrast, similarities in monetary policies, measured by interest rate differentials, have emerged as a robust determinant of business cycle synchronisation. Also, differences in the size of industrial sectors, stock market comovement and similar competitiveness situations appear to have good explanatory power. On the other hand, we could not detect any robust impact of nominal exchange rate variability, bilateral bank capital flows or differences in labour market flexibility on cycle synchronisation. The missing effect of mere exchange rate stabilisation on the synchronisation of business cycles is in line with the endogeneity hypothesis of optimum currency areas which predicts that only irrevocably fixed exchange rates, i.e. currency union, would unleash synchronisation dynamics.

The remainder of this chapter is structured as follows. Section 3.1 provides an overview of the recent literature, introduces the potential determinants of cycle correlation and presents some stylized facts. Section 3.2 outlines the methodology of the extreme-bounds analysis (EBA) and presents the EBA results. Section 3.3 summarises and concludes.

3.1 The potential factors behind business cycle synchronisation in the euro area

This section deals with the potential determinants of business cycle synchronisation. The first sub-section reviews traditional factors from the recent literature and suggests new indicators that seem particularly relevant in the context of EMU. Based on these considerations, we then specify our variables and present some stylised facts.

3.1.1 Traditional and new factors

The foremost candidate expected to influence business cycle synchronisation is *trade*. In theory, however, it is unclear whether intensified bilateral trade relations result in more or in less synchronised business cycles. Spill-overs due to common aggregate demand and productivity shocks would result in a positive effect of trade integration on business cycle synchronisation.[2] On the other hand, intensified trade relations may also lead to a higher degree of specialisation in different sectors across countries, due to the exploitation of comparative advantages. As a result, business cycles may become more asynchronous.[3] The underlying question is whether bilateral trade occurs mainly in similar or different sectors. If trade flows are predominantly intra-industry, as it is the case for most of the trade among industrialised countries, then we would expect the first effect to materialise. If bilateral trade is, or increasingly becomes, mostly inter-industry, the second prediction would dominate. Whether an intensification of bilateral trade relations will result in more or less synchronous business cycles can be assessed by paralleling the evolution of bilateral trade and of relative trade specialisation. Smaller cross-country differences in *trade specialisation* would indicate an intensification of intra-industry trade conducive of more synchronous business cycles.

Given the unclear theoretical case, the question is fundamentally an empirical one. In

[2] See Gruben et al. (2002).

[3] This point was made by Krugman (1993). He postulates that, due to a specialisation effect of trade, even an *ex ante* OCA may turn out to be non-optimal *ex post*.

their seminal contribution on "the endogeneity of the optimum currency area criteria", Frankel and Rose (1998) estimated a single-equation model based on a large sample of developing and industrialised countries and found a strong and robust positive relationship between bilateral trade and cycle synchronisation. This result is confirmed by Baxter and Kouparitsas (2004). Imbs (2004) employs a simultaneous-equations approach and verifies the overall positive impact of trade on business cycle synchronisation but points out that "a sizable portion is found to actually work through intra-industry trade."[4]

The effects of *economic specialisation* on cycle synchronisation have also been measured directly. Stockmann (1988) emphasises the importance of sectoral shocks for the business cycle since two countries will be hurt similarly by sector-specific shocks if they have economic sectors of similar nature and size. Hence, we would expect the degree of differences in sectoral specialisation to be negatively related to cycle synchronisation, i.e. the more dissimilar the economies, the less correlated their cycles. Empirical studies however, find conflicting evidence regarding the robustness of this effect.[5]

Financial integration is the third major field of determinants. The literature is ambiguous on the effect of financial integration on the synchronisation of business cycles. Kalemli-Ozcan et al. (2003) argue along the lines of Krugman (1993) that countries with a high degree of financial integration tend to have more specialised industrial patterns because firms need not spread production risk geographically. Hence, business cycles will be less synchronised. Evidence from the financial crises and contagion literature, however, points out the role of psychological spill-overs and indicates a direct, positive effect of capital flows to business cycle synchronisation.[6] Kose et al. (2003) find that financial integration tends to enhance international spillovers of macroeconomic fluctuations leading to more business cycle synchronisation. Imbs (2004, 2006) tests this direct link and finds a positive effect dominating the indirect link via specialisation dynamics.

[4] Imbs (2004: 733).

[5] While Imbs (2004) asserts that specialisation patterns play an independent role in cycle correlation, this notion is rejected by Baxter and Kouparitsas (2004).

[6] See ECB(2004).

Even if financial integration leads to intensified specialisation, the latter may occur in similar and not different sectors, as argued by Obstfeld (1994). In his model, countries that gain new access to international financial markets can now all equally well afford to specialise on risky high-tech industries. This catching-up process leads, despite specialisation, to more and not less similar economic structures across countries.

The three major determinants of business cycle synchronisation and the various, partly opposite transmission channels are illustrated in figure 3.1.

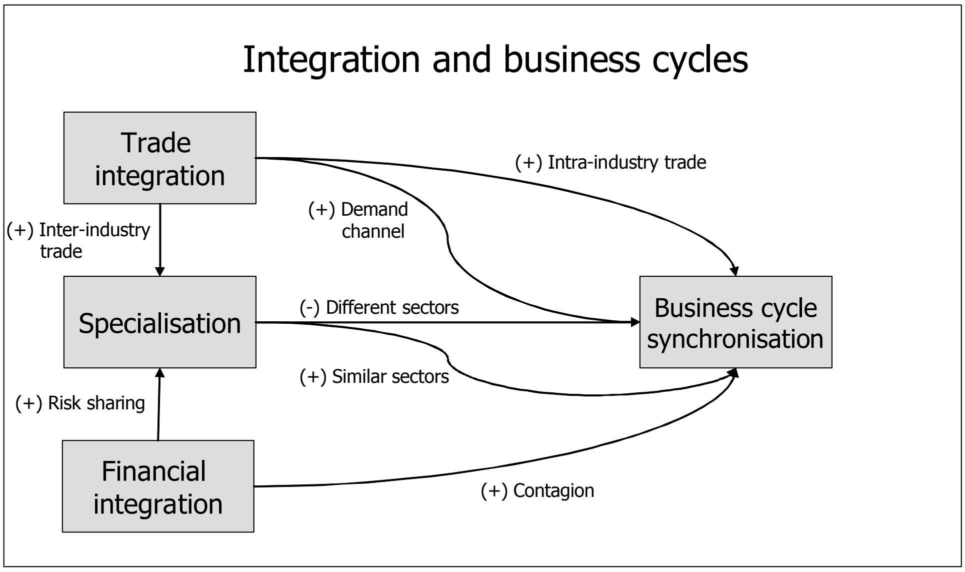


Figure 3.1: Major determinants of business cycle synchronisation and channels of influence; adapted from ECB (2004) and Imbs (2004).

In addition to the above variables used in the literature, we suggest a number of *policy and structural indicators* that seem particularly relevant for the euro area. We ask whether the degree of similarity in various economic variables between two countries has influenced the bilateral synchronisation of business cycles. The policy indicators include bilateral differentials in the real short-run interest rate as a measure of the monetary policy stance, nominal exchange rate variations, and differentials in fiscal deficits. The structural indicators capture competitiveness differentials, stock market comovements, and labour market flexibility. Finally, we add geographical distance between countries and relative country size in terms of population, in order to control for exogenous factors. The following sub-section specifies these variables in detail.

3.1.2 Data and variable specification

As a measure of *business cycle synchronisation* in the euro area, we compute bilateral correlation coefficients between the cyclical part of real annual GDP for each pair of countries, drawing 66 pairs among the twelve euro area countries over the 1980-2004 period.[7] The cyclical parts are obtained by applying the Baxter-King band-pass filter, which Baxter and King (1999) suggested specifically in order to measure business cycle correlations.[8]

The remainder of this sub-section provides detailed information on the specification of variables which we selected as potential determinants of business cycle synchronisation. In general, we take averages of the annual data which cover the period 1980-2004. Exceptions due to missing years or countries are indicated in the respective sub-sections. The data apply to the twelve individual euro area countries. We use bilateral country data where available and construct them from individual country data otherwise. Hence, the terminology in the following equations corresponds to the country indices $i = 1, \ldots, 12$ and $j = 1, \ldots, 12$, $i \neq j$, as well as the time index $t = 1, \ldots, 25$. The first set of variables draws largely on the determinants used by Baxter and Kouparitsas (2004)[9]

[7] For the pre-euro period, national currencies are converted using the fixed euro exchange rate as to exclude the influence of exchange rate fluctuations. We use annual data for GDP because, for a number of euro area countries, quarterly data are unavailable prior to 1997.

[8] For the Baxter-King filter, we employ the standard Burns-Mitchell settings for annual data, i.e. maximum lead/lag length k = 3, shortest cycle pass p =2 and longest cycle pass q = 8. We are aware that, due to the one-sided filtering windows at the margins of the sample, the estimates of the cyclical components may decrease in accuracy at the beginning and the end of the data period.

[9] Baxter and Kouparitsas (2004) use initial values for the determinants of business cycle correlation. We prefer full-sample and sub-sample averages based on the consideration that cross-country correlations of business cycles may not be appropriately explained solely by the initial values of the potential determinants since nearly all variables have undergone major changes since 1980.

and Imbs (2004). The second set of variables consists in policy and structural indicators which appear particularly relevant in the context of EMU. Appendix table A.1 gives an overview of the variables and provides the data sources.

Traditional determinants of business cycle synchronisation

The independent variable *bilateral trade* is constructed in two alternative ways. First, it is defined as the average of the sum of bilateral exports and imports, divided over the sum of total exports and imports, denoted by BTT_{ij}.

$$BTT_{ij} = \frac{1}{T}\sum_{t=1}^{T}\frac{x_{ijt} + m_{ijt} + x_{jit} + m_{jit}}{x_{it} + m_{it} + x_{jt} + m_{jt}},$$

where x_{ij} denotes the exports of country i to country j at time t, m_{it} stands for the imports of country i from country j at time t, and x_{it} and m_{it} represent total exports and imports of country i.

Second, the sum of national GDPs, y_i and y_j, serves as scaling variable which gives

$$BTY_{ij} = \frac{1}{T}\sum_{t=1}^{T}\frac{x_{ijt} + m_{ijt} + x_{jit} + m_{jit}}{y_{it} + y_{jt}}.$$

The variable *trade openness* is calculated as the sum of total exports and imports of both countries, divided by the sum of national GDPs:

$$TTY_{ij} = \frac{1}{T}\sum_{t=1}^{T}\frac{x_{it} + m_{it} + x_{jt} + m_{jt}}{y_{it} + y_{jt}}.$$

We expect the bilateral trade and trade openness indicators to be positively correlated with business cycle correlation.

Trade specialisation is measured by the cross-country difference between the average share across time of a particular sector in total exports. To obtain an overall sectoral distance measure for total exports, we add up the distances calculated for all sectors:

$$TRADEPAT_{ij} = \sum\nolimits_{n=1}^{N} \left[\left(\frac{1}{T} \sum\nolimits_{t=1}^{T} e_{int} \right) - \left(\frac{1}{T} \sum\nolimits_{t=1}^{T} e_{jnt} \right) \right]$$

where e_{int} stands for the share of sector n in total exports of country i, at time t. For instance, the share of the chemical sector in Belgium's overall exports is first averaged over the number of annual observations, then subtracted from the average chemicals share of, say Greece's total exports. This gives the economic "distance" between the two countries for the trade in the chemical sector. Total exports of a country are divided into the ten first-digit sub-sectors of the United Nation's Standard International Trade Classification (SITC), revision 2. These sub-sectors are (i) food and live animals, (ii) beverages and tobacco, (iii) crude materials, inedible, except fuels, (iv) mineral fuels, lubricants and related materials, (v) animal and vegetable oils, fats and waxes, (vi) chemicals and related products, n.e.s., (vii) manufactured goods, (viii) machinery and transport equipment, (ix) miscellaneous manufactured articles, and (x) commodities and transactions not classified elsewhere in the SITC.[10] Differences in trade specialisation patterns should be negatively related to business cycle correlation.

*Economic specialisatio*n is defined along the same lines as trade specialisation, as the sum of the differences of sector shares in the national economies:

$$ECOPAT_{ij} = \sum\nolimits_{n=1}^{N} \left[\left(\frac{1}{T} \sum\nolimits_{t=1}^{T} s_{int} \right) - \left(\frac{1}{T} \sum\nolimits_{t=1}^{T} s_{jnt} \right) \right].$$

s_{int} now represents the share, in terms of total output, of sector n in country i, at time t. Intuitively, we would expect a larger distance in economic patterns to have a negative impact on business cycle synchronisation. Hence we expect a negative coefficient for this variable, as for differences in trade specialisation. National value added divides into six sub-sectors, based on the International Standard Industrial Classification

[10] The data source is the NBER World Trade Flows Database, as documented in Feenstra and Lipsey (2005). We calculate the average over the years 1980, 1989, and 2000. Luxembourg is not covered by this dataset.

(ISIC): (i) agriculture, hunting, forestry, and fishing, (ii) industry including energy, (iii) construction, (iv) wholesale and retail trade, (v) financial intermediation and real estate, and (vi) other services.[11] Ideally we would have needed to use a more detailed decomposition of value-added in order to construct indices representing product-differentiation. A comprehensive data for more detailed sectors of the economy was unfortunately not readily available for all countries over the entire sample.

There is a variety of strategies of how to measure *financial integration*. A recent ECB survey on financial integration indicators by Baele et al. (2004) identifies two major measurement categories. The first category comprises *price-based* measures. According to the law of one price, a financial market is completely integrated if all differences in asset prices and returns are eliminated which stem from the geographic origin of the assets. Hence, the degree of price-based financial integration is measured by interest rate spreads of comparable assets across countries. Other authors resort to the second major category, *quantity-based* measures. These include asset quantities and flows across countries and attempt to measure capital flows and cross-border listings among countries; hence, they can be regarded as measures of financial intensity.[12] One pitfall of price-based and of most quantity-based measures is the lack of bilateral, country-to-country information. Only Papapioannou (2005) explores actual bilateral flows between country pairs as a quantity-based measure, employing data on bank flows. In addition, Imbs (2006) uses bilateral asset holdings data which are, however, survey-based and limited in their country coverage.[13] We adopt Papapioannou's approach and employ bilateral bank flows as a quantity-based proxy of country-to-country flows. We are aware that bank flows are an imperfect measure of financial integration but we believe that the

[11] The ISIC dataset includes all twelve euro area countries but the data period is limited to 1980-2003.

[12] See, for example, the financial integration studies by Imbs (2004), Kose et al. (2003), Lane and Milesi-Ferretti (2005); in addition to price-based and quantity-based measures, Baele et al. (2004) define a third, specialised category, news-based measures, which we neglect here.

[13] The data used by Imbs (2006) stem from the IMF's Coordinated Portfolio Investment Survey (CPIS) and apply to 2001 only; a number of financially important countries are not covered, such as Germany and Luxembourg.

bilateral characteristic of the bank flows suits particularly well to our econometric set up of country pairs.

We use as a proxy bilateral bank flows data provided by Papaioannou (2005). The source of the data is the BIS International Locational Banking Statistics. The aggregate bank flows are defined as the change in international financial claims of a bank resident in a given country vis-à-vis the banking and non-banking sectors in another country. The asset and liability flows are adjusted for exchange rate movements. Although similar, these two sets of series are not strictly equivalent. Asset flows from country i to country j are the assets held by banks in country i on all sectors in country j. They are not exactly the opposite of liabilities from country j to country i, since that variable represents the liabilities of banks in country j on all sectors in country i. The pair-wise series is calculated by taking the log of the average sum of bilateral asset (liability) flows between two countries.[14] The bilateral averages express a measure of financial intensity, regardless of whether flows occur in one direction or in the other. Hence, the log-bank flows of assets (LBFA) and of liabilities (LBFL) is expressed as

$$LBFA_{ij} = \left| \frac{1}{T} \sum_{t=1}^{T} \log \left(a_{ijt} + a_{jit} \right) \right|, LBFL_{ij} = \left| \frac{1}{T} \sum_{t=1}^{T} \log \left(l_{ijt} + l_{jit} \right) \right|,$$

with a_{ijt} as the change in assets of a country i bank towards all sectors in country j, at time t and l_{ijt} as the change in liabilities of a country i bank towards all sectors in country j, at time t.[15] The more intensive bank flows between two countries, the stronger we expect the correlation between their business cycles to be.

[14] Since the dependent variable, business cycle synchronisation, is by definition a ratio and all the other explanatory variables are either ratios themselves or are expressed as ratios, it is possible to compare the logarithm of financial flows to the other variables.

[15] The bank flows dataset generally covers the years 1980-2002. Some country series are, however, incomplete. Data for Luxembourg starts only in 1985, Portuguese data are available only from 1997. Greece's data are missing.

Policy and structural indicators relevant in the context of EMU

We consider short-term *real interest rate differentials*, in order to determine whether differences in the monetary policy stance can be related to business cycle synchronisation.[16] In theory, the direction of the effect is ambiguous. On the one hand, monetary policy shocks are one source of business cycles, and hence countries with a similar policy stance may react in a similar way or stand at around the same point of the business cycle. In this case, we would expect smaller interest rate differentials to be associated with larger cycle correlations. On the other hand, we can think of a reverse effect: if the economies were hit by asymmetric external shocks, business cycles may be less correlated due to the inability to respond by individual monetary policy in the presence of policy coordination. Then we would see small interest rate differentials corresponding to small cycle correlations. The same argument holds true for fiscal policy which we specify below. Therefore, the direction of the effect is ultimately an empirical one. To proxy the monetary policy stance, we use short-term three-month money market rates deflated by consumer prices (private consumption deflator), and take the absolute value of the mean sample of pair-wise differences:

$$IRSCDIFF_{ij} = \left| \frac{1}{T} \sum\nolimits_{t=1}^{T} (r_{it} - r_{jt}) \right|,$$

where r_{it} and r_{jt} represent the short-term real interest rates of countries i and j at time t.[17]

Nominal exchange rate fluctuations played a major role in the convergence process

[16] We employ the real and not the nominal interest rate for the following reason. Although a central bank sets the nominal interest rate, it does so taking the actual inflation rate into account, in order to achieve a certain real interest rate level. For household and firm decision-making, the real interest rate is the decisive number. In fact, in the presence of high inflation rates, a large nominal interest rate contains no information per se whether the central bank's monetary policy stance is effectively contractive or expansionary. Although we are aware that real interest rate differentials can also be seen as financial integration indicators, we presently focus on their characteristic as monetary policy measures.

[17] The interest rates dataset ranges from 1980-2004, except for Portugal where the series starts only in 1985.

prior to 1999. Exchange rate volatility should be negatively correlated with business cycle synchronisation. To capture the effect of variations in nominal exchange rates on business cycle synchronisation, we use the standard deviations of the bilateral nominal exchange rates between countries i and j across time t, $\sigma(E_{ijt})$, calculated via the ECU exchange rates. The standard deviations are scaled by the mean of the bilateral exchange rates over the sample period and can be written as

$$SD\_NERE_{ij} = \frac{\sigma(E_{ijt})}{\frac{1}{T}\sum_{t=1}^{T} E_{ijt}}.$$

Another convergence measure is given by the *fiscal deficit differentials*. As in the case of monetary policy, the effect of similar fiscal policy is unclear from a theoretical point of view. Two countries with a small difference in their general government balance may exhibit more or less similar business cycles. To explore this question empirically, we use net borrowing or net lending as a percentage of GDP at market prices of countries i and j at time t, d_{it} and d_{jt}, as defined by the European Commission's excessive deficit procedure. The variable is constructed as the mean sample of the bilateral differences of deficit ratios, and taken as the absolute value:

$$DEFDIFF_{ij} = \left| \frac{1}{T} \sum_{t=1}^{T} (d_{it} - d_{jt}) \right|.$$

As a *national competitiveness indicator (NCI)*, we use real effective exchange rates, weighted by intra-euro area trade partners and deflated by the HICP. Since the introduction of the euro in 1999, real effective exchange rates measure competitiveness based on relative price levels. As a distance measure, we compute the bilateral differences between countries i and j at time t and take the absolute value of the sample mean.

$$NCIDIFF_{ij} = \left| \frac{1}{T} \sum_{t=1}^{T} (nci_{it} - nci_{jt}) \right|.$$

The sectoral *stock market indicator* is built as the difference between stock market indices. We use the Datastream Total Market Index (TOTMK) and the Cyclical Services Index (CYSER).[18] To explore this finding in the context of cycle comovement, we expect a smaller cross-country difference in the stock market indices to be associated with more synchronised business cycles. We calculate country-pair differences in the values of these indices, scale them by national nominal GDPs and take the absolute value of the sample mean. Since the stock market indicators are expressed in terms of differences, we expect a negative relation with business cycle correlation. The corresponding equations read

$$TOTMKDIFF_{ij} = \left| \frac{1}{T} \sum\nolimits_{t=1}^{T} \left(\frac{totmk_{it} - totmk_{jt}}{y_{it} + y_{jt}} \right) \right|$$

and

$$CYSERDIFF_{ij} = \left| \frac{1}{T} \sum\nolimits_{t=1}^{T} \left(\frac{cyser_{it} - cyser_{jt}}{y_{it} + y_{jt}} \right) \right|.$$

Labour market flexibility indicators may play a role in the process of business cycle synchronisation. The more similar two countries are in terms of labour market flexibility, the more similar we expect their adjustment to shocks, leading to smoother cycles and less idiosyncracy. We employ two indicators from the OECD Labour Market Statistics. The first indicator is trade union density, measured as the percentage of organised workers in percent. We calculate the average over the sample and compute the bilateral differences in order to obtain a distance measure expressed in absolute value.[19] The second indicator is the OECD index of strictness of employment protection legislation. This index ranges from 0 (no protection) to 5 (strict protection) and is given for both permanent and temporary employment. We calculate the average of the permanent and temporary employment protection indices. Since data is available only for the years

[18] The CYSER index includes retail firms, hotel chains, media corporations and transports (such as airlines and railroads). Data are available from 1980-2004 except for Greece (1989-2004), Spain (1988-2004), Luxembourg (1993-2004), Portugal (1991-2004) and Finland (1989-2004). We also test other sectoral indices but report only those that deliver meaningful results.

[19] Trade union density data are available for all countries but only for the years 1980-2001.

1990, 1998, and 2003, we average these values for each country before we compute the bilateral differences as our distance measure of employment protection.

Finally, we apply *gravity variables* that are commonly used in the literature to account for exogenous aspects. Bilateral trade flows have been well explained by the "gravity" measures of geographical distance and relative size. Geographical distance is expressed in terms of distance between national capitals, in 1000 kilometre units.[20] Relative size is measured as the average of the bilateral difference in population between two countries, divided by the sum of their population.[21] The greater the distance, the smaller the expected correlation of business cycles.

3.1.3 Stylised facts of cross-country developments in the euro area

Before estimating the extreme-bounds analysis, we explore some descriptive properties of the core variables.

First, we inspect the country-specific cycles graphically. Figure 3.2 illustrates the cyclical parts of the annual real GDP series of the 12 euro area countries, scaled by overall GDP. All series exhibit the boom in the late 1980s and early 1990s, followed by a downturn. The German series reveals the 1990 unification boom and the successive period of high interest rates. This pattern seems to have spilled over particularly to France, Ireland, Italy, and Portugal. The Finnish series exhibits the strongest downturn of about 8 percent in magnitude, amplified by the breakdown of the Soviet Union in 1991. Apart from this exception, all cycles move within a band of ± 3 percent. The remainder of this sub-section further investigates the properties of the core bilateral variables, namely business cycle correlation, trade and specialisation.

[20] For Germany, the distance refers to Bonn, the capital of former West Germany. This makes sense economically because Bonn is located closer to Germany's main industrial areas than remote Berlin.

[21] We use population and not GDP to measure relative country size because GDP is already included in our left-hand side variable, i.e. business cycle synchronisation.

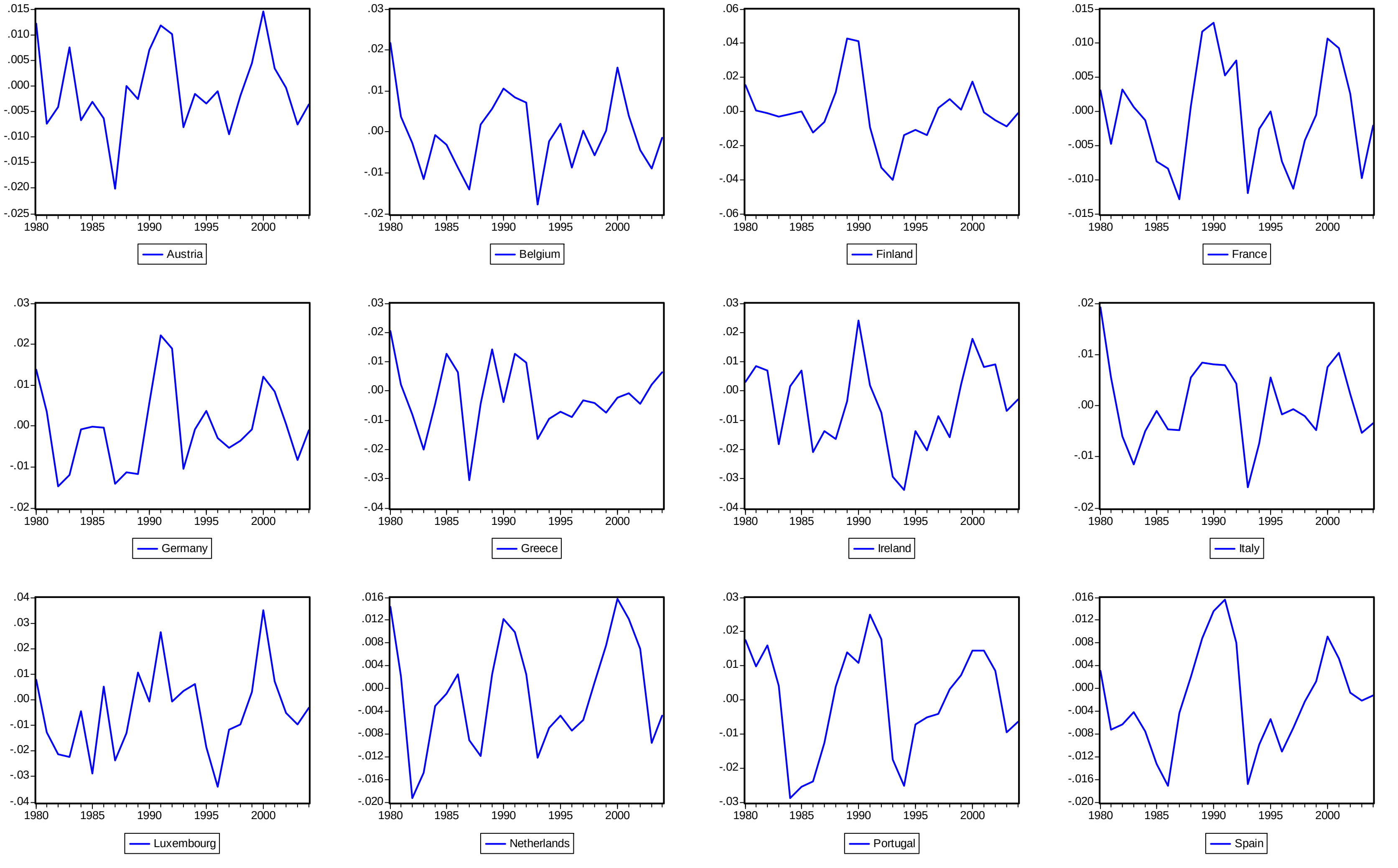


Figure 3.2: Business cycles of the 12 euro area countries (Baxter-King-filtered cyclical GDP components, scaled by overall GDP), 1980-2004.

Correlation of business cycles

Forming country-by-country pairs delivers 66 bilateral combinations. Figure 3.3 presents the largest and smallest ten coefficients of bilateral cycle correlation. Surprisingly, the largest correlation coefficient applies to Belgium-Italy, amounting to 0.85. The remaining top ten coefficients appear more intuitive, including neighbouring countries such as Spain-Portugal, Belgium-France, Germany-Austria or Germany-Netherlands.

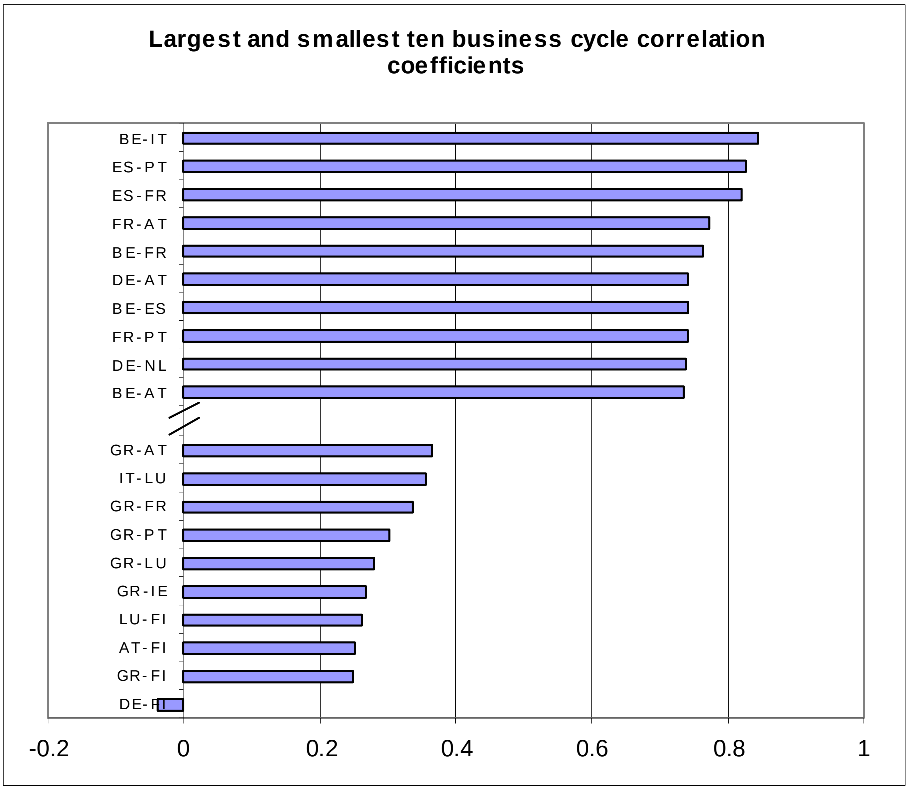


Figure 3.3: Largest and smallest business cycle correlation coefficients among the 66 euro area country pairs, 1980-2004.

The ten combinations with the smallest coefficients are often, although not always, between countries that are separated by a large geographical distance. This confirms the importance of geographical distance in the literature explaining differences in business cycles, as well as the need to include geographical distance as a control variable in regressions, provided it does not overlap with other explanatory variables. With a negative value that differs significantly from that of other country pairs, the Germany-Finland country pair stands out. The negative correlation is due to a one-off event. The German and Finnish economies were affected asymmetrically by the same external shock, namely the breakdown of the Communist regimes in Europe. Germany's unification boom peaked when Finland's cycle was already bust due to the collapse of the Soviet Union, one of its main trading partners.

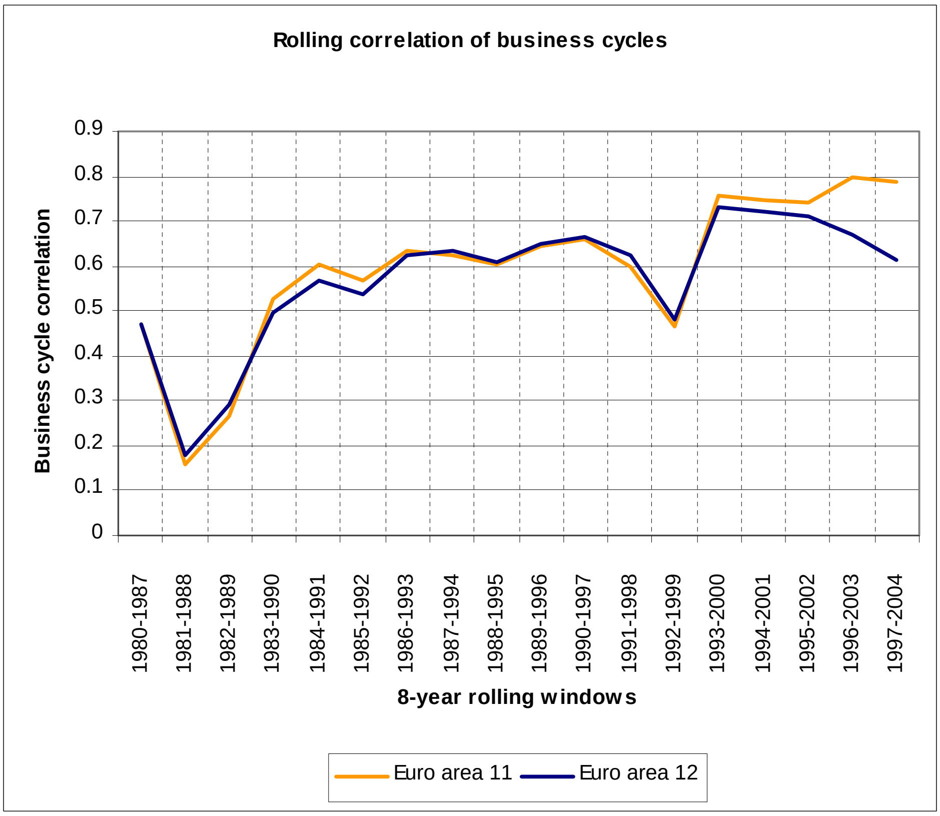


Figure 3.4: Average correlation coefficients of euro area business cycles (euro area 11 excludes Greece), 1980-2004, 8-year rolling windows.

Turning to time-varying aspects, we present rolling windows and sub-samples of the cycle correlations. Figure 3.4 illustrates the average correlations of the 66 country combinations in rolling windows. We choose 8-year windows corresponding to the maximum length of the business cycle in the Baxter-King filter which we applied to de-trend the real GDP series. The average correlation reaches a minimum of 0.18 in the period 1981-1988 before it increases in the late 1980s and early 1990s. It peaks in the period of 1993-2000 with a coefficient of 0.73 before declining to 0.62 in the most recent period, from 1997 to 2004. Excluding Greece however, the correlation of business cycles continued to increase after 1993 up to the most recent period, as illustrated by the euro area-11 line[22].

To analyse the background of the correlation variation over time, we divide the sample into three sub-samples, namely (i) 1980-1988, (ii) 1989-1996, and (iii) 1997-2004. Sub-samples of smaller size than eight years would be less likely to capture a full business cycle. In addition, the three periods broadly capture the successive stages of European integration. Economic and financial integration gained momentum in the late 80s and early 90 with the completion of the Single European Act in 1992 and later with the Treaty on the European Union of Maastricht. The third sub-sample can be regarded as the period of EMU, plus a two-year anticipation period. While the single monetary policy came into force in 1999, the definite timetable for its implementation gained credibility after the agreement on the Stability and Growth Pact in June 1997. Empirical studies have confirmed 1997 as the start of the convergence process towards monetary union.[23]

Figure A.1 in appendix A illustrates the average bilateral cycle correlations for the entire sample as well as for the three sub-samples. Given the overall average correlation of 0.57, the sub-sample value increased markedly from 0.42 in (i) to 0.65 in (ii). Period (iii) is characterised by a slight decrease to a correlation coefficient of 0.62. The latter

[22] We note that, due to the detrending filter, the cycle data may exhibit a certain degree of instability at the beginning and end points.

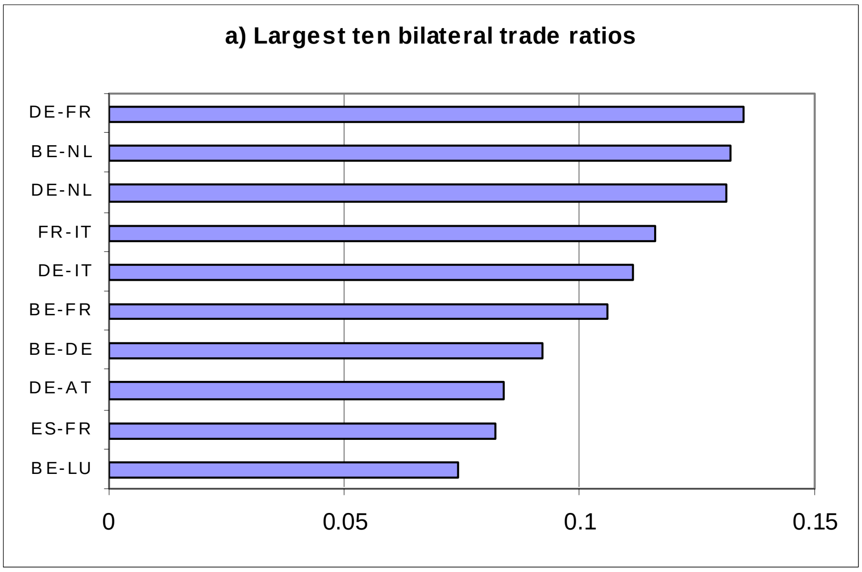
[23] See Frankel (2005) who considers June 1997 as the "breakpoint in perceptions"; according to Goldman Sachs estimations, the expectations of EMU taking place in 1999 shot up to a probability of 75%.

result becomes clear when looking at the largest and smallest ten coefficients for the three sub-samples, presented in figures A.2-4. While the presence of some minor negative coefficients is not surprising for period (i), we see a different picture in period (ii). Now, only the country pair Germany-Finland displays a negative coefficient, for the above-mentioned reasons. In period (iii), however, a large number of negative coefficients re-emerges. In fact, all of these negative values involve Greece.

The fall in the average correlation during the period of preparation for EMU and since Monetary Union is entirely due to specific developments in Greece. Excluding Greece, cross-country correlation coefficients indicate that EMU has been characterised by a greater synchronisation of business cycles among the other 11 euro area countries. The cross-country correlation of business cycles averaged 0.79 from 1997 to 2004, which was higher both than during the previous 1989-96 period (0.65) and than in the full sample (0.60).

Trade

Figure 3.5 illustrates bilateral trade ratios, scaled by total trade. The largest ratios correspond to the well-known examples of trade-integrated country pairs Germany-France, Belgium-Netherlands, and Germany-Netherlands. For instance, trade between Germany and France amounted to an average of 13.5 percent of their overall total trade over the period 1980-2004. Among the smallest ratios, we again find either Greece or Luxembourg in most of the pairs, confirming their special position among the euro area member states. Both countries have strong service sectors which are not captured by the merchandise trade measures.



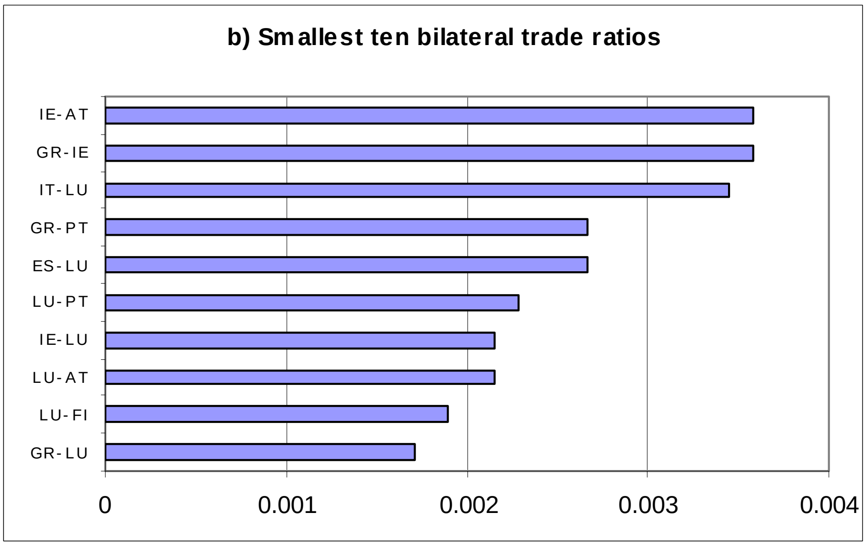


Figure 3.5: Largest and smallest ratios of bilateral trade ratios, scaled by total trade, 1980-2004.

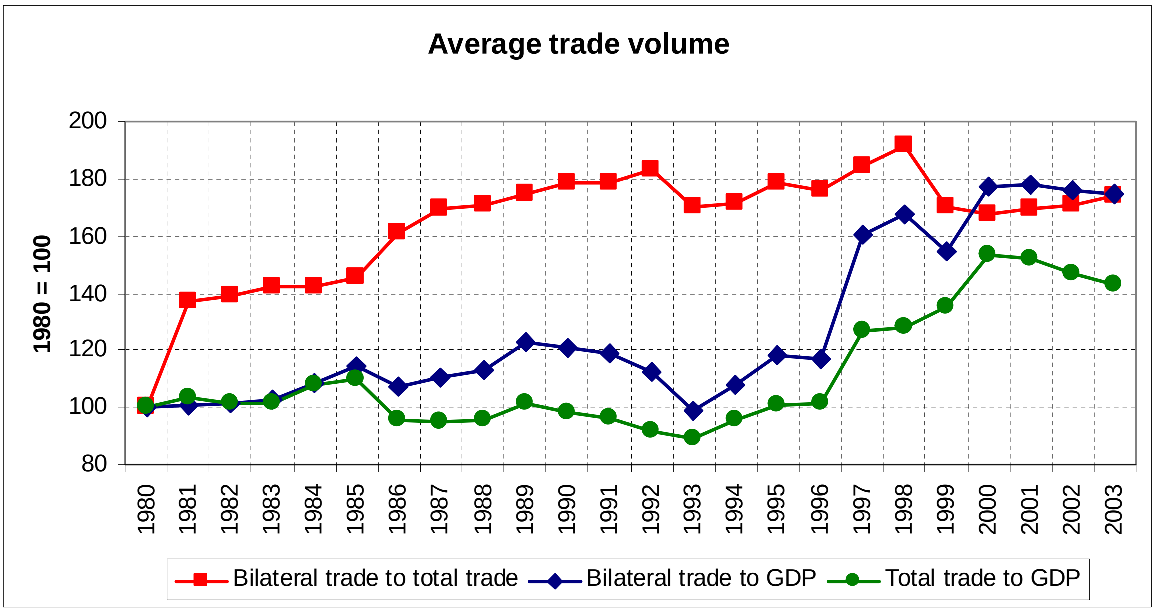


Figure 3.6: Average trade ratios, 1980-2003, scaled to 1980 = 100.

Inspecting the average trade ratios over time in figure 3.6, we note a sharp increase of bilateral trade, scaled by total trade, during the 1980s.[24] This may be partly due to the decrease of total trade which serves as the denominator in that bilateral trade ratio. At the same time, bilateral trade between the euro area countries in relation to GDP increased moderately. We take this as an indication of a trade diversion effect since trade with non-euro area countries seems to have gone down whereas intra-euro area trade has increased relatively. It is likely that this developement was spurred by intensified European economic integration in the late 1980s in the form of the Single Market programme and exchange rate coordination. During the late 1990s, we observe a sharp rise in bilateral trade, scaled to GDP, among euro area countries. Total trade has gone up as well which may have caused the ratio of bilateral trade to total trade to fall slightly. It seems that during the period of preparation and launch of monetary union, not only bilateral trade among the participating countries increased substantially

[24] Total trade refers to trade with the rest of the world, including euro area and non-euro area countries.

but also trade with the rest of the world. Although we recognise the contribution of other factors, it seems that the trade diversion effect turned into a trade creation effect so that trade increase among the euro area countries evolves no longer at the expense of third-country trade but rather in addition.[25]

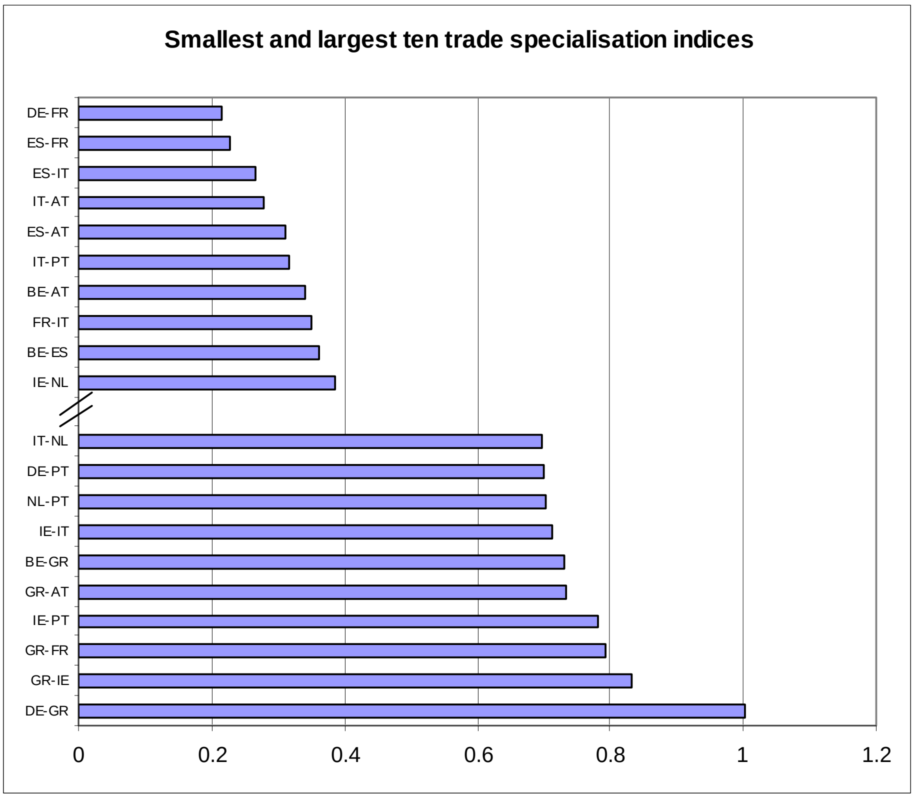


Figure 3.7: Smallest and largest indices of trade specialisation differences.

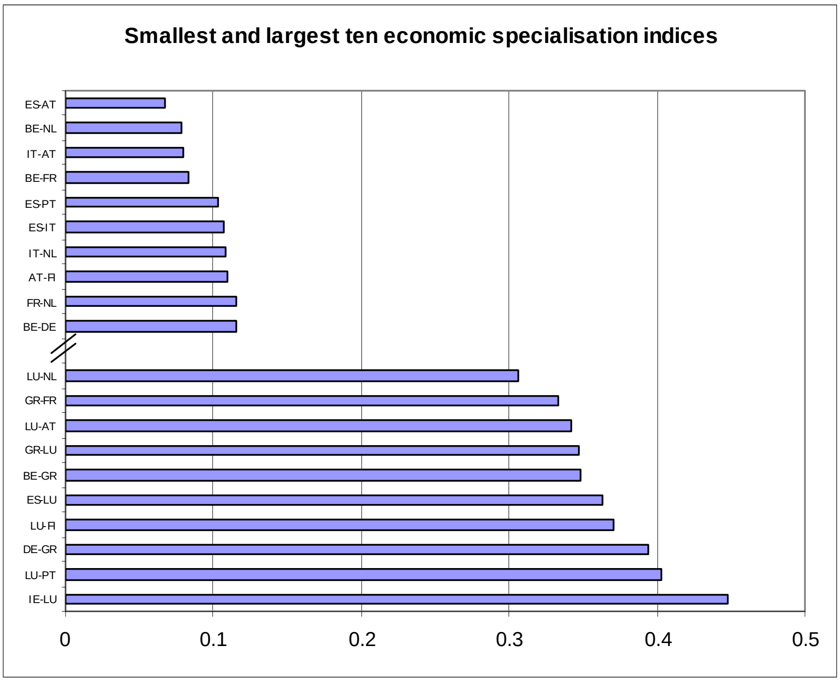
Regarding trade structure, the trade specialisation indicator reflects the cross-country differences in ten export sectors and thus focuses explicitly on tradables. The smallest and largest ten values are shown in figure 3.7, with small values indicating a low degree of specialisation differences, whereas large values stand for very different specialisation patterns. In other words, a small trade specialisation value indicates a high degree of

[25] This argument finds empirical support in Micco et al. (2003). For an overview, see Baldwin (2005).

intra-industry trade between two countries while country pair with a large index trades mostly *inter*-industry. The lowest trade specialisation position is taken by Germany-France which is often quoted as the classical example of intra-industry trade. Hence, these two countries do not only trade most with each other as indicated by the bilateral trade ratios, they also trade most in similar sectors. The most different country pairs involve Greece in six out of ten values. Greek exports exhibit markedly larger shares of trade in food and beverages while the exports of Greece are at the same time characterised by smaller shares of machinery and transport equipment than that of most other euro area countries. Luxembourg does not appear because of data unavailability.

Figure 3.8: Average indices of trade specialisation differences.

Across time, euro area countries have converged in terms of trade specialisation as shown in figure 3.8. From 1980 through 2000, differences in trade specialisation declined. The trade specialisation measures indicate that euro area countries have become more similar in terms of trade structure. Combined with the above evidence that EMU contributed to trade creation, this provides an indication that the intensification of trade relations due to the single currency was characterised by the development of *intra*-industry trade by opposition to *inter*-industry trade. Thus, as conjectured by Frankel and Rose (1998), the introduction of the single currency may have given a substantial impetus for trade expansion.



3.9: Smallest and largest indices of economic specialisation differences.

Economic specialisation

Furthermore, we consider bilateral economic specialisation indices across six sub-sectors of the economy. Again, a small value indicates a small specialisation difference, i.e. large similarity in the share of economic sectors in value-added. A large index value, in turn, stands for highly different sectoral shares across countries. In general, we expect small values for specialisation to be associated with large coefficients of cycle correlation.

Figure 3.9 presents the smallest and largest ten economic specialisation indices. Spain and Austria share the most similar economic structure as indicated by the small value of the specialisation index. Although this result may appear surprising at first sight, it does not reflect an actual product specialisation. The small index means that the shares of industry, construction, wholesale and retail trade and financial services are similar in the Spanish and Austrian economies. While this seems like a lot of similarity, the product specialisation — in particular in tradable goods and services — may differ considerably. Other country-pairs are less unexpected, such as Belgium-Netherlands, or Spain-Portugal.

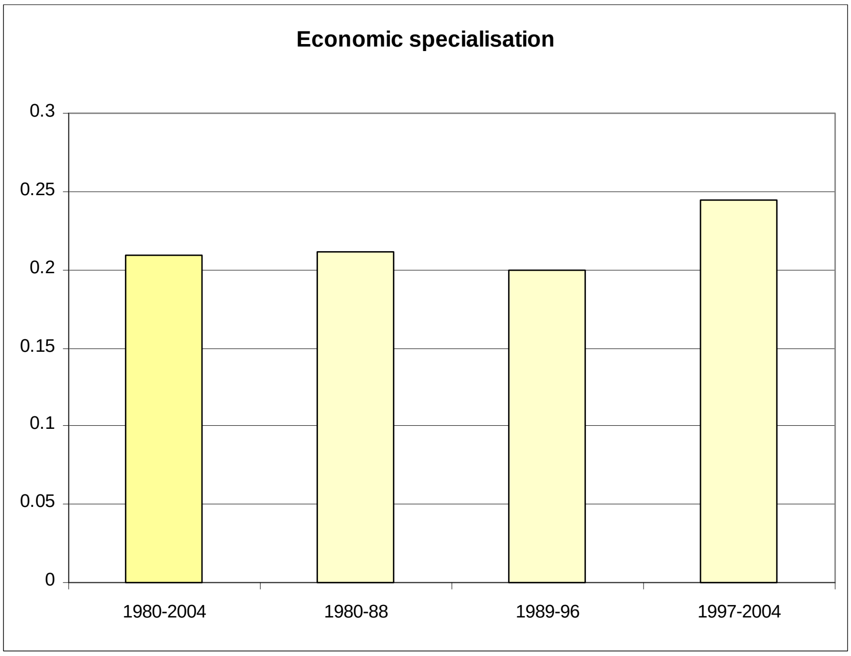


Figure 3.10: Average indices of economic specialisation differences.

Analysing the countries with the most different structures, it strikes that again either Greece or Luxembourg are involved in each of the pairs. In this case, Luxembourg's large financial service sector gives rise to larger values in overall economic specialisation differences. Greece stands out with a fairly large agricultural and rather small industrial sector.

Over time, cross-country differences in terms of broad economic specialisation have remained fairly stable during the 1980 and early 1990s, as figure 3.10 illustrates.Since 1997, we observe a modest increase. The ECB (2004) report on sectoral specialisation comes to a similar result and attributes a slight increase in specialisation for some smaller euro area countries to developments in business sector services. Furthermore, an analysis by the European Commission (2004: 149) matches our results in observing that "the specialisation of production has gradually increased [...] while export specialisation has decreased." While this finding appears puzzling at first glance, the Commission argues that production adjusts more slowly than trade. Also, it supports the notion of increased intra-industry trade measured by a rising Grubel-Lloyd index between 1980 and 2001. Given that trade in similar industries is a key channel of spill-overs across countries, we expect trade specialisation, more than economic specialisation, to play a key role in the synchronisation of business cycles.

Bank flows

Bilateral bank flows are presented in figure 3.11, again for the largest and smallest ten values. The country pair Germany-Luxembourg ranks top and reflects, on the one hand, the capital-strong position of Germany, and on the other, the outstanding importance of Luxembourg's financial service industry. Among the smallest values, Finland seems to have had a particularly low integration with the euro area countries over the past 25 years. Figure 3.12 illustrates how average bank flows evolved across the three sub-periods. It is obvious that the average bank asset flows increased steadily over time across euro area countries which is in line with increasing capital market liberalisation.

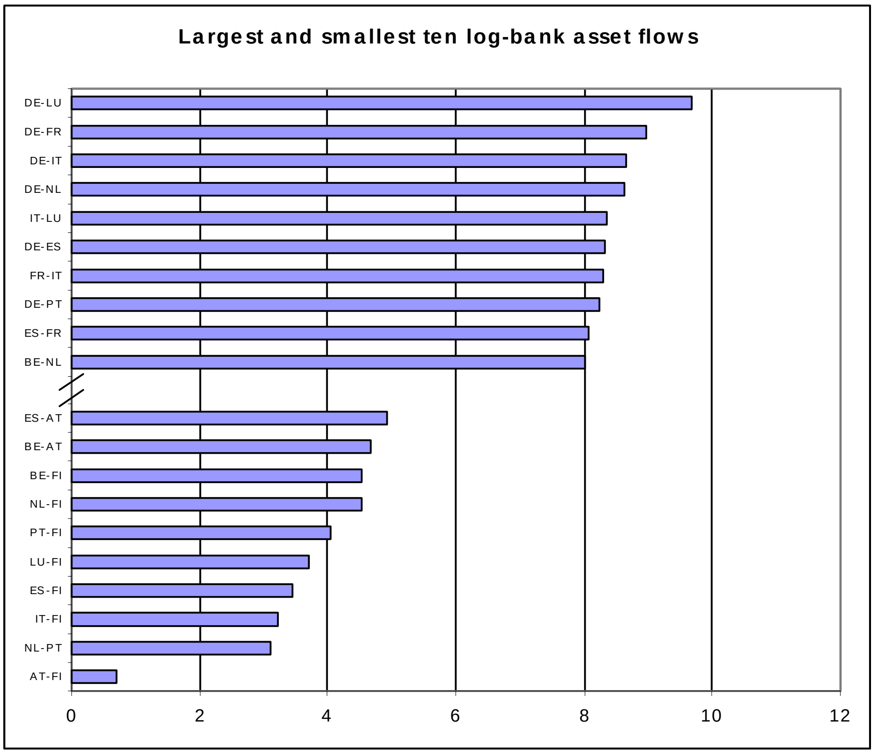


Figure 3.11: Largest and smallest bilateral bank flow indicators (assets, in logs).

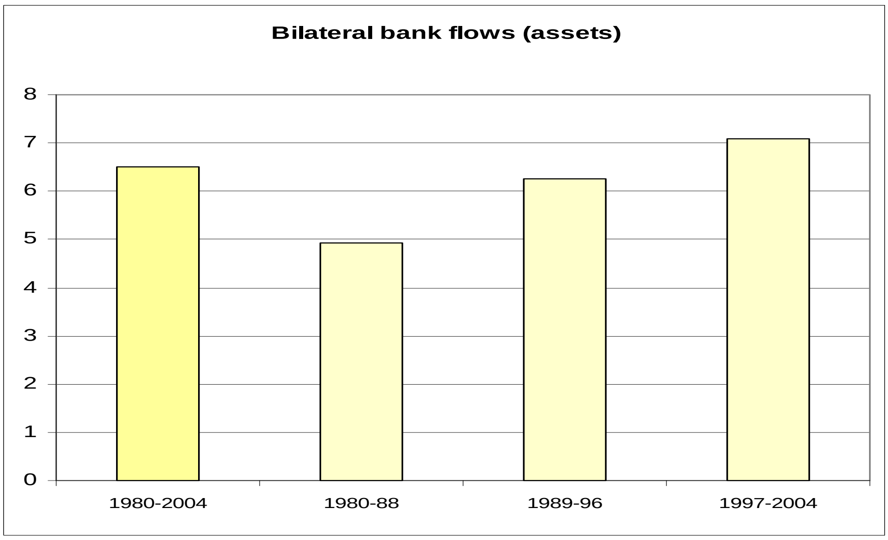


Figure 3.12: Average bilateral bank flows (assets, in logs).

3.2 A "robust" estimation approach: The extreme-bounds analysis

In this section, we introduce the econometric methodology and present the main results of the analysis of the determinants of business cycle synchronisation across euro area countries.

3.2.1 Methodology

To identify the key determinants of business cycle synchronisation in the euro area, we employ the extreme-bounds analysis (EBA) as proposed by Leamer and Leonard (1981), Leamer (1983) and further developed by Levine and Renelt (1992), Levine and Zervos (1993), and Sala-i-Martin (1997) in the context of empirical growth analysis. Baxter and Kouparitsas (2004) employ an EBA estimation to explain business cycle synchronisation across a large sample of developing and industrialised countries.

Estimation framework

In empirical studies, the researcher is often faced with the decision which determinants to include in an analysis. Sometimes, various possible regression set-ups have equal theoretical status but the resulting coefficients may depend heavily on the set of control variables employed. Hence, the choice of right-hand side variables is often based on assumptions and, in the end, left to the researcher's discretion.[26] This dilemma, which Brock and Durlauf (2001) refer to as the "open-endedness of theories", may result in incomplete econometric models suffering from specification bias.

The EBA framework is one attempt to respond to this dilemma by considering a large number of alternative specifications and filtering out those determinants that do not turn insignificant with the alteration of the conditioning set of information. In this sense of

[26] See Durham (2001) and Levine and Renelt (1992).

robustness, the significance of the "robust" determinants cannot be eliminated by any other variable. Otherwise, the variable is considered "fragile" even if it is significant in a bivariate or in some multivariate set-ups.

In practice, the robustness of the potential determinants is investigated by testing each candidate variable (M-variable) against a varying set of other conditioning variables (Z-variables). A necessary condition for a variable to be a meaningful determinant of business cycle correlation is that it should be significant in a bivariate regression. Its explanatory power may however vary considerably when other determinants are added to the baseline regression. The basic equation can be expressed as

$$Y = \beta_i I + \beta_m M + \beta_z Z + u, \tag{3.1}$$

where Y denotes a vector of coefficients of bilateral business cycle correlations. The M-variable is the candidate variable of interest which is tested for robustness. This robustness test is conducted by including a varying set of conditioning or control variables, Z, and checking β_m's sensitivity to alterations in Z. For each M-variable, we first run a baseline regression without any Z-variables, then successively include one, two, and three Z-variables in every possible combination.[27] The I-variable, on the other hand, controls for initial conditions that are exogenous. The "gravity variables", geographical distance and relative population size, fall into that group. We run alternative set-ups with and without the I-variables.

For every M-variable under consideration, the EBA identifies the "extreme bounds" by constructing the highest and lowest values of confidence intervals of the estimated β_m coefficients. In other words, the *extreme upper bound (EUB)* is equal to the maximum estimated β_m, plus two times its standard error,

[27] This strategy follows Levine and Zervos (1993).

$$EUB = \beta_m^{\max} + 2\sigma(\beta_m^{\max}),$$

the *extreme lower bound (ELB)* is the minimum estimated β_m, minus two times its standard error,

$$ELB = \beta_m^{\min} - 2\sigma(\beta_m^{\min}),$$

The M-variable is then regarded as robust, if the EUB and the ELB exhibit the same sign and if all estimated β_m coefficients are significant.

Leamer's standard methodology is based on OLS estimates. Estimates of the parameters in cross-section regressions are subject to sampling uncertainty and to correlations between sampling errors. Frankel and Rose (1998) and Imbs (2004) use the White correction for heteroskedasticity to account for possible sampling errors. Clark and Van Wincoop (2001) argue that this does not allow to correct for dependencies in the residuals and use GMM methods to calculate the variance-covariance matrix of the parameters. GMM nevertheless gives imprecise variance estimates in small samples and would therefore not have been appropriate in our case, given the relatively small size of our sample consisting in the 66 euro area country pairs. Instead, in order to get robust estimators for the coefficients of the candidate explanatory variables, we apply to the OLS regressions a Newey-West correction for heteroskedasticity and autocorrelation in the residuals which is less dependent on large sample properties.

The decision rule first outlined by Levine and Renelt (1992) was derived from the statistical theory expounded in Leamer and Leonard (1981). It has often been criticised for being too restrictive. In practice, an explanatory variable might fail to qualify for robustness because of one statistical outlier in one single equation. Using least absolute deviation (LAD) estimators to deal with potential outliers is, however, not an option for our study because LAD is particularly inappropriate in relatively small samples. Also, when compared with OLS, LAD is not a robust estimation method in the statistical sense of the word. It indeed requires extra assumptions for the estimation of conditional mean parameters that are not necessarily met in the actual population. Nevertheless, we consider two other criteria in addition to the decision rule defined by Levine and Renelt (1992).

The first additional criteria is the percentage of significant coefficients of the same sign. Sala-i-Martin (1997) argues that running a sufficiently large number of regressions increases the probability of reaching a non-robust result, pointing that "if one finds a single regression for which the sign of the coefficient β_m changes or becomes insignificant, then the variable is not robust."[28] He suggests to assign a certain "level of confidence" to each M-variable by investigating the share of significant β_m coefficients. An M-variable with a share of significant coefficients of 95% may be considered as "significantly correlated" with business cycle synchronisation. In the results tables, we therefore not only state the robust/fragile result but also indicate the share of significant coefficients.[29]

The second criteria we consider in the cases where one of the bounds changes sign, is whether the value of the extreme bound is large compared with the corresponding coefficients. In some cases, after adding (or subtracting) two standard deviations to the maximum (or minimum) estimated β_m coefficient, the extreme upper (or lower) bound changed sign but remained close to around zero while all β_m coefficients were significant and of the same sign. When the value of the upper (lower) bound was less than 5% the maximum (minimum) coefficient, we have considered that the variable was significant in explaining business cycle correlation.

These two criteria do not affect our fundamental results but allow to qualify the evidence in one or two limit cases.

[28] Sala-i-Martin (1997: 178)

[29] We state the share of significant coefficients only for the cases in which at least the bivariate estimation coefficient is significant.

Information set

The dependent variable is a vector of bilateral pairs containing the 66 correlation coefficients between the cyclical part of real GDP for the 12 euro area countries. The candidate explanatory variables are drawn from the set of potential determinants presented above. They include: bilateral trade, trade openness, trade patterns, economic patterns, bilateral bank flows, real short-term interest rate differentials, nominal exchange rate fluctuations, fiscal deficit differentials, national competitiveness indicators, differences in stock market indices, labour market flexibility indicators and gravity variables.

Among this set of indicators, we select four main categories of *M-variables of interest* which we think should be key determinants of the business cycle as indicated by the literature. These variables are: bilateral trade and openness to trade, trade specialisation, economic specialisation and bilateral bank flows. Regarding the group of Z-variables, we agree with the selection process used by Levine and Zervos (1993) and try to avoid including series that may overlap with the M-variable under review. This amounts to minimising multicollinearity problems between the explanatory variables which might be a drawback of the EBA analysis. For instance, a similar trade specialisation pattern between two countries may be related to strong intra-industry trade, which would result in an intensification of bilateral trade. The similarity of economic structures may also be reflected in the similarity of trade patterns. Strong trade relations may contribute to intensify the flow of credits between two countries. In addition, we test successively for different alternative measures of these M-variables.

The robustness of the M-variables was tested by estimating multivariate regressions where all possible combinations of 1 to 3 explanatory variables, drawn from a pool of six Z-variables and one I-variable, were added successively to the bivariate regression.

The core group of *control Z-variables* which may be related to the business cycle includes: bilateral exchange rate volatility (SD_NERE), differences in fiscal deficits (DEFDIFF), differences in national price competitiveness (NCIDIFF), differences in the performance of stock markets (TOTMKDIFF for the overall market index; alternatively CYSERDIFF for cyclical services), differences in trade union density (TUDDIFF). The employment protection indicator EPADIFF was not used in the multivariate regressions due to the lack of data and absence of significance in the bivariate regression. The Z-variables may also turn out to be potentially important explanatory variables and have also been identified, directly or indirectly, as key determinants of business cycle synchronisation.

To the group of initial Z-variables, we added the gravity variables which we first considered as *I-variables,* and which represent external non-economic factors. However, systematically including geographical distance (GEODIST) in all equations created partial correlation problems because several explanatory variables are closely related to geographical distance, bilateral trade in particular. As in Baxter and Kouparitsas (2004), we treated geographical distance as a "not-always" included variable. Including or not differences in population size (POPDIFF) as an I-variable did not make any difference to the EBA analysis. In the tables in appendix B we present the results of the EBA estimates without population differences because of the complete absence of significance of that variable in our estimates.

Robustness tests were conducted also for the variables which we designated ex-ante as Z-variables and I-variables. In order to ensure the comparability of results, the additional explanatory variables were always drawn from the same pool of explanatory variables,[30] as for the M-variables.

Samples

In the following sub-sections, for each group of possible explanatory variable, we present the bivariate relations with business cycle and discuss the EBA results. The robustness of the variables is tested for the full sample from 1980 to 2004. It is of particular interest

[30]BTT, TOTMKDIFF, IRSCDIFF, NCIDIFF, DEFDIFF, SD_NERE, TUDIFFF and GEODIST.

to know whether the determinants of business cycle correlation have changed since the implementation of a common monetary policy. We therefore conducted tests for two sub-periods. The first period runs from 1980 to 1996, the second period starts in 1997 and ends in 2004. For the above mentioned reasons, we consider the second period as the "EMU period".

Since the analysis is a cross-section analysis, across countries and for one point in time, the sample size for the estimates is always the same whatever the number of years in the period of estimation, and corresponds to the 66 country pairs. Since the series entering the regressions are calculated in terms of averages, the cross-country observations might be more dispersed when calculated over a shorter period of time than when calculated over a period of several years. This is not however the case: the standard deviations of the series scaled by their means are not always higher in the two sub-samples than in the full sample, and in the last sub-sample than the first one.

Regarding parameter uncertainty, the standard error of the coefficients tend to increase in the 1997-04 sample (see tables of results in appendix B) which could lead to more frequent rejection of robustness. However, there is no automatic link between the size of standard errors and the acceptation or rejection of robustness. The "robustness" of the explanatory variable is accepted also in the cases where the standard error of the explanatory variable's coefficient increases considerably in the third sample (for instance TRADEPAT in table B.3 or IRSCDIFF in table B.6 in appendix B).

3.2.2 Results for core explanatory variables

Bilateral trade and trade openness

Different measures of trade The three measures of trade are considered successively. For these variables we expect a positive coefficient: the more intensive trade between two countries (or the more open to trade), the higher the trade variable, and the more synchronous the business cycles. Business cycle correlation increases with the intensification of bilateral trade, both relative to total trade and to GDP. Through bilateral trade, spill-over effects appear to affect simultaneously business cycles in two countries regardless of their relative openness to trade.

The first measure, bilateral trade as a ratio to total trade (BTT), is plotted against business cycle correlation in figure 3.13. The vertical axis represents business cycle correlation and the horizontal axis the explanatory variable, the bilateral trade to total trade ratio in the present case. The plot shows the equation corresponding to the regression line and the associated R^2. The bivariate regression of business cycle correlation on bilateral trade reveals a positive-sloping trend. With a t-statistic of 3.9, the point estimate is significant at the 5% level. The goodness of fit amounts to 0.2 which appears acceptable for a bivariate regression. It is, however, clearly visible from the chart that the upward slope is generated by approximately a third of the observations while the remaining points form a cloud close to the vertical axis. The outlier with the negative correlation estimate pertains to the German-Finnish country pair as discussed above.

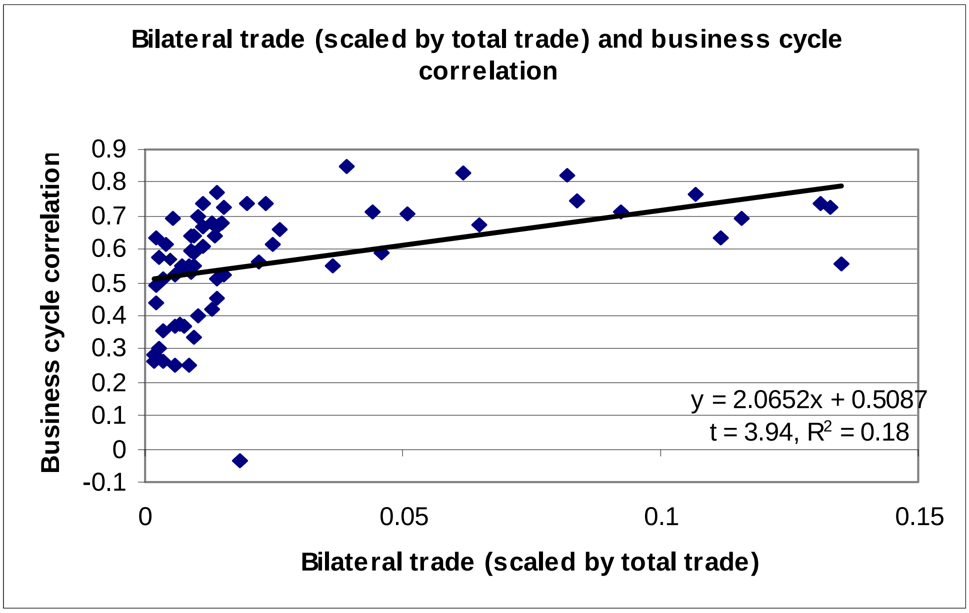


Figure 3.13

The plot of the second trade measure, bilateral trade to GDP (BTY), is shown in figure 3.14 and exhibits the same positive-sloping trend. The coefficient on BTY is also positive, the t-statistics significant at the 5% level, and the R^2 acceptable.

By contrast with BTT and BTY, the third trade measure, overall openness to trade (TTY), fails to be significant in a bivariate regression. Figure 3.15 indicates little connection between similarities in openness and cycle correlation. Since the total trade to GDP ratio is not significant in the bilateral regression and the first necessary condition is not fulfilled, we do not test that variable for the EBA.

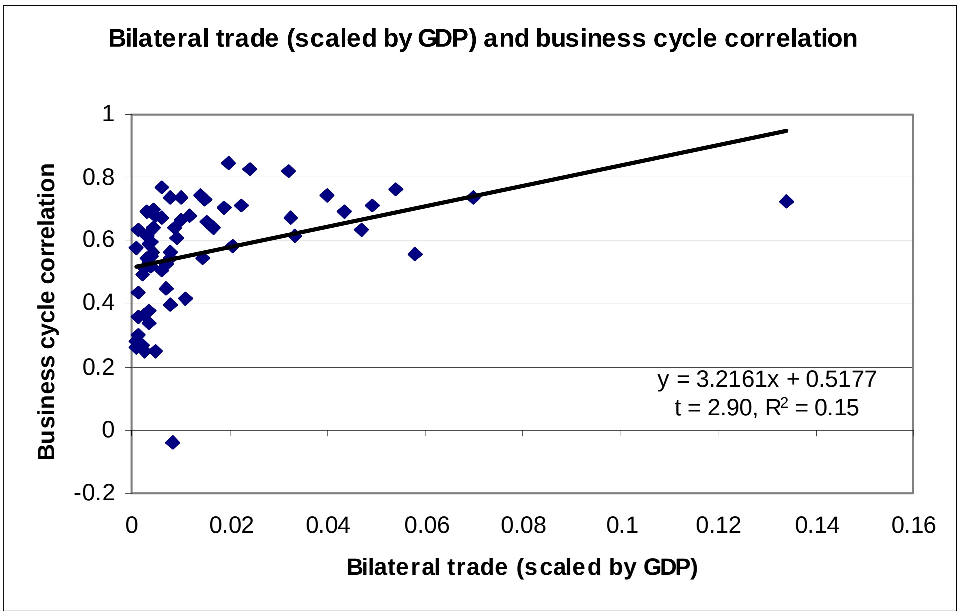


Figure 3.14

EBA results Over the full sample, both BTT and BTY come out clearly as robust, in the case of BTT including or not geographical distance, and in the case of BTY without geographical distance. The results are reported for the two variables without geographical distance.[31] For BTT, without geographical distance, the lower and upper bounds of all estimates range from 0.1 to 3.1. The β_m coefficients range between 1.0 and 2.1 and are all significant at the 5% level. Although the lower bound is close to zero, the associated equation has a fairly good explanatory power. Indeed, the associated R^2 reaches 0.4 and is twice as big as for the upper bound and as in the bivariate case. For BTY, also without geographical distance, both the extreme β_m coefficients and the extreme bounds tend to be higher than for BTT (from 1.5 to 3.2 for the extreme coefficients), probably because the BTT ratio tends to be lower than BTY. However, the explanatory power of BTY is not greater than that of BTT, as indicated by the similarity in the R^2s. Among the three Z-variables for which the lower bound is reached are the national competitiveness indicator and differences in fiscal deficits, both in the case of BTT and of BTY.

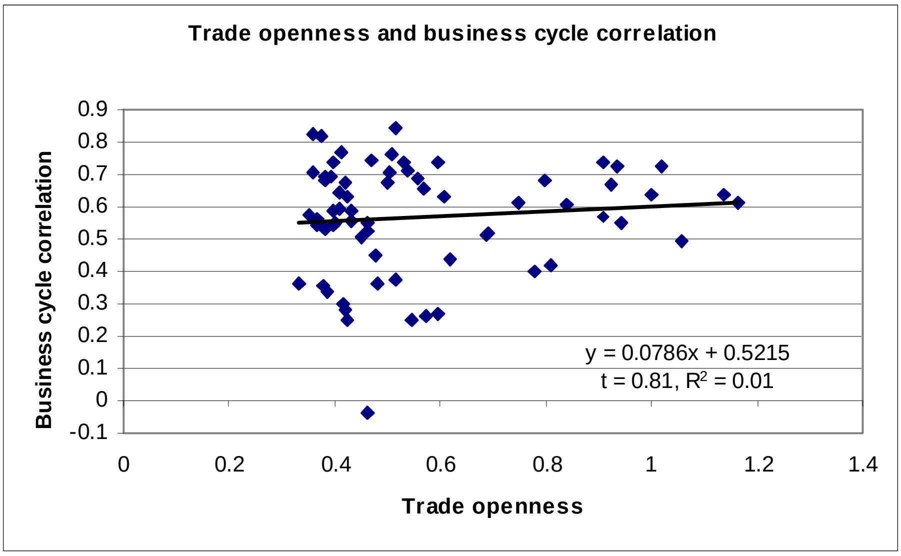


Figure 3.15

[31] In that particular case, geographical distance may create multicollinearity problems if included among the regressors. Geographical distance is indeed a strong determinant of bilateral trade itself.

Turning to the sub-samples, for the 1980-96 period, both BTT and BTY remain robust determinants of business cycle correlation. The range for the extreme bounds tends to be larger than for the full sample, due to larger standard errors. Nevertheless, the range for the actual β_m coefficients is smaller, indicating that the power of BTT and BTY to explain business cycle synchronisation is less conditioned by other variables than in the full sample. However the explanatory power of bilateral trade ratios for the 1980-1996 period is very low (the R^2s are around 0.1), indicating that bilateral trade explained only a small part of business cycle correlation.

While bilateral trade appears to have been a key element in the synchronisation of business cycles before monetary union, its importance to explain business cycle correlation has clearly decreased since then. For both BTT and BTY, over the 1997-2004 period, the lower bound turns clearly negative as the minimum β_m becomes insignificant in particular when the fiscal deficit differential are added as explanatory Z-variable. However, the upper bounds increase markedly. In the bivariate case and when only difference in trade union membership is added to the equation, the maximum β_m coefficients increase to 4.1 for BTT and to 5.9 for BTY .

Trade specialisation

The trade specialisation indicator (TRADEPAT) is presented in figure 3.16 where the expected negative relation to cycle correlation is confirmed. In other words, the more similar the trade structures of two countries, the higher is cycle correlation. The t-statistics amounts to -3.1, respectively and the R^2 is fairly large (0.2) for a bivariate regression.

EBA results

Over the full sample, trade specialisation fails to qualify as robust by only a small margin. All the coefficients have the expected negative sign and are significant at the 10% level but the upper bound turns positive in the case of the maximum coefficient (-0.2). The minimum coefficient (-0.4) is reached in the bivariate case and in the case with one Z-variable, difference in trade union density. Noticeably, bilateral exchange rate volatility when introduced in the estimate seems to reduce sensibly the explanatory power of trade specialisation.

As the case for trade specialisation is somewhat undetermined, we conducted tests replacing it with selected components: differences in the share in total trade of mineral fuels (CD_FUEL), machinery and transport equipment (CD_MACH), other manufacturing products (CD_MANU) and chemicals (CD_CHEM). These products were selected for their greater sensitivity to fluctuations in the business cycle. None of the four components comes out as a robust over the full sample but, with all the coefficients significant at the 10% level, trade in machinery and equipment comes very close to it. Machinery and equipment is indeed widely considered as a leading indicator of the business cycle, and a substantial part of intra-industry trade between euro area countries occurs in that sector.

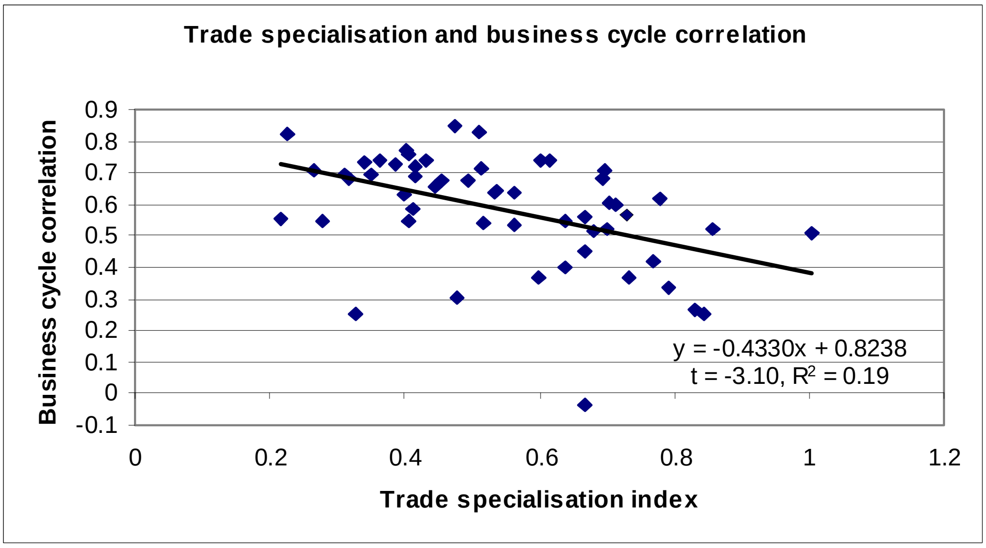


Figure 3.16

Over the 1980-1996 period, trade specialisation fails to qualify as robust. Even in the bivariate regression, the coefficient on trade specialisation remains insignificant. The upper bound which was more sensitive to changes in the information set in estimates for the full sample, becomes even more clearly insignificant when the national competitiveness indicator is included as a control variable. None of the components of trade specialisation qualifies as robust and not even as significant in the case of two Z-variables.

By contrast, trade specialisation becomes clearly robust in the 1997-2004 sample. The maximum and minimum β_m coefficients are all significant at the 5% level, ranging from -0.5 to -1.5 with fairly large R^2s (0.6 and 0.4, respectively). As for the full sample, most of the impact of trade specialisation on business cycle synchronisation seems to be driven by trade specialisation in machinery and transport equipment (CD_MACH). For that sector, the results are even more significant than for total trade, Importantly, the R^2s are very large, in particular in the case of the upper bound (0.8), including three Z-variables: the real interest rate differentials, the competitiveness indicator and differences in fiscal deficits.

Economic specialisation

The economic specialisation indicator (ECOPAT) is presented in figure 3.17. As for trade specialisation, the expected negative relation to cycle correlation is confirmed. Although the t-statistics on the coefficient is significant at the 5% level, the R^2 of the regression (0.05) is not meaningful. This suggests that an overall similarity in the relative shares of broad economic sectors provides little information to explain business cycle correlation.

EBA results Indeed, in the EBA analysis, economic specialisation fails to reach the robustness status with the extreme bounds ranging from 0.3 to -1.0. The upper bound becomes insignificant and of the wrong sign when the total stock market index, the fiscal deficit differentials and bilateral exchange rate volatility are included as control variables. As for trade specialisation, we also analysed the robustness of some of the components of economic specialisation: industry (CD_IND), construction (CD_CNT), wholesale and retail trade (CD_TRA), financial intermediation (CD_FIN). Out of the five sectors, only the differences between the share of industrial sectors (CD_IND) come out as significant, regardless of the combination of Z-variables included in the equation. In the full sample, from 1980 to 2004, all the β_m coefficients significant at the 5% level and negative, ranging from -1.2 to -2.2. The statistics presented in the tables in the appendix are based on short-term interest rates deflated by the GDP deflator. On a yearly basis, interest rate differentials deflated by the national GDP deflators or by the national consumption deflators differ little. Nevertheless in the case of industrial differences, the upper bound turned to the wrong positive sign by a very small margin (less than 5% of the absolute value of the extreme coefficients), when using interest rates deflated by consumer prices. When using differentials of interest rates deflated by the GDP deflator, they remained clearly negative. By comparison using either deflator did not make any difference to the results in the case of the other variables that were tested for robustness.

Regarding the 1980-96 sub-sample, economic specialisation fails again to qualify as robust but both the relative shares of industrial sectors (CD_IND) and the relative shares of financial sectors (CD_FIN) come close to robustness.[32] The relative importance of financial specialisation in explaining business cycle synchronisation over the first sub-sample may reflect the impact on economic activity of the liberalisation, development and internationalisation of financial services during that period. Even though all the β_m coefficients are again significant at the 5% level and of the right sign, the relative size of the industrial sector in value-added does not comes out as robust. Due to a marked increase in the standard errors of the estimated coefficients, the upper bound turns out very positive.

[32] Construction also appears as robust but with the wrong expected sign.

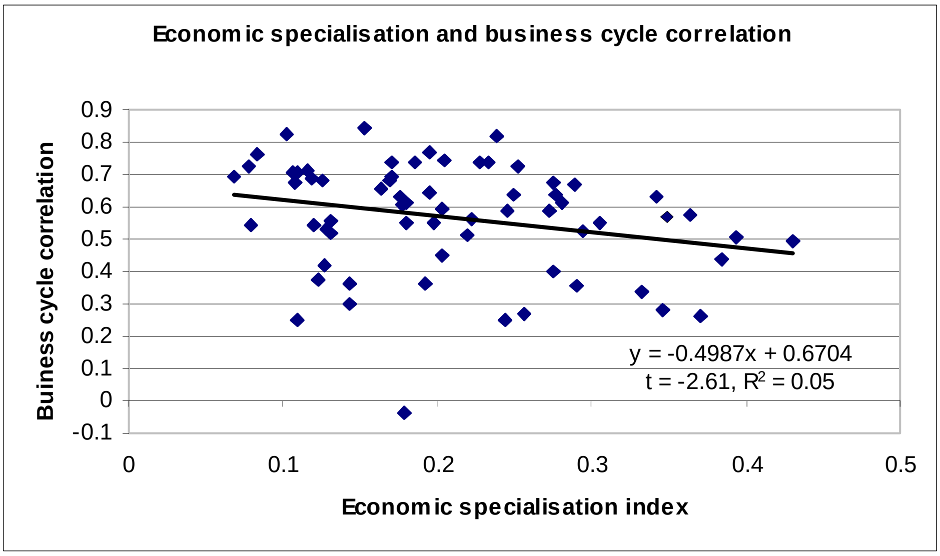


Figure 3.17

Over the 1997-2004 period, neither overall economic specialisation nor any of its components comes out as robust. In addition the β_m coefficients are insignificant and often of the wrong sign, even in the case of industrial and financial specialisations. Also, as for the full sample and for the previous sample, the explanatory power of economic specialisation appears limited as indicated by the fairly small R^2s.

As supposed in sub-section 2.3, the absence of clear-cut results for economic specialisation and its components might be due to the fact that the impact of economic specialisation on the business cycle would be better captured by a narrower breakdown of value-added, allowing to account for product-specialisation in tradable goods and services.

Bilateral bank flows

The measure of bank flows, log-bilateral flows of bank assets (LBFA), is plotted against business cycle correlation in figure 3.18. The slope of the regression line is positive (0.04) and significant at the 1% level with an R^2 of 0.2. This suggests that, on a bivariate basis, larger amounts of bilateral bank flows are associated with higher correlation of the business cycles.

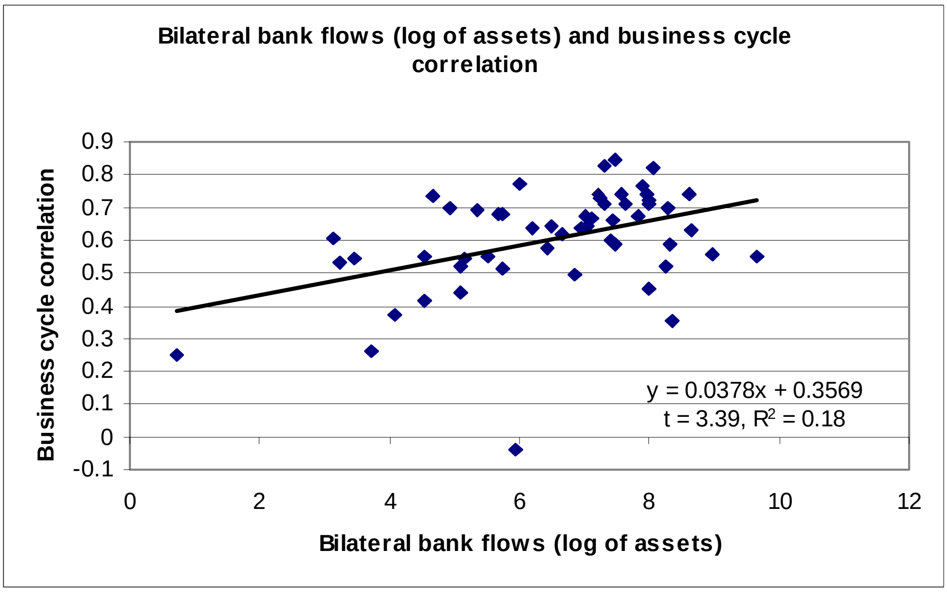


Figure 3.18

EBA results Over the full sample, bilateral asset flows fail to qualify as robust, including or not geographical distance in the group of Z-variables. Although most β_m coefficients are positive and significant at the 5% or 1% level, the coefficients of the equations including the national competitiveness indicator or real interest rate differentials as control variables, are insignificant. Turning to the sub-samples, asset flows do not qualify as robust in either case but are more significant in the second period. From 1997 to 2004, bilateral asset flows are close to becoming a "robust" determinant of business cycle correlation, whereas from 1980 to 1996 none of the coefficients are significant and most of them have the wrong sign. The series representing bilateral flows of bank liabilities broadly follow the series of the asset flows and are not explicitly reported; they never appeared as robust.

3.2.3 Results for policy indicators

Real short-term interest rates

The relation between real short-term interest rates differentials (IRSCDIFF) and business cycle correlation is illustrated in figure 3.19. The regression line is negatively sloped which indicates more highly correlated cycles in the presence of more similar monetary policy. The coefficient is significant at the 10% level but the R^2 (0.03) is far too small for the bivariate regression to be meaningful at all.

EBA results In the full sample, real short-term interest rate differentials do not appear as robust. When negative as expected, the β_m coefficients are far from the significance level and the R^2s of the equations are close to zero. When interest rate differentials turn out as significant, they have a positive sign. The same characteristics apply to the 1980-96 period as for the full sample.

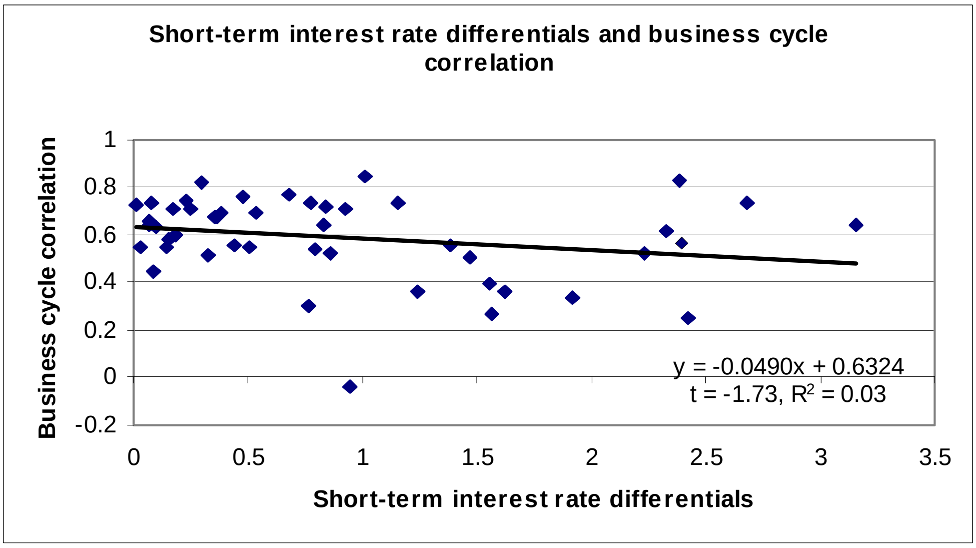


Figure 3.19

More interesting is the fact that real interest rate differentials clearly appear robust when used as a variable of interest in the second period from 1997 to 2004. The result is also robust to the choice of the pool of Z-variables. The coefficients are very significant at the 1% level and the R^2 very large, ranging from 0.6 to 0.7 in the multivariate regressions. The actual coefficients vary between -0.3 and -0.6, which corresponds to extreme bounds of -0.2 and -0.8.[33] Since the preparation for and the implementation of monetary union, business cycle synchronisation and real interest-rate differentials have become more closely related. This result indicates that monetary policy shocks may act as a source of business cycles in themselves. Increasingly coordinated monetary policy could therefore lead to more closely correlated cycles.

Nominal exchange rate variations

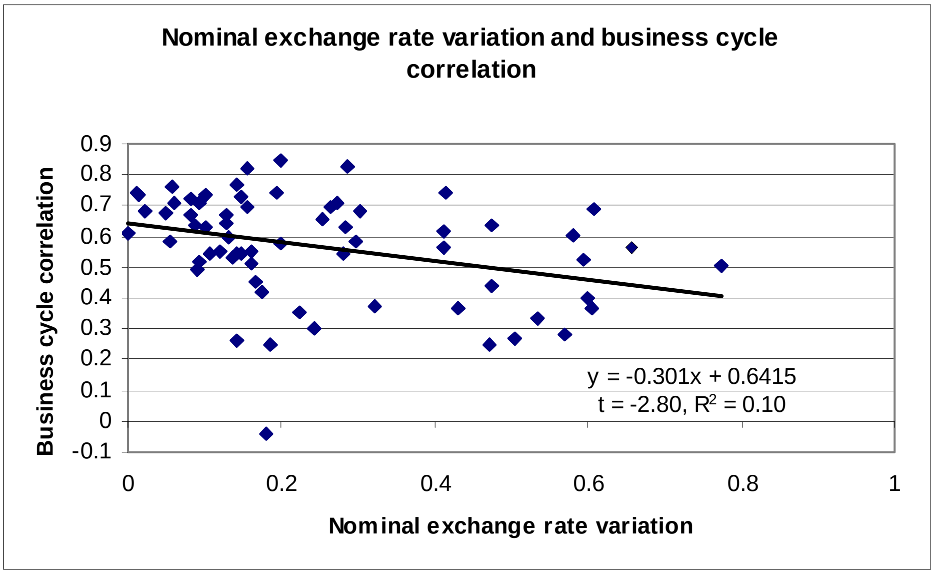


Figure 3.20

[33] The pool of Z-variables include: BTT, TOTMKDIFF, NCIDIF, DEFDIFF, TUDIFF AND GEO-DIST.

We now turn to the relation of nominal exchange rate fluctuations (SD_NERE) and the correlation of business cycles across the euro area. Figure 3.20 suggests a clearly negative relationship according to which a lower standard deviation in the bilateral nominal exchange rates is associated with a higher degree in business cycle comovement. The t-statistic of -2.80 indicates statistical significance and the R^2 of 0.10 is in the medium range when compared to the other bivariate regressions.

EBA results In the full sample and over the 1980-96 period, nominal exchange rate fluctuations do not qualify as a robust determinant of business cycle synchronisation.[34] Nearly all β_m coefficients are negative but many are not significant. It seems that nominal exchange rate stabilisation alone is not sufficient for the synchronisation of business cycles. According to Rose (2000), it takes irrevocably fixed exchange rate in the form of currency union to achieve that goal.

Fiscal deficits

The effects of similar fiscal policies are estimated by the bilateral differentials in fiscal budget deficits as shares of GDP (DEFDIFF). More similar fiscal policies correspond to increased correlation between business cycles as implied by the negative slope of the regression line as presented in figure 3.21. With a t-statistic of -5.2 and an R^2 of 0.2, the relation proves significant. In the case of fiscal deficits, however, we may face a particularly strong case of reverse causation: not only may similar fiscal policies lead to more synchronous cycles but common positions in the business cycle are likely to induce similar fiscal policy responses as well.

[34] In the case of exchange rates, the full sample comprises 1980-1998. The pool of Z-variables include: BTT, TOTMKDIFF, NCIDIFF, DEFDIFF, IRSCDIFF, TUDIFF.

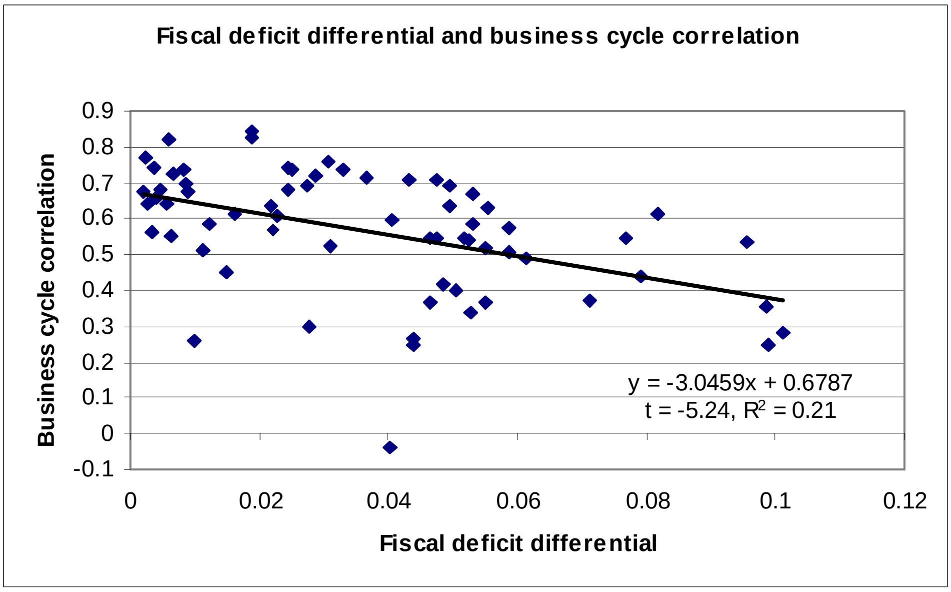


Figure 3.21

EBA results Over the full sample, the fiscal policy indicator appears robustly related to business cycle synchronisation, with extreme bounds ranging from -0.8 to -4.2.[35] All the t-statistics are significant at the 1% level. Over the 1980-1996 period, the case for the fiscal policy indicator comes very close to qualify as robust. All the β_m coefficients are negative and significant at or close to the 5% level but the upper bound becomes positive. The upper bound becomes positive by a small margin. However, a close investigation of the residuals showed that the Germany-Finland pair acted as an outlier in the equation corresponding to the upper bound.[36] This outlier can be easily explained by the shock created by the collapse of the Soviet system in Europe. In Western Europe, Germany and Finland were the countries most affected by that event but the shock had a diverging impact on the two economies. Over the 1980-1996 period, the dummy for

[35] The pool of Z-variables include: BTT, TOTMKDIFF, IRSCDIFF, NCIDIFF, SD_NERE, TUDIFF AND GEODIST.

[36] The residual for Germany-Finland was 3.9 times the standard deviation of the residuals of the equation.

Germany-Finland is significant in all the equations. In addition, the extreme bounds of the fiscal deficit indicator keep the right sign, remaining clearly negative.

As expected, given the timing of the external shock, the Germany-Finland dummy has no significant impact on the results for the full sample and for the second sample. Over the 1997-2004 period, the fiscal policy indicator fails to qualify as robust, with or without dummy for the Germany-Finland pair. Nevertheless, more than 95% of the coefficients remain significant with the right expected negative sign.

The apparent weakening in the power of fiscal deficit differentials to explain business cycle differentials might be related to the Stability and Growth Pact. Since the implementation of the Pact, fiscal policy has become less pro-actively used as a policy instrument to fine tune economic growth. Compared with the 1980-96 period, fiscal deficits may have become more determined by the business cycle and have become less a causing variable of the business cycle.

Table 3.1: Test results for business cycle correlation as a robust determinant of fiscal deficit differentials (1997-2004)

Result	**Estim.**	**Bound**	**Coef.**	**Stdd err.**	**T-Stat.**	**R^2 adj.**	**Z control variables**	No of Z-var.	Out-liers
Fragile	Bivariate		-0.017	0.004	-4.56	0.12			
	High	**0.004**	-0.008	0.006	-1.36	0.31	BTT, IRSCDIFF, TUDDIFF	1,2 and 3	
	Low	**-0.046**	-0.029	0.009	-3.33	0.12	TOTMKDIFF, IRSCDIFF, NCIDIFF		
	High	**0.004**	-0.008	0.006	-1.36	0.31	BTT, IRSCDIFF, TUDDIFF	3	5%
	Low	**-0.046**	-0.029	0.009	-3.33	0.12	TOTMKDIFF, IRSCDIFF, NCIDIFF		
	High	**-0.002**	-0.011	0.004	-2.52	0.26	BTT, NCIDIFF	2	0%
	Low	**-0.043**	-0.029	0.007	-3.89	0.14	IRSCDIFF, NCIDIFF		
	High	**-0.002**	-0.011	0.004	-2.50	0.26	BTT	1	0%
	Low	**-0.031**	-0.019	0.006	-3.03	0.11	IRSCDIFF		

In order to test that hypothesis, we conducted tests on the robustness of business cycle correlation as a determinant of fiscal deficit differentials over the 1997-2004 period. Although robustness was rejected, it was so by a very small margin, suggesting that reverse causation from business cycle correlation to fiscal deficit differential became stronger in the 1997-2004 period.

3.2.4 Results for the structural indicators

Competitiveness

Bilateral differences in competitiveness (NCIDIFF) are plotted against cycle correlation in figure 3.22. As hypothesised, the relationship is clearly negative: the lower the differences in national competitiveness, the larger is the degree of cycle correlation. The more similar countries are in terms of relative price competitiveness, the more comparable will be their ability to adjust to international shocks. With a t-statistic of -4.8, the relation is highly significant. In addition, the R^2 of 0.3 is the highest of all bivariate regressions in this section.

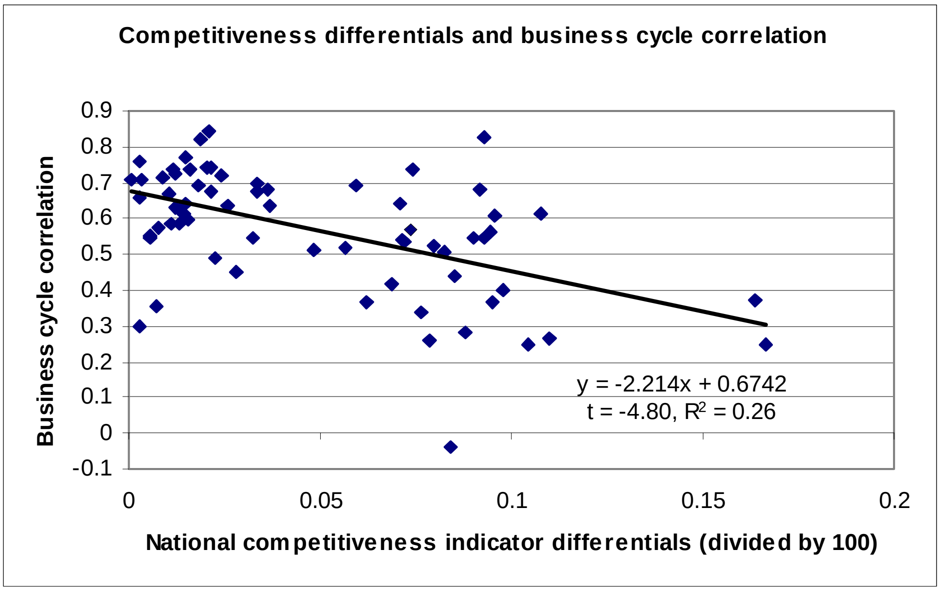


Figure 3.22

EBA results In the multi-regression estimates, excluding geographical distance, national price competitiveness differentials comes out as significant. All coefficients are negative and significant with the extreme bounds ranging from -0.03 to -4.8. When geographical distance was included, NCIDIFF failed to qualify as robust by a small margin.

Nevertheless, all the β_m coefficients were significant and negative. The upper extreme bound coefficient turned slightly positive but remained close to zero when the control Z-variables included geographical distance.

In the sub-samples, including or not geographical distance, the competitiveness indicator clearly fails to qualify as robust. In the first sample from 1980 to 1996, the reason why competitiveness differentials fail to qualify as robust is unclear. Including or not exchange rate volatility in the set of control Z-variables does not affect sensibly the results. Furthermore, although the upper bound becomes strongly positive when bilateral trade or the fiscal deficit differentials are included in the equation, none of these two variables is strongly correlated with the competitiveness indicator which would indicate some multicollinearity. The reason why NCIDIFF does not qualify as robust in the first sub-sample may be plainly due to its weak own explanatory power as indicated by the fairly low t-statistics in the bivariate regression. In the second sample, competitiveness differentials are not even significant in the bivariate regression.[37]

Stock market indices

Figures 3.23 and 3.24 present cross-country differences between the total market indices (TOTMKDIFF) and the cyclical service indices (CYSERDIFF), each plotted against the correlation of business cycles. The two plots display negatively sloped regression lines: the difference between stock markets performances is negatively related to business cycle synchronisation. However, only the cyclical service indicator appears to be significantly correlated to business cycle correlation, with an R^2 of 0.2 and a coefficient significant at the 1% level. The total market indicator does not have a significant coefficient and the

[37] Since the launch of the single currency, differences in national competitiveness are driven essentially by trade-weighted inflation differentials with other euro area countries. Real short-term interest rate differentials also capture essentially changes in national inflation but on a bilateral basis. Over the 1997-2004 period, the two series tend to reflect more the same shocks than in the previous samples, due to the fixed exchange rates. Nevertheless, tests conducted by replacing real short-term interest rate differentials with nominal short-term interest rate differentials in the group of control Z-variables, also led to the rejection of robustness for NCIDIFF over the 1997-2004 sample.

R^2 is too small to be meaningful.

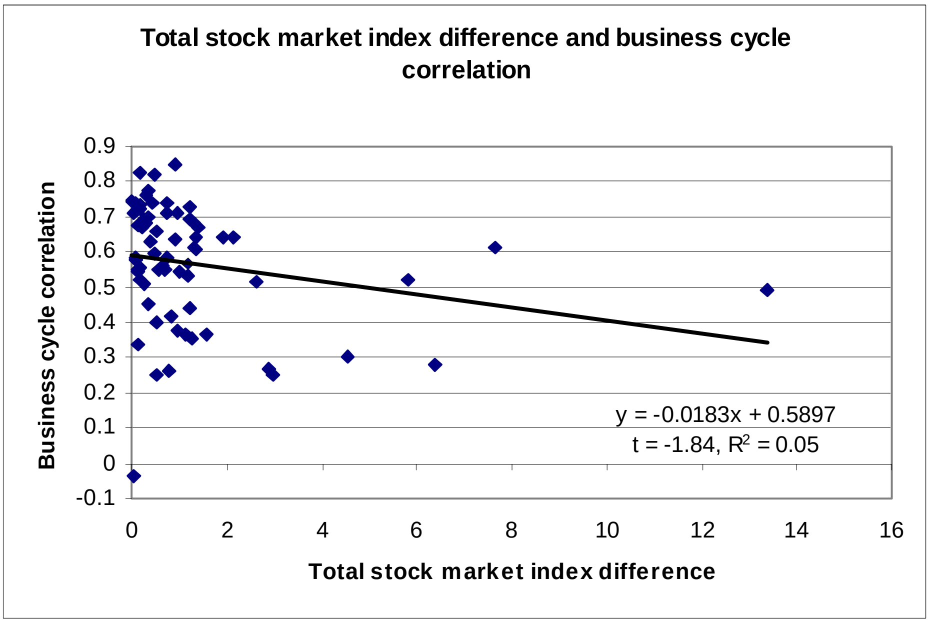


Figure 3.23

EBA results Although the difference between total stock market indices (TOTMKDIFF) did not appear significant on a bilateral basis over the full sample, we tested it in multivariate regressions (Table B. 10a). Overall stock market performance is indeed a key financial indicator and may have turned robust in the sub-samples. Although over the 1980-96 period, TOTMKDIFF is significant at the 1% level in the bivariate regression, it fails to qualify as robust for that period, as well as in the second sample.[38]

By contrast, the relative stock market performance in the sector of cyclical services (CYSERDIFF) is clearly significant over the 1980-04 and 1997-04 periods. Over the full sample, CYSERDIFF comes clearly out as robustly related to business cycle correlation Table B. 10b). All the β_m coefficients are significant at the 1% level. The extreme

[38] When substituting economic specialisation for bilateral trade in the standard pool of explanatory variables, overall stock market differentials came out as robust in the 1980-1996 sample but the R^2s were all very small at less than 0.1 in most equations.

bounds range from -0.001 to -0.012, with R^2s of 0.4 and 0.2, respectively. However, differences between national total stock market indices does not appear related at all to business cycle correlation, either in the full sample or in the sub-samples.

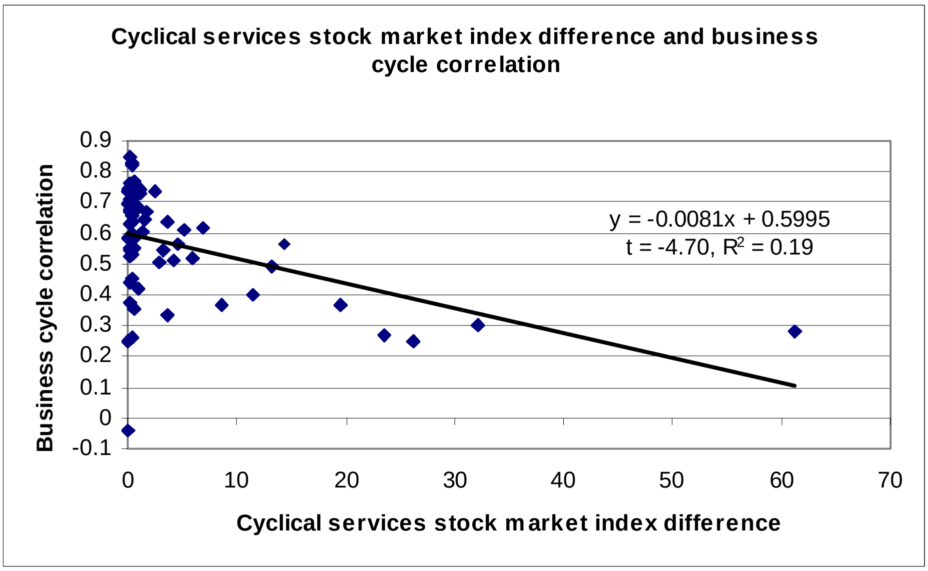


Figure 3.24

In the first sample period from 1980 to 1996, the cyclical service indicator does not qualify as robust but in the second sample from 1997 to 2004, it clearly appears robust with all β_m coefficients significant at the 5% level. Although the upper bound is very small, the R^2 is very high at 0.8. In the last sample, the standard errors of the β_m coefficients are noticeably larger than in the full sample and than in the first period, probably due to the overall increase in stock market volatility.

Labour market flexibility

In theory, more flexible labour markets should help an economy to adjust to asymmetric shocks and hence lead to more synchronous cycles even in the presence of idiosyncratic shocks. However, labour market flexibility is difficult to measure. We apply two alternative indicators, trade union density and an employment protection index and use the bilateral differences (TUDDIFF and EPADIFF, respectively) to measure the degree of similarity across countries. High values indicate very different flexibility regimes whereas low values suggest rather similar labour market conditions. Both indices are plotted against cycle correlation as shown in figures 3.25 and 3.26. Although the coefficients exhibit the expected negative sign, neither of them is statistically significant. The trade union density differential's t-statistic is -0.7, the corresponding value for the employment protection index differential is -0.7. The R^2s are around zero.

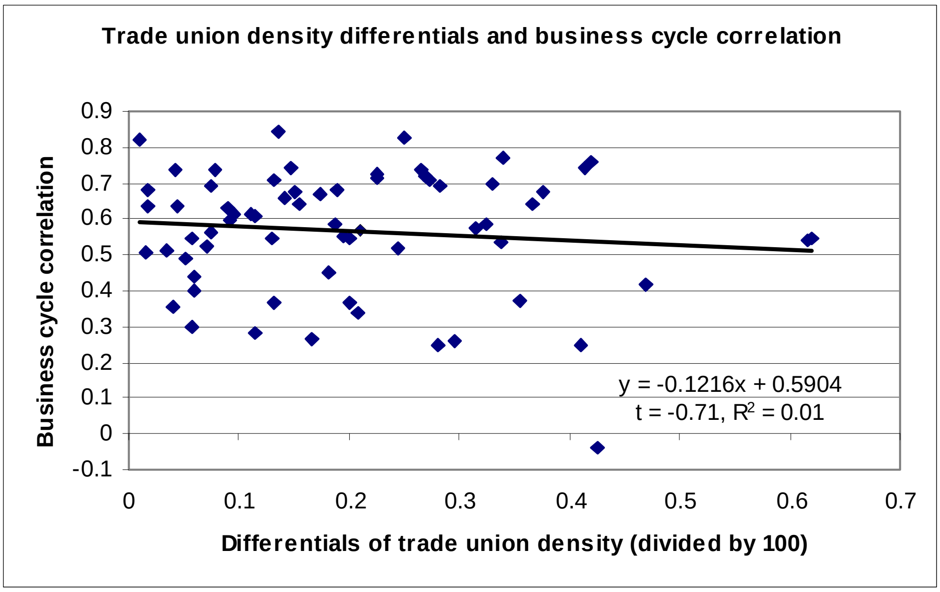


Figure 3.25

EBA results In the multivariate regressions we focus on the trade union density differential due to limited data for the EPA indicator which is available for only three years available from 1990 to 2003. In none of the estimates and sub-samples, the trade union differential qualifies as robust.

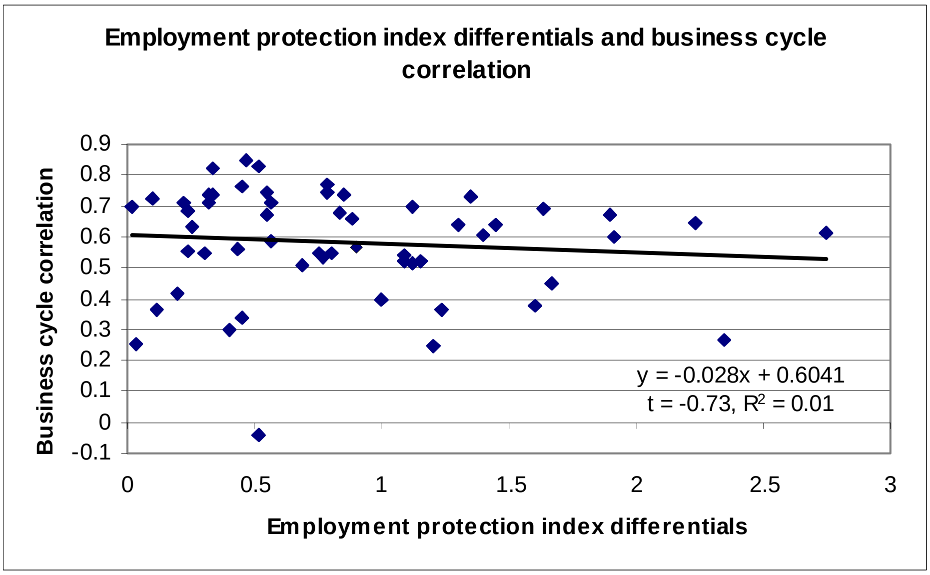


Figure 3.26

Gravity variables

Gravity variables have been used extensively in the empirical trade literature to account for exogenous factors. Traditionally, geographical distance and relative size are the core gravity measures. Figures 3.27 and 3.28 provide the corresponding scatter plots, relating the gravity variables to business cycle correlation. In the case of geographical distance, the case is surprisingly clear. The closer countries are located next to each other, the more synchronous are their business cycles. With a t-statistic of -5.2 and an R^2 of 0.3, the relation exhibits strong significance and a fair goodness of fit. We would not have expected such a clear result, given the relatively small distances and low transport costs in Europe.

The second gravity variable, relative population size, is plotted against cycle correlation in figure 3.28. We would expect a negatively sloped regression line, hypothesising that countries of similar size may have more synchronised business cycles. The scatterplot falsifies this hypothesis. Although the line slope is slightly negative, it is not significant; the t-statistic is only -0.4. Neither is the goodness of fit satisfactory, with an R^2 around zero.

We did not test for the robustness of the relative population size, because coefficients on that variable not only failed to be significant in the bilateral and in the multilateral regressions, but were also of the wrong expected sign.

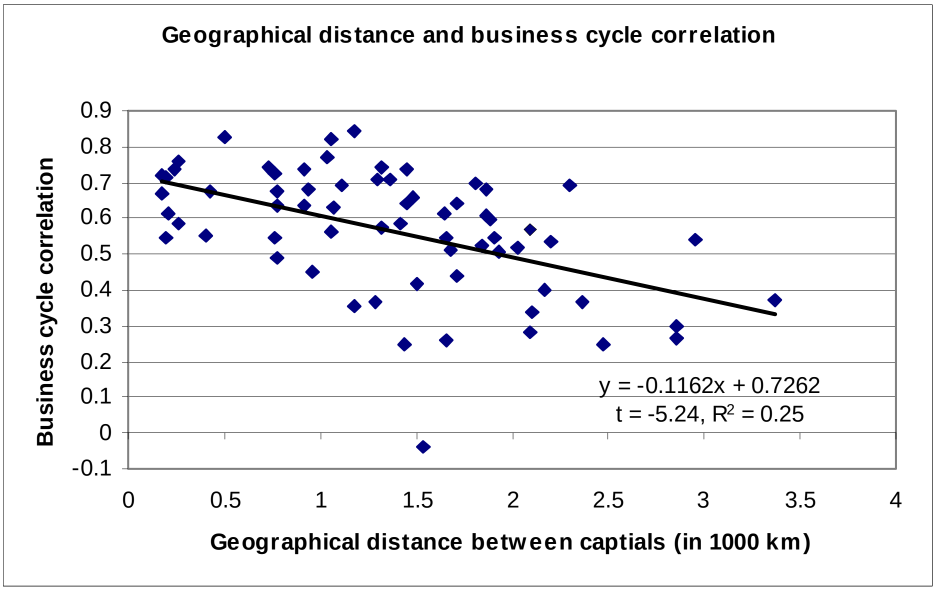


Figure 3.27

EBA results Surprisingly, geographical distance appears robust in the period from 1997 to 2004 but not in the previous period and not in the full sample.[39] The difference of result between the different samples may have reflected a partial correlation problem between geographical distance and the ratio of bilateral trade to total trade (BTT). Indeed, the pool of Z-variables we drew from to test the robustness of geographical distance also includes the ratio of bilateral trade to total trade which emerged as a robust

[39] The pool of Z-variables include: BTT, TOTMKDIFF, NCIDIFF, DEFDIFF, IRSCDIFF, SD_NERE AND TUDIFF.

determinant of business cycle correlation in the full sample and in the first sub-sample but not in the second one. Bilateral trade is also strongly related to geographical distance. However, tests conducted by replacing bilateral trade with economic specialisation in the pool of Z-variables, did not support that assumption. Although economic specialisation is not at all correlated to geographical distance, the latter came out again as nearly robust in the last sample,[40] whereas for the 1980-04 and 1980-96 periods the rejection of robustness was clear-cut.

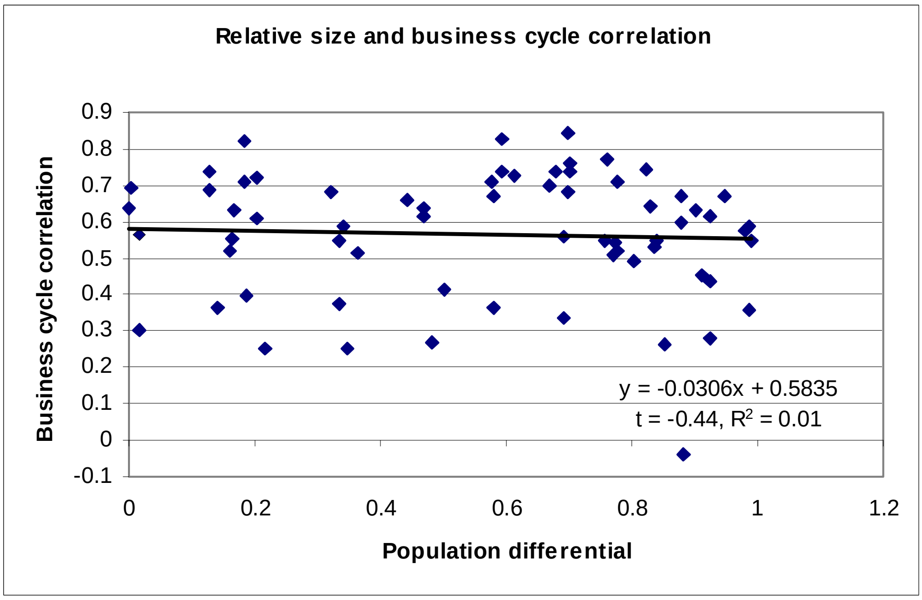


Figure 3.28

3.3 Conclusion

This chapter dealt with the determinants of business cycle synchronsiation among euro area countries. In the context of the endogeneity hypothesis of optimum currency areas, we investigated whether business cycles are likely to become more or less synchronised

[40] The coefficients are all negative and significant at the 5% level but the upper bound is around zero.

under the influcen of EMU. Since it is still too early to isolate a direct "euro effect" reliably, we followed Frankel and Rose (1998) in their approach to estimate the effect of trade on business cycle synchronisation. In theory, it is unclear whether increased trade leads to more synchronised cycles or, as Krugman (1993) suggests, to more specialisation and hence less cycle synchronisation. In addition to trade, we tested a large number of other potential determinants and apply the extreme-bounds analysis (EBA) by Leamer (1983). We split our 25-year period in sub-samples to learn more about time-variant effects.

The main results of the EBA analysis are presented in Table 3.2. The table shows the variables that qualify as "robust" in the strict sense and those for which robustness is rejected by a very small margin ("quasi-robust"); cases when more than 95% of coefficients are significant but robustness is rejected are also reported.

We need to take into account that, as emphasised by Levine and Renelt (1992), the EBA is not a causality analysis. For that reason, the choice of variables as potential determinants of business cycle synchronisation relies on economic theory. The upper panel presents the variables which were selected as potential determinants of business cycle synchronisation, the so-called "M-variables of interest". For these variables, economic literature indicates that they should influence business cycle synchronisation. The lower panel presents variables which were used as "control Z-variables". Economic theory tells us that several of these variables should have something to do with economic growth and with the business cycle. However the direction of the causality is far less clear than in the case of the M-variables. This is particularly obvious in the case of fiscal deficits and of the exchange rate where the relation works both ways, especially in the short run. This does not mean that the Z-variables are not determinant of the business cycle but indicates that the relationship is more likely to be two-way than in the case of the M-variables.

Variable[1]	**1980-2004**	**1980-1996**	**1997-2004**
M-variables: traditional determinants of business cycle synchronisation			
Ratio of bilateral trade to total trade (BTT)	**Robust**	**Robust**	Fragile
Ratio of bilateral trade to GDP (BTY)	**Robust**	**Robust**	Fragile
Trade specialisation (TRADEPAT)	Fragile (significant)	Fragile	**Robust**
Fuels	*Fragile*	*Fragile*	*Fragile*
Machinery and transport equipment	*Fragile (significant)*	*Fragile*	***Robust***
Other manufacturing	*Fragile*	*Fragile*	*Fragile*
Chemicals	*Fragile*	*Fragile*	*Fragile*
Economic specialisation (ECOPAT)	Fragile	Fragile	Fragile
Industry	**Robust**	**Quasi-robust** (significant)	*Fragile*
Construction	*Fragile*	***Robust[2]***	*Fragile*
Wholesale and retail trade	*Fragile*	*Fragile*	*Fragile*
Financial intermediation	*Fragile*	**Quasi-robust** (significant)	*Fragile*
Bilateral flows of bank assets (LBFA)	Fragile	Fragile	Fragile
Z-variables: policy and structural indicators			
Real short-term interest rate differential (IRSCDIFF)	Fragile	Fragile	**Robust**
Nominal exchange rate volatility (SD_NERE)	Fragile	Fragile	--
Fiscal deficit differential (DEFDIFF)	**Robust**	**Robust[3]**	Fragile (significant)
Price competitiveness differential (NCIDIFF)	**Robust**	Fragile	Fragile
Stock market differential, cyclical services (CYSERDIFF)	**Robust**	Fragile	**Robust**
Trade union membership differential (TUDDIFF)	Fragile	Fragile	Fragile
Geographical distance (GEODIST)	Fragile	Fragile	**Robust**

1. As they failed to be significant in the bivariate baseline regression, we do not report the EBA results for the following variables: Trade openness (TTY), log-bilateral bank liability flows (LBFL), employment protection differential (EPADIFF), and relative population (POPDIFF).
2. Qualifies as robust but the coefficient has the wrong (positive) expected sign.
3. Including a dummy for the Germany-Finland country pair.

Table 3.2: Summary of EBA results

In the full sample, among the potential determinants of the business cycle, the ratios of bilateral trade to total trade and to GDP as well as the fiscal deficit differentials, the stock market differentials for cyclical services and the differentials in national competitiveness come out as robust. While overall economic specialisation does not qualify as a robust determinant of business cycle synchronisation, differences between the shares of industrial sectors in total value-added meet the criteria. Similarities in overall trade specialisation and in the relative specialisation particular in machine and equipment have a significant coefficient in all equations but do not qualify as a robust determinant in the strict sense because of the relatively large standard errors on the estimated coefficients.

When considering the results for the sub-periods, the variables robustly related to business cycle synchronisation from 1980 to 1996 are the ratios of bilateral trade and the fiscal deficit differentials. The relative shares of the industrial and financial sectors and the fiscal deficit differentials do not fully qualify for robustness but are very close to it. Over the period from 1997 to 2004, trade specialisation in particular in machinery and transport equipment, the real short-term interest rate differentials and the stock market differentials for cyclical services all appear robustly related to business cycle synchronisation.

The EBA results confirm external trade as a key determinant of business cycle synchronisation in the context of the euro area. Given the theoretically unclear case of the trade effect on cycle correlation, our results support the OCA endogeneity view of Frankel and Rose (1998). They find a strongly positive effect for a wide array of countries and on these grounds postulate the "endogeneity of the optimum currency area criteria": if trade promotes the comovement of business cycles, then a common currency that fosters trade would endogenously lead to more synchronised cycles in the monetary union. Also in keeping with the results of Rose (2000) and its "Rose effect"[41] we fail to identify a direct "robust" relation between exchange rate volatility and business cycle correlation.

The effect of monetary union is closely related to our second major finding on the impact of trade specialisation and the degree of intra-industry trade. The positive trade effect on cycle correlation hinges on the degree of intra-industry trade, i.e. the similarity of trade specialisation patterns. The more intra-industry trade, the more likely is

[41] "Entering a currency union delivers an effect that is over an order of magnitude larger than the impact of reducing exchange rate volatility from one standard deviation to zero", Rose (2000: 17).

the positive trade effect to materialise. Empirical evidence indicates an increased degree of intra-industry trade over time across euro area countries, even though the very broad economic structures do not seem to have not converged. The EBA analysis shows that similar trade specialisation emerges as a robust determinant of cycle correlation in the 1997-2004 period. Taken together, these findings support the Frankel and Rose (1998) prediction that EMU would lead to trade expansion and to the development of intra-industry trade, rather than to greater trade specialisation, which in turn would result in more highly correlated business cycles. The transmission of industry-shocks via intra-trade seems to be concentrated in the sector of machinery and equipment: trade specialisation in machinery and equipment alone explains 61% of cycle correlation in 1997-2004.

The positive impact of stock market comovements in the cyclical service sector on cycle correlation can be interpreted either as an indication that financial integration has been conducive of greater cycle symmetry or that cyclical services themselves have become a channel of transmission of business cycle fluctuations across countries. The second hypothesis of a direct link seems more appropriate since the relative performance of overall stock market indices does not appear clearly as a major determinant of business correlation.

Further research would be required on financial integration. Although the bivariate correlation between bank flows and cycle synchronisation is quite strong the EBA results remain weak, partly due to incomplete data sets. Another area of research is competitiveness differentials which would require more in-depth investigation of the interactions with the synchronisation of business cycles.

We acknowledge that more time is needed to make definite statements on the effects of the euro on business cycle synchronisation. As of today, however, we believe that our results indicate a cautiously optimistic view on *ex post* optimality of the euro area.

Chapter 4

Risk sharing, financial integration and Mundell II in the enlarged European Union

This chapter deals with the the latest advancement in OCA theory, known as Mundell II. In contrast to the initial OCA literature (Mundell I) and in interaction with the endogeneity hypothesis of optimum currency areas, Mundell II draws attention to risk sharing and the role of financial markets in a currency union. In the presence of financial market integration, it is argued, those countries with little business cycle synchronisation may benefit even more from adopting a common currency. This benefit arises from new consumption risk sharing opportunities because, in a financially integrated currency union without exchange rate fluctuations and risk premia, national consumption patterns should be diversified across the union and less contingent on home income. Even if the degree of financial integration is limited in the first place, the common financial market created by the currency union would unfold beneficial risk-sharing effects and make the adoption of the common currency more and not less attractive if cross-country business cycles lack synchronisation.

We investigate some measures of risk sharing and financial integration for the enlarged European Union. Given that the degree of business cycle comovement of the new member states (NMS) with the euro area is still limited, it is interesting to learn more about the past degree and future potential of risk sharing in the context of euro adoption. The prevailing policy question is whether it makes sense for the NMS to wait until their cycles are sufficiently synchronised with the euro area or whether they should join early and benefit from the euro area's risk-sharing property, even and especially in the presence of non-synchronised cycles.

In the following, we examine the eight Central and Eastern European NMS[1] in relation to the euro area as an aggregate. For comparison, we apply similar tests to the "old" EU members. We use correlation and codependence measures of cross-country consumption and output comovement to proxy the degree of risk sharing before we analyse financial market integration by employing a number of interest rate comovement indicators. We find that the degree of risk sharing between the new member states and the euro area is limited, hence the potential gain from euro adoption may be substantial. Furthermore, we note that both consumption and output comovement have been increasing over time, for the NMS as well as for the EU-15. The reasons for little risk sharing may be attributed to a relatively low degree of financial integration between the NMS and the euro area which is revealed by the analysis of real interest rate comovement. The introduction of the euro may, however, unfold endogenous effects particularly on financial markets which may change the picture. Results from the EU-15 countries indicate a rising degree of financial integration during the preparation for monetary union.

4.1 Risk sharing

This section portrays the conceptual framework of the risk-sharing analysis and presents the empirical results of consumption and output comovement in the enlarged EU. We investigate the degree of risk sharing between the NMS and the euro area during the last decade and compare their experience to that of the "old" EU countries. The the-

[1] The Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia and Slovenia.

oretical foundation of analysing consumption correlations in the context of risk sharing is based on models of markets for contingent claims. In a world of complete markets, consumers can diversify risk by investing in Arrow-Debreu securities. These financial assets constitute contingent claims and deliver a state-contingent pay-off. By purchasing and selling Arrow-Debreu securities, households can consume the same amount of resources in varying states of the world. In other words, they can effectively insure against domestic risks and decouple their consumption patterns from domestic income flows. In equilibrium, cross-country consumption should be highly correlated because national consumption is internationally diversified and thus invariant to domestic output shocks. On these grounds, Backus, Kehoe and Kydland (1992) construct a calibrated international real business cycle model which predicts consumption to be more highly correlated than output across countries.

Empirical analysis, however, has not substantiated this prediction. In fact, cross-country consumption tends to be less highly correlated than output. The resulting consumption correlation puzzle is one of the "six major puzzles in international macroeconomics" as pointed out by Obstfeld and Rogoff (2000). Various reasons may be responsible for this puzzle. Low degrees of financial integration may prevent consumers from diversifying their portfolios internationally to the Arrow-Debreu degree. Also, trade costs and other barriers to international trade may inhibit risk sharing across countries. Moreover, a large degree of non-traded goods may contribute to the puzzle since risk sharing is possible only for risk to tradable output. Hence, measuring cross-country correlation in consumption of tradables only may alleviate the puzzle. Another measurement issue pertains to the output side. Given that only output remaining after investment and government consumption can be shared by private consumers, consumption correlations should rather be compared to correlations in GDP net of investment and government consumption, see Obstfeld and Rogoff (2000). In practice, however, limited data availability often restricts this type of analysis. Furthermore, the model proposed by Stockman and Tesar (1995) emphasises the role of influences from the demand side, particularly taste shocks, which may be responsible for low cross-country consumption correlation. Finally, Imbs (2006) investigates potential interactions between financial integration, output and consumption correlation. He finds that increased financial integration does not only raise consumption correlations across countries but that it boosts output correlation to an even larger degree. As a result, he argues, "the bulk of the quantity puzzle originates in the tendency for GDP correlations to increase with financial links, not in low risk sharing" (Imbs 2006: 315).

The following empirical investigation of risk sharing in the EU proceeds in two steps. We explore consumption and GDP comovement first by looking at cross-country correlations and then move on to the codependence analysis.

4.1.1 Consumption correlation

In a first step, we compare cross-country correlations of consumption and GDP. We use quarterly data of real private consumption and real GDP for the euro area and the eight NMS over the time period 1995Q1-2005Q4.[2] For comparison, we also investigate 14 "old" EU countries.[3] Data mostly stem from Eurostat, supplemented by national sources. Given that we are interested in the new member countries' prospective adoption of the euro, we correlate each country with the aggregate euro area.

Table 4.1 presents consumption correlation coefficients of growth rates and various cycle specifications. We derive the latter by detrending real GDP applying the Hodrick-Prescott (HP) filter and the Baxter-King (BK) band-pass filter.[4]

[2] Private consumption includes consumption of households and non-profit institutions serving households (NPISH). All data are in euro, scaled to 1995 prices and exchange rates, indexed and taken in logs. At this stage, we use seasonally-adjusted data. In the following section, we apply the codependence framework which incorporates seasonal adjustment within the statistical model and hence employs non-adjusted data.

[3] These include the EU-15 without Luxembourg, due to data constraints.

[4] See Hodrick and Prescott (1997) and Baxter and King (1999).

Table 4.1: Consumption correlation

Country	Growth rates	HP cycles	BK cycles ($k = 4$)	BK cycles ($k = 8$)
Czech Rep.	-0.06	-0.51	-0.20	-0.21
Estonia	-0.06	-0.62	-0.24	-0.52
Hungary	-0.16	0.23	0.07	0.06
Latvia	-0.31	-0.44	-0.32	-0.33
Lithuania	-0.28	-0.34	-0.04	0.03
Poland	0.20	-0.28	-0.23	-0.26
Slovakia	-0.02	-0.39	-0.26	-0.37
Slovenia	0.15	0.12	0.05	0.22
Austria	0.34	0.76	0.50	0.60
Belgium	0.17	0.72	0.54	0.65
Denmark	0.26	-0.26	0.00	-0.30
Finland	0.38	-0.12	0.07	0.03
France	0.63	0.82	0.59	0.79
Germany	0.76	0.83	0.60	0.79
Greece	-0.25	-0.20	-0.04	-0.03
Ireland	0.49	0.83	0.42	0.75
Italy	0.41	0.64	0.48	0.61
Netherlands	0.33	0.73	0.49	0.79
Portugal	0.38	0.70	0.33	0.67
Spain	0.31	0.79	0.61	0.80
Sweden	0.45	0.78	0.52	0.68
UK	0.39	0.57	0.21	0.45

Note: Correlation coefficients of real private consumption vis-à-vis the aggregate euro area in growth rates and cycles, applying the Hodrick-Prescott (HP) filter and the Baxter-King (BK) filter, the latter with alternative lead/lag parameters $k = 4$ and $k = 8$.

Both filters have been used extensively in business cycle analysis. The BK filter identifies the cyclical component by removing very high and very low frequency fluctuations from the data but the choice of the lead/lag parameter k involves a trade-off particularly in small samples like ours. The larger k, the more periods need to be cut off at the beginning and at the end of the sample. A smaller k, however, reduces the reliability of the results. The HP filter involves minimising the variance of the cyclical component but has been critised for the arbitrariness of the smoothing parameter employed. Although the HP filter does not reduce the sample size like the BP filter, the HP marginal values tend to be biased due to the required estimation of values for differencing.

According to table 4.1, the correlation of consumption with the euro area is very low for all NMS. In fact, the majority of coefficients is even negative, regardless of the specification of the indicator. Estonia, Latvia and Slovakia exhibit the lowest correlation whereas only Slovenia's consumption is positively correlated with euro area consumption throughout specifications, with coefficients ranging from 0.05 to 0.22. Not surprisingly, consumption correlation is much higher for EU-14 countries. France and Germany are characterised by top values between 0.59 and 0.83 while this is, of course, partly due to their large weight in euro area aggregate consumption. Depending on the specification, large correlation coefficients also pertain to Ireland, Spain and Sweden. We note that the correlation coefficients of the non-euro area members Sweden and the UK are not considerably lower than than those of euro area countries. Low and partly negative coefficients can be observed, however, in the cases of Denmark, Finland and Greece.

Table 4.2 presents the same growth rate and cycle specifications for GDP correlations. For the NMS, most coefficients take positive values although the sizes vary across specifications. Hungary stands out with the largest correlation coefficients of up to 0.88. Also, Slovenia and, in part, Poland show a relatively large degree of output correlation with the euro area. Lithuania, Slovakia and, partly, the Czech Republic have rather low, if not negative coefficients. For the EU-14 countries, France and Germany again exhibit the largest correlation values, between 0.72 and 0.97. Other countries with large coefficients include Belgium, Italy, the Netherlands and the UK. Greece has again by far the lowest correlation coefficients.

Regarding the consumption correlation puzzle, we turn to the differences between consumption and GDP correlations across countries. Figure 4.1 illustrates this gap at the example of the HP-filtered series.[5]

[5] We acknowledge that both the HP and the BP filters deliver imperfect results in the presence of

Table 4.2: GDP correlation

Country	Growth rates	HP cycles	BK cycles ($k = 4$)	BK cycles ($k = 8$)
Czech Rep.	-0.13	0.05	0.04	0.54
Estonia	0.20	-0.06	0.59	0.33
Hungary	0.43	0.78	0.76	0.88
Latvia	0.06	0.03	0.40	0.27
Lithuania	-0.18	-0.51	-0.04	-0.19
Poland	0.32	0.18	0.68	0.62
Slovakia	0.05	-0.42	0.18	-0.06
Slovenia	0.08	0.41	0.35	0.58
Austria	0.43	0.72	0.69	0.80
Belgium	0.63	0.83	0.83	0.85
Denmark	0.34	0.53	0.39	0.77
Finland	0.17	0.51	0.36	0.61
France	0.72	0.93	0.93	0.94
Germany	0.78	0.92	0.93	0.97
Greece	0.05	0.06	-0.13	0.00
Ireland	0.45	0.67	0.65	0.82
Italy	0.65	0.87	0.89	0.94
Netherlands	0.68	0.79	0.84	0.82
Portugal	0.30	0.64	0.21	0.41
Spain	0.46	0.83	0.61	0.75
Sweden	0.53	0.80	0.70	0.75
UK	0.37	0.71	0.71	0.85

Note: Correlation coefficients of real GDP vis-à-vis the aggregate euro area in growth rates and cycles, applying the Hodrick-Prescott (HP) filter and the Baxter-King (BK) filter, the latter with alternative lead/lag parameters $k = 4$ and $k = 8$.

It is very obvious that the consumption-GDP gap is negative and with down to -0.56 very large for most NMS, i.e. the consumption correlations are considerably lower than the GDP correlations. This is a first indication that the consumption correlation puzzle applies for the NMS. The only two positive gaps in the cases of Slovakia and Lithuania stem from the fact that both consumption and GDP correlations are very negative, with GDP even exceeding consumption correlation in absolute value.

small samples. To avoid further reduction of our sample, we employ the HP filter for the following exercise.

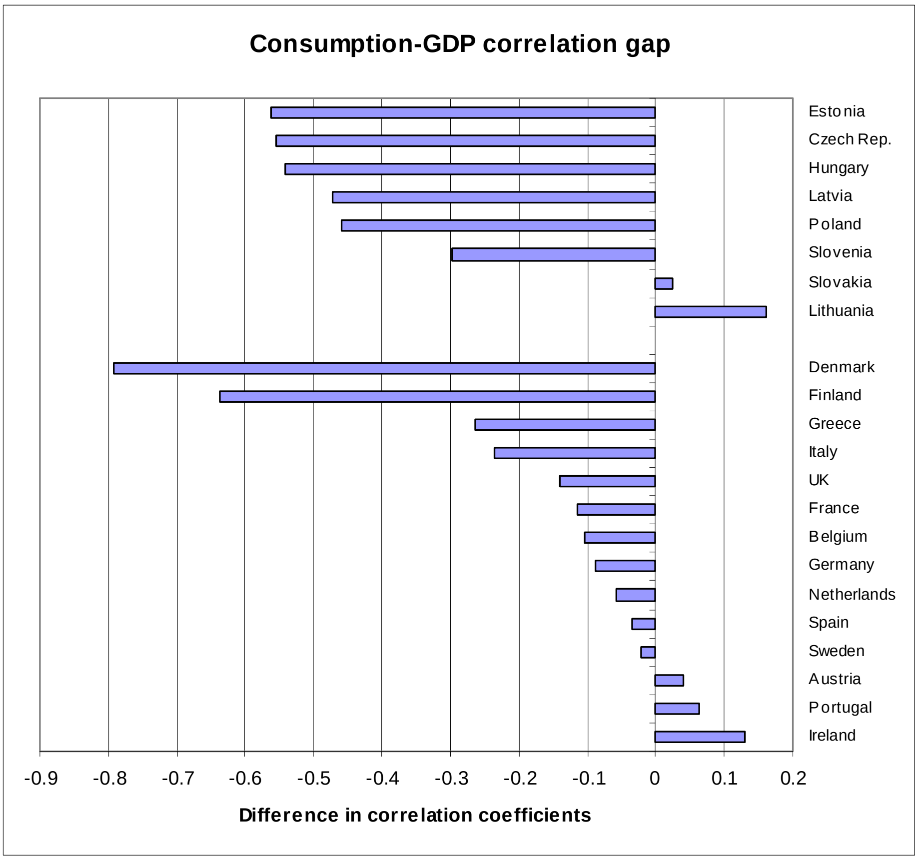


Figure 4.1: Differences in the correlation coefficients of real consumption (households and NPISH) and real GDP vis-à-vis the euro area, 1995-2005 (HP-filtered series).

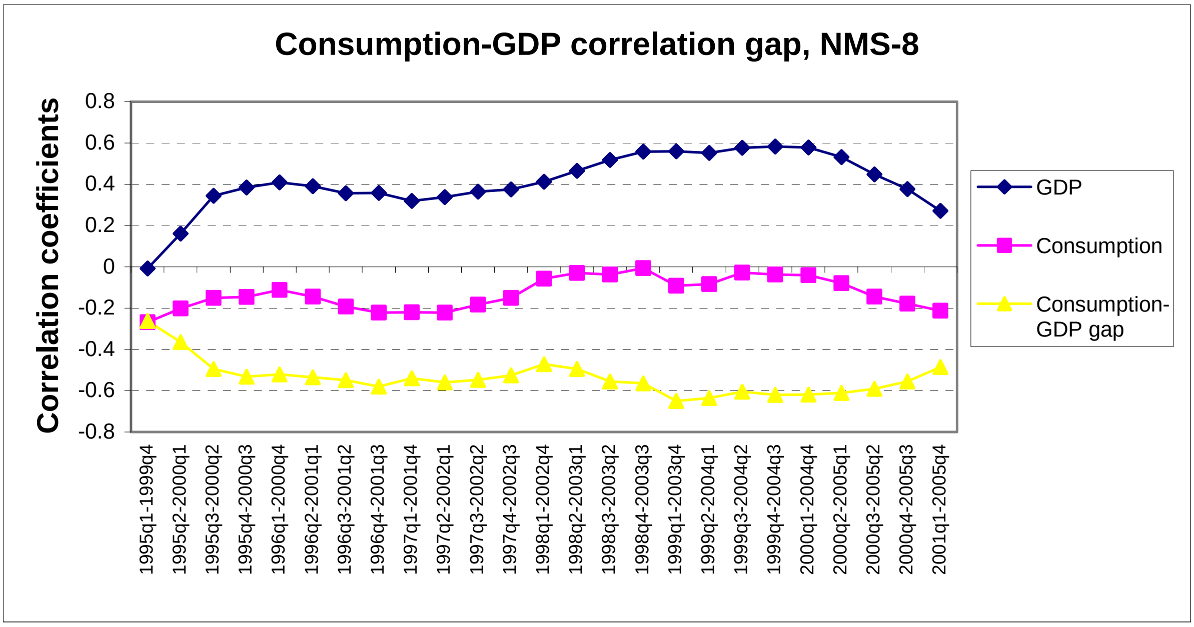
For the EU-14 countries, we identify large negative gaps for Denmark (-0.79) and Finland (-0.69) whereas the remaining countries are characterised by much smaller or even positive gaps. Except for Greece and Italy, all remaining countries have values above -0.20. Austria, Portugal and Ireland have positive gaps, i.e. for these countries, consumption correlation exceeds output correlation - an indication of functioning risk sharing with the euro area. Taken together, the consumption-GDP correlations seem to indicate that those countries which have shared years of economic integration already (EU-14) tend to have much smaller consumption-GDP gaps than those still in economic transition. Hence, the consumption correlation puzzle may decline as integration proceeds.

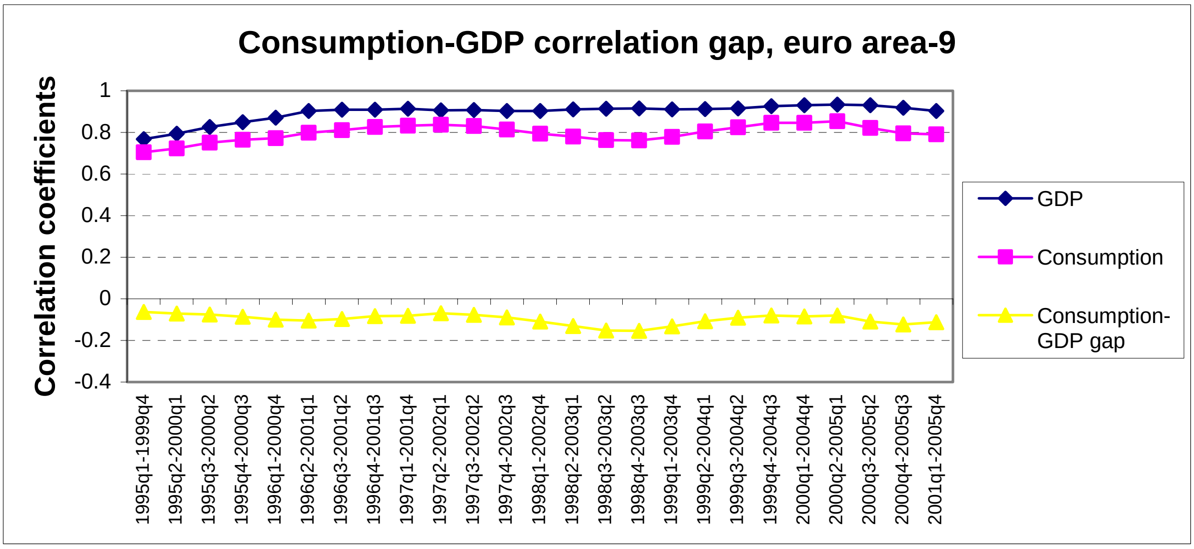
To find out more about the dynamics of risk sharing, we investigate rolling correlation windows. Figures 4.2-4.4 depict 5-year rolling windows ranging from 1995Q1-1999Q4 to 2001Q1-2005Q4. Due to the large number of countries, we form country groups composed of weighted averages of correlation coefficients.[6]

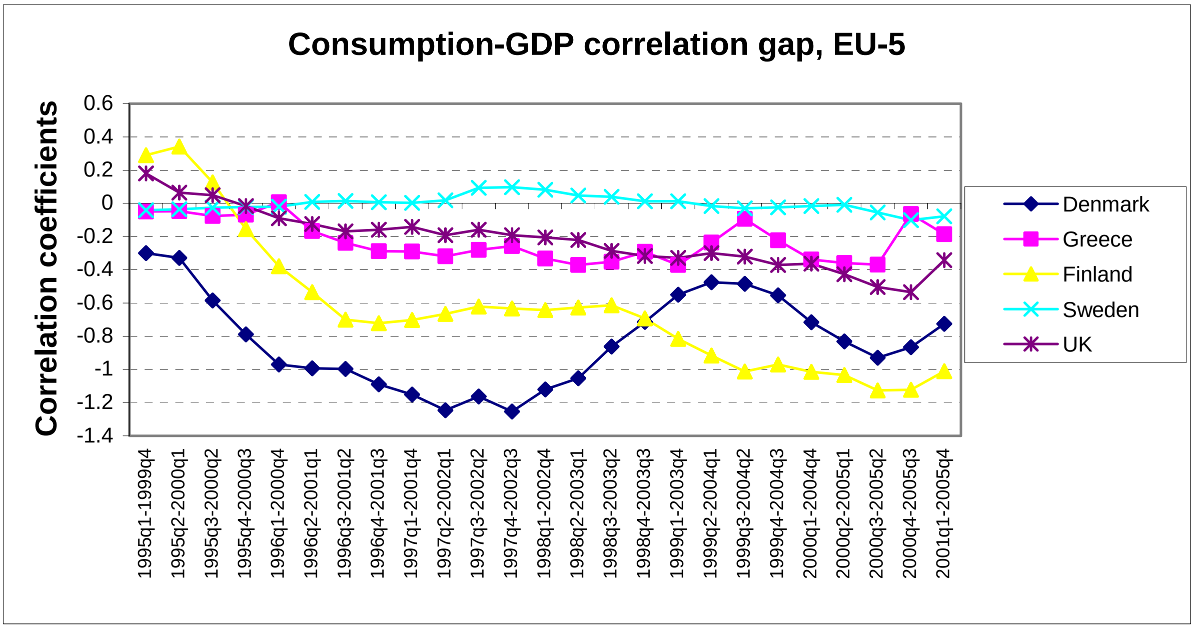
Figure 4.2 includes the eight NMS and shows that the average degree of GDP correlation with the euro area has increased markedly from -0.01 up to 0.58 during the 1999Q4-2004Q3 window before it declined to 0.27 in the most recent period. The very last windows may, however, be subject to some endpoint instability of the detrending filters and hence not be overestimated. The average consumption correlation of the NMS-8 with the euro area is clearly below GDP correlation. It has, however, risen from a starting value of -0.27 to a maximum of 0.00 in 1998Q4-2003Q3 and then moved down to -0.21. The distance between consumption and GDP correlation is illustrated by the bottom line in the graph. On the whole, the gap has widened over time.

Figure 4.3 averages nine euro area countries (EA-9) which seem to behave roughly similar. The euro area countries Finland and Greece, in contrast, appear idiosyncratic and hence grahped together with the non-euro area countries in figure 4.4 (EU-5). Although GDP correlation exceeds consumption correlation for the EA-9 countries, both lines are at far higher levels and have a more narrow gap than the NMS-8. GDP correlation of the EA-9 increased from 0.77 to 0.93 in 2000Q2-2005Q1 before it fell slightly to 0.90. Consumption correlation also rose on average from 0.70 to 0.85 in the same peak window as GDP, then decreasing somewhat to 0.79. As in the case of the NMS-8, we observe increasing rates of both GDP and consumption correlations, though at a lower rate for consumption. This finding is summarised by the negative and decreasing gap

[6] We use relative GDP as weighting factor for averaging the respective correlation coefficients. Applying unweighted averages instead did not have a major impact on the results. The presented figures are based on HP-filtered data.







Figures 4.2-4: 5-year rolling correlation windows of quarterly HP-filtered consumption, GDP and their difference, against the euro area. See the text for exact country coverage.

line. However, the EA-9 gap never touches the -0.20 shreshold.

The experience of the remaining EU-5 countries is less uniform. Figure 4.4 graphs only the consumption-GDP gaps but for each country individually. While the gap lines of Sweden and the UK declined moderately, we observe a massive decline in the case of Finland and a very volatile behaviour for Denmark and Greece.

On the whole, our correlation results confirm the consumption correlation puzzle for the NMS and the EU-14 countries as GDP correlations frequently exceed consumption correlations. However, the correlation levels of the EA-9 countries are much higher than for the NMS. Also, the gaps are more narrow. This may lead us to the conclusion that, as integration between the NMS and the euro area makes progress, the consumption-GDP gap may go down. Another interesting overall observation is that both GDP and consumption correlations increased on average over time. This may, without having performed any causal analysis, be interpreted as supportive evidence of the hypothesis by Imbs (2006). He suggests that the consumption-GDP gap widens not because of lacking risk sharing. Instead, he argues, it is financial integration with promotes both GDP and consumption correlation. According to his estimates, the effect of financial integration on GDP, or business cycle correlation is much stronger than that on consumption correlation. As a result, a widening consumption-GDP gap may be a more ambiguous phenomenon than previously assumed.

4.1.2 Consumption codependence

In addition to the correlation analysis above, we explore the data using the codependence framework. This method is a more sophisticated time-series technique which takes both long-run and short-run comovement into account. Also, the codependence analysis explicitly incorporates the seasonal adjustment into the statistical model. Hence we use non-adjusted data in this section. For more detailed information on the methodology of codependence, we refer to Chapter 2. In this section, we consider quarterly real household consumption and real GDP the eight new EU member states (NMS-8) and 13 "old" EU countries, again covering 1995Q1-2005Q4.[7] Since we are mostly interested in short-term comovement of consumption and output, we omit the cointegration results at this stage and turn directly to the short-term analysis of common cycles.[8]

Since codependence operates with difference-stationary data, we conduct unit root tests for all data in levels and seasonal differences, employing the Dickey-Fuller General Least Squares (DF-GLS) unit root test by Elliot et al (1996). This test is a modified version of the Augmented Dickey-Fuller (ADF) test and involves transforming the time series via a generalised least squares regression. It has been shown that the DF-GLS test, as compared to the standard ADF test, tends to be substantially more powerful, i.e. it is more likely to reject the null hypothesis of a unit root when the alternative hypothesis of stationarity is true.[9]

The series of the NMS, presented in the upper panel of table C.1 in appendix C, reveal a considerable amount of unstability in the data. For consumption, five out of eight countries cannot be considered difference-stationary. Among the EU countries, Irish and Dutch consumption show non-stationary behaviour in differences. In the case of GDP, we cannot reject the unit root hypothesis for three NMS and three EU-13 countries. These countries basically disqualify for the codependence analysis. However, given the uncertainties of unit root testing with a relatively short time sample, all countries are tested for codependence with borderline cases receiving special attention.

Starting with consumption codependence, the results show again that comovement of consumption is weaker than that of GDP. Also, the relative comovement levels of EU-13 countries tends to be higher than that of the NMS.

[7] Greece and Portugal are not included due to data unavailability. For Ireland, the data span begins in 1997Q1.

[8] For the series under investigation, hardly any cointegration relations can be detected. Only France shows some indication of common stochastic trends with the euro area at the standard frequency. We do, however, find seasonal cointegration for a number of countries which hints at seasonal unit roots in the data and supports the idea of using non-adjusted figures.

[9] See Obstfeld and Taylor (2002), Stock and Watson (2003).

Table 4.3: Consumption codependenc results, NMS-8

Country	rank	Common features	Codependence Order 1	Codependence Order 2	Codependence Order 3
Czech Rep.	$m = 1$	34.17***	13.71***	17.90***	1.34
	$m = 2$	108.68***	24.44***	24.33***	5.91
Estonia	$m = 1$	21.77***	1.57	4.09	2.85
	$m = 2$	67.15***	11.81*	9.03	4.83
Hungary	$m = 1$	34.07***	10.70***	4.18	1.90
	$m = 2$	91.42***	21.75***	12.92**	4.77
Latvia	$m = 1$	24.90***	9.04	16.24***	6.34
	$m = 2$	110.43***	21.85**	23.92**	11.22
Lithuania	$m = 1$	11.95**	1.77	8.64	2.35
	$m = 2$	64.52***	12.89	16.69	4.86
Poland	$m = 1$	40.32***	9.60**	12.13**	5.20
	$m = 2$	105.60***	23.97***	21.28**	11.79
Slovakia	$m = 1$	22.63***	3.87**	0.57	0.00
	$m = 2$	60.11***	13.51***	5.34	2.37
Slovenia	$m = 1$	24.16***	1.88	15.27***	2.48
	$m = 2$	68.47***	12.28	23.01**	11.28

*Note: Codependence results of real private consumption of each country vis-à-vis the euro area. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector* ($m = 1$) *and rejecting a second vector* ($m = 2$)*, we conlude the existence of* $n - m = 2 - 1 = 1$ *common cycle.*

Table 4.3 reports the consumption codependence results of the NMS. We find no clear-cut case of common features or, in other words, codependence of zero order $CD(0)$. Hence, no NMS seems to have synchronised common consumption cycles with the euro area. Considering borderline cases, we note that for Lithuania the hypothesis of one common feature vector is rejected with a p-value of 0.02. Applying the 5 percent significance criterion, Lithuania does not qualify for a common feature - applying 1 percent, however, it does. Another borderline case is Slovakia which exibits codependence of first order, $CD(1)$, with a p-value of 0.049.

Table 4.4: Consumption codependenc results, EU-13

Country	rank	Common features	Codependence Order 1	Order 2	Order 3
Austria	$m = 1$	5.65	1.40	9.74**	1.86
	$m = 2$	56.02***	14.42*	18.42**	9.77
Belgium	$m = 1$	3.38	0.96	10.11**	2.29
	$m = 2$	61.04***	14.45**	18.22**	8.74
Denmark	$m = 1$	3.21	0.91	2.55	0.03
	$m = 2$	64.55***	11.25*	9.20	4.09
Finland	$m = 1$	19.69***	4.31	4.67*	1.62
	$m = 2$	68.51***	13.91**	10.48	6.78
France	$m = 1$	18.17***	3.47*	0.46	0.02
	$m = 2$	60.06***	13.56***	5.01	2.68
Germany	$m = 1$	23.60***	8.22**	6.41*	2.31
	$m = 2$	81.14***	21.78***	14.79*	16.18**
Ireland	$m = 1$	18.92***	3.53	0.60	0.58
	$m = 2$	91.40***	16.79**	7.84	6.41
Italy	$m = 1$	20.04***	4.23**	0.85	0.07
	$m = 2$	64.11***	13.99***	6.08	3.76
Luxembourg	$m = 1$	19.68***	5.84*	1.13	0.88
	$m = 2$	71.13***	15.18**	5.57	2.87
Netherlands	$m = 1$	6.91***	3.93**	4.51**	0.01
	$m = 2$	59.05***	16.77***	11.04**	3.79
Spain	$m = 1$	5.51**	0.43	0.00	0.46
	$m = 2$	43.35***	9.92**	4.78	1.94
Sweden	$m = 1$	32.48***	9.31	5.68	8.56
	$m = 2$	108.22***	24.78	17.99	19.73
UK	$m = 1$	20.33***	9.80	4.43	3.24
	$m = 2$	80.16***	24.61**	19.79	14.84

*Note: Codependence results of real private consumption of each country vis-à-vis the euro area. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector* ($m = 1$) *and rejecting a second vector* ($m = 2$)*, we conlude the existence of* $n - m = 2 - 1 = 1$ *common cycle.*

However, Slovakia's unit root test concluded an optimal autocorrelation lag length of 1 which would exclude any codependence of order higher than zero. Since the choice of the unit root lag length tends to be ambiguous, we consider Slovakia a candidate for $CD(1)$, i.e. common but non-synchronised consumption cycles with the euro area. In other words, the Slovak consumption cycles may not be perfectly synchronised with the one of the euro area but it may adjust after one lag period. On the whole, however, consumption codependence results for the NMS with the euro area are largely negative and the only indications of comovement are burdened with uncertainty.

Turning to consumption codependence of the EU-13 countries, the evidence is only slightly more favourable. Austria, Belgium and Denmark are the only clear cases of synchronised common consumption cycles with the euro area as table 4.4 makes clear. In all of these cases, the notion of one common feature vector cannot be rejected with p-values above the 0.10 threshold whereas second vectors are rejected at the 1 percent levels throughout. For Austria and Belgium, this is in line with the correlation results that indicated a large degree of consumption comovement for these countries with the euro area. Interestingly, Denmark shows signs of zero-order codependence whereas the consumption correlation results were rather poor. Other countries which were ascribed a high consumption correlation coefficient in the analysis above do not qualify for consumption codependence.

Neither France nor Germany exhibit synchronised common correlation cycles with the euro area. In the cases of France and Luxembourg, we find find evidence of non-synchronised common cycles, i.e. $CD(1)$. These results, however, depend on the true autoregressive order which may be 1 or 2. The Netherlands, on the other hand, would qualify for $CD(1)$ if they did not fail to be difference-stationary. Spain is another borderline case which hinges on the level of significance applied. In the standard case of the 5 percent level, it fails but it qualifies if we use the 1 percent criterion - the corresponding p-value for the rejectance of one common feature vector is 0.02. In sum, the consumption codependence results for both the NMS and the EU-13 countries with the euro area turn out to be weak, with the EU-13 slightly more positive than the NMS.

Table 4.5: GDP codependenc results, NMS-8 plus Turkey

Country	rank	Common features	Codependence Order 1	Order 2	Order 3
Czech Rep.	$m = 1$	27.01***	7.03	3.85	1.18
	$m = 2$	149.03***	22.30**	11.91	7.68
Estonia	$m = 1$	31.93***	11.78**	9.73**	2.99
	$m = 2$	96.98***	24.03***	16.09*	11.28
Hungary	$m = 1$	9.54*	0.96	5.96	4.23
	$m = 2$	44.78***	9.71	12.77	10.04
Latvia	$m = 1$	19.77***	1.66	0.05	1.65
	$m = 2$	75.83***	13.37***	5.54	5.90
Lithuania	$m = 1$	4.32	2.07	8.90**	1.69
	$m = 2$	58.35***	12.49	13.02	9.19
Poland	$m = 1$	11.61***	3.33*	1.73	0.47
	$m = 2$	48.14***	11.21**	12.42**	7.40
Slovakia	$m = 1$	13.70***	4.27	7.99**	3.76
	$m = 2$	58.40***	14.40*	13.34	8.16
Slovenia	$m = 1$	3.65*	7.31***	2.84*	0.49
	$m = 2$	41.74***	17.01***	9.29**	9.07*
Turkey	$m = 1$	22.37***	9.09**	3.88	06.69
	$m = 2$	60.28***	17.40**	15.61**	10.59

*Note: Codependence results of real GDP of each country vis-à-vis the euro area. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector ($m = 1$) and rejecting a second vector ($m = 2$), we conlude the existence of $n - m = 2 - 1 = 1$ common cycle.*

Not unexpectedly, the common GDP cycles are more pronounced. Table 4.5 provides the results for the NMS. Lithuania and Slovenia exhibit one common feature vector which indicates synchronised common cycles with the euro area. Hungary also qualifies according to the codependence test but the non-stationarity result for Hungary's GDP growth rates calls that result in question. The Czech Republic, Estonia, Latvia and Poland show signs of first-order codependence, i.e. common but non-synchronised common cycles. Slovakia, on the contrary, clearly fails both in terms of difference-stationarity and codependence. In addition to the above countries, we consider the EU candidate country Turkey but find no evidence of any codependence. In brief, the results on common GDP cycles of the NMS with the euro area are clearly better than in the case of consumption which tends to lend support to the consumption correlation puzzle.

Table 4.6: GDP codependenc results, EU-13

Country	rank	Common features	Codependence		
			Order 1	Order 2	Order 3
Austria	$m=1$	7.49*	0.58	3.73	3.92
	$m=2$	45.75***	8.30	7.67	12.83
Belgium	$m=1$	5.51**	1.89	0.01	0.28
	$m=2$	45.92***	10.78**	8.22*	11.04**
Denmark	$m=1$	35.01***	8.29**	3.32	7.99**
	$m=2$	77.65***	15.78**	15.90**	11.87
Finland	$m=1$	22.76***	6.21	2.71	1.93
	$m=2$	55.44***	14.25*	7.85	7.67
France	$m=1$	4.11**	1.17	1.74	0.03
	$m=2$	35.16***	7.87*	7.83*	2.09
Germany	$m=1$	0.07	0.00	1.99	0.09
	$m=2$	37.37***	7.60	4.01	6.28
Ireland	$m=1$	0.20	1.90	0.12	1.26
	$m=2$	32.77***	8.76*	2.24	4.80
Italy	$m=1$	11.55***	1.52	0.01	0.00
	$m=2$	44.41***	10.21**	6.54	9.93**
Luxembourg	$m=1$	7.29**	0.97	21.91***	1.47
	$m=2$	59.10***	12.63**	32.40***	15.03**
Netherlands	$m=1$	15.78***	3.26**	0.94	0.06
	$m=2$	72.20***	15.31***	8.44	6.30
Spain	$m=1$	2.21	0.37	0.00	0.58
	$m=2$	38.12***	7.01	4.09	4.56
Sweden	$m=1$	12.35***	2.57	0.58	0.03
	$m=2$	46.91***	9.67**	3.29	0.59
UK	$m=1$	0.01	0.04	0.47	0.05
	$m=2$	32.48***	7.42	4.62	9.19*

*Note: Codependence results of real GDP of each country vis-à-vis the euro area. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector ($m = 1$) and rejecting a second vector ($m = 2$), we conlude the existence of $n - m = 2 - 1 = 1$ common cycle.*

Next we turn to GDP codependence of the EU-13 countries vis-à-vis the euro area, see table 4.6. Again, we generally find a larger degree of GDP than consumption comovement. Austria, Germany and the UK qualify for synchronised common GDP cycles with the euro area. Borderline cases for $CD(0)$ are Belgium, France and Luxembourg for which the p-value of rejecting one common feature vector is below 0.05 but above 0.01. Ireland and Spain seem to qualify for common features but both suffer from non-stationarity results in the unit root test. Italy and Sweden seem to have common but non-synchronised cycles with the euro area, i.e. they exhibit one codependence vector of order one. This $CD(1)$ result holds also true for Belgium and Luxembourg who were considered borderline for $CD(0)$. These results largely correspond with the correlation evidence concerning Austria, Belgium, France, Germany and the UK. For other countries, the codependence results tend to be weaker than the correlation evidence. However, simple correlations do not provide a clear benchmark threshold and are a more simplistic concept per se.

Summing up, we make two general observations. First, the degree of consumption comovement tends to be weaker than that of GDP comovement which, at first glance, hints at a low degree of risk sharing. However, the rolling correlations seem to indicate that both consumption and GDP comovement vis-à-vis the euro area have been increasing over the recent years, for both the NMS and the "old" EU countries. Considering the argumentation of Imbs (2006) who sees increased financial integration behind the rise of both consumption and GDP comovement, we may not draw unequivocal conclusions from our evidence on the consumption correlation puzzle on the underlying degree of risk sharing.

The second observation pertains to the fact that the overall levels of consumption and GDP comovement to the euro area tend to be larger among the EU-13 countries than among the NMS. This is not surprising given the longer integration history among the "old" EU and the fact that most EU-15 are actually included in the euro area aggregate. It may indicate, however, that with ongoing economic integration, the obstacles to risk sharing may continue to shrink and hence the improve the future perspective of risk sharing among the member states of the enlarged EU. To shed more light on the degree and dynamics of financial integration, we now turn to the analysis of real interest rate comovement.

4.2 Financial integration

The argument of Mundell II postulates that, in the presence of financial integration, countries with asymmetric business cycles benefit most from joining a currency union because consumers can diversify their portfolios across the region and decouple their consumption patterns from potentially idiosyncratic output cycles at home. The previous section presented evidence that the degree of risk sharing, measured by consumption comovement, is to date limited in the NMS. The "old" EU members, however, enjoy a larger degree of risk sharing which is a likely result from their common integration history.

This section investigates financial integration for both the NMS and the EU-15 countries. It finds that the degree of financial integration as measured by real interest rate comovement is limited for the NMS. The EU-15 countries have, however, made considerable progress in financial integration from the 1980s to the 1990s. This development can be expected to have contributed to higher levels of risk sharing and may be anticipated for the NMS as they continue to integrate with the EU-15.

One way to measure financial integration is to compare cross-country interest rates. If financial markets are integrated, identical financial assets should have the same price whether they are traded at home or abroad. As a result, we would expect to see equalised real interest rates between countries that share a perfect financial market. Various concepts capture the different dimensions of interest parity. Uncovered interest parity states that differences in nominal returns across countries should equal expected exchange rate changes. Covered interest rate parity uses the forward rate instead of spot rates. According to real interest parity, the expected difference between domestic and foreign real interest rates is zero. We follow Kugler and Neusser (1993) who investigate long-run and short-run comovement of real interest rates across countries using the codependence technique. While they foucs on pairwise codependence among five G7 countries and Switzerland, we consider the 23 countries of the enlarged EU vis-à-vis the euro area aggregate. Before conducting the correlation and codependence analyses, we discuss the ambiguous issue of stationarity in the context of interest rates.

4.2.1 Interest rates and stationarity

It has been an issue of debate whether interest rates should be regarded as stationary or non-stationary. A stationary time series is characterised by constant expected mean and variance and is hence considered mean-reverting. For consumption and GDP, the case seems clear: Most countries exhibit long-run positive trends which turn the series non-stationary. Growth rates or cyclical components, however, tend to be stationary, i.e. they fluctuate around a constant mean and have a finite variance.

The case of interest rates is less clear. In theory, the life cycle model of consumption predicts consumption growth rates to have similar time-series properties as real interest rates.[10] Hence, interest rates would be expected to be stationary, similar to consumption growth rates. But empirical evidence on interest rate stationarity is mixed. Kugler and Neusser (1993) confirm the theoretical proposition for their 1980s sample of industrialised countries and find that the unit root hypothesis can be easily rejected. Rose (1988), in contrast, suggests that interest rates in the U.S. and elsewhere tend to be non-stationary. Also, Obstfeld et al. (2005) find that, at least during the post-Bretton Woods era, interest rates are overwhelmingly non-stationary. However, they admit that interest rates are unlikely to follow a literal unit root process - otherwise we would see interest rates rise unboundedly. This is hardly the case. Driffell and Snell (2003) propose that

[10] See Kugler and Neusser (1993).

the unit root result my stem from the high persistency of interest rates and not from a truely non-stationary process. Moreover, they argue that what seems like a unit root process may often be a result of regime shifts in otherwise stationary data. Also, Garcia and Perron (1996) make this point and treat interest rates as stationary.

In our dataset of the enlarged EU, evidence on stationarity is mixed and thus reflects the ambiguity of the literature. The following sections present the unit root test results in the context of the correlation and codependence analyses. The corresponding tables can be found in appendix C.

4.2.2 Interest rate correlation

We consider quarterly short-term interest rates for the eight NMS and the EU-15 countries. We employ three-months money market rates from the IMF's International Financial Statistics, supplemented by Eurostat data. All data are deflated by CPI.[11] In the case of the NMS, our time frame is 1995Q1-2005Q4 and we pair each country with the euro area aggregate. For the EU-15 countries, we apply the pre-EMU time frame 1980Q1-1998Q4 which we divide into two subsamples at 1990Q1.[12] We use Germany as the reference country for the EU-15 countries because it served as benchmark and role model in the run-up to EMU.

Given the ambiguous stationarity situation for interest rates, we first conduct unit root tests for all real interest rate series. Tables C.2 and C.3 summarise the results for the NMS and the EU-15 for their respective time frames in levels and first differences.[13]

[11] Although the Harmonised Index of Cosumer Inflation (HICP), compiled by Eurostat, would be preferable for the comparison of European countries, it is not available for all countries in all periods. Hence, we resort to the commonly used consumer price index (CPI), provided by IFS. Quarterly inflation rates are calculated on a year-on-year basis and then subtracted from the quarterly nominal interest rate. Following Obstfeld and Taylor (2002), we make the standard assumption that the observed *ex post* real interest rates are equal to the *ex ante* real rate plus a white-noise stationary forcast error.

[12] We analyse only pre-EMU data because with the start of the single monetary policy, nominal short-term interest rates are equalised across the euro area. Hence, real interest differentials would only be due to inflation differentials which are, in itself, not a prime measure of financial integration.

[13] We calculate the first differences from the interest rate levels, not logs. The reason is that logs cannot be computed for negative real interest rates which tend to prevail for quite a number of observations.

The evidence is irregular. Some of the NMS seem stationary in levels whereas for others, the test cannot reject the hypothesis of a unit root. In differences, all countries but Lithuania seem stationary at least on the 10 percent level of significance. In case of the EU-15 countries, only five countries show stationary behaviour in levels but in nearly all cases, the differences are stationary. For France and Ireland, we cannot reject a unit root either in levels or in differences. Given the ambiguity of interest rate stationarity, we present correlation results for both levels and differences in the following.

Table 4.7: Real interest rate correlation, NMS-8

Country	Levels			First differences		
	95-05	95-99	00-05	95-05	95-99	00-05
Czech Rep.	0.55	-0.20	0.32	0.20	0.26	0.16
Estonia	-0.43	-0.53	0.42	0.33	0.45	0.31
Hungary	-0.44	-0.53	-0.61	0.29	0.15	0.40
Latvia	0.02	-0.38	0.75	0.21	0.13	0.39
Lithuania	-0.34	-0.80	0.53	-0.02	-0.15	0.20
Poland	0.18	-0.29	0.87	0.28	0.31	0.29
Slovakia	0.50	-0.08	0.07	0.26	0.27	0.33
Slovenia	-0.05	-0.02	-0.77	-0.15	-0.36	0.25

Note: Correlation coefficients of real interest rates vis-à-vis the euro area.

Table 4.7 presents the correlation coefficients of NMS real interest rates vis-à-vis those of the euro area aggregate. We calculate correlation coefficients of levels and differences for the entire 1995Q1-2005Q4 period as well as for two sub-periods, 1995Q1-1999Q4 and 2000Q1-2005Q4. In the levels case, we observe correlation coefficients of up to 0.55 for the Czech Republic and 0.50 for Slovakia. Three out of the eight countries show negative coefficients: Estonia, Hungary and Lithuania. Comparing the two sub-samples, it becomes clear that, except for Hungary and Slovenia, all countries exhibit increasing correlation which may hint at improved financial integration with the euro area. The correlation coefficients of the first differences are less dispersed. Generally, all coefficients

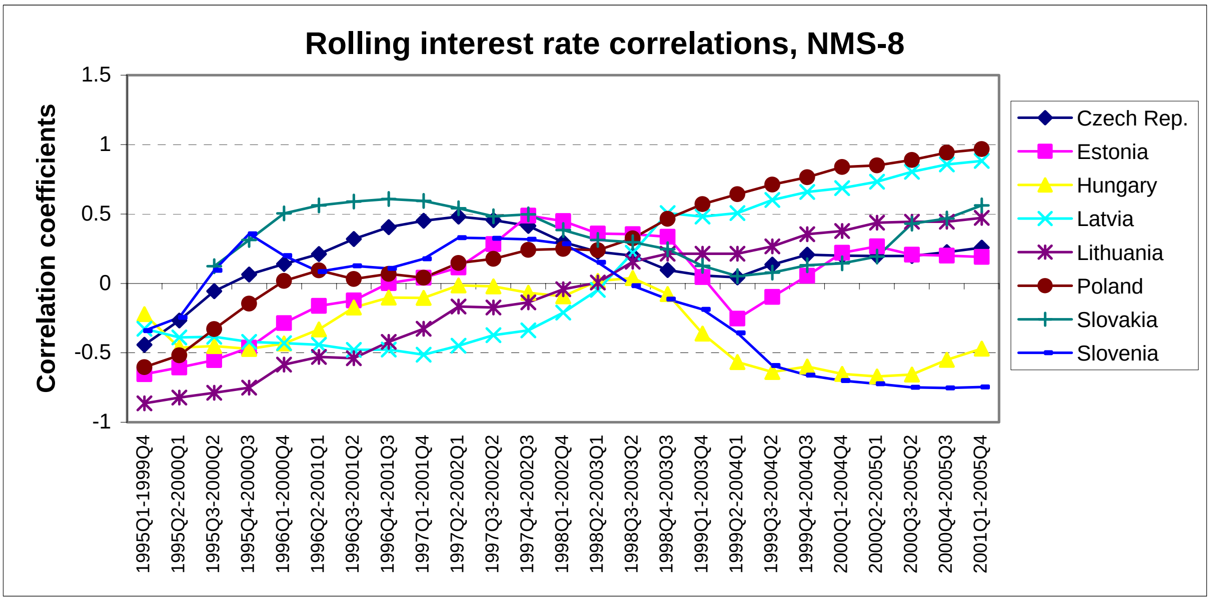
Using instead the logs of the interest rate factors, log(1+R), as suggested by Obstfeld et al. (2005), would yield factor growth rates when differenced. Their correlation coefficients, however, are almost equal to those of the simple first differences of non-log levels since $d[\log(1+R)] \approx d[R]$ for small Rs.

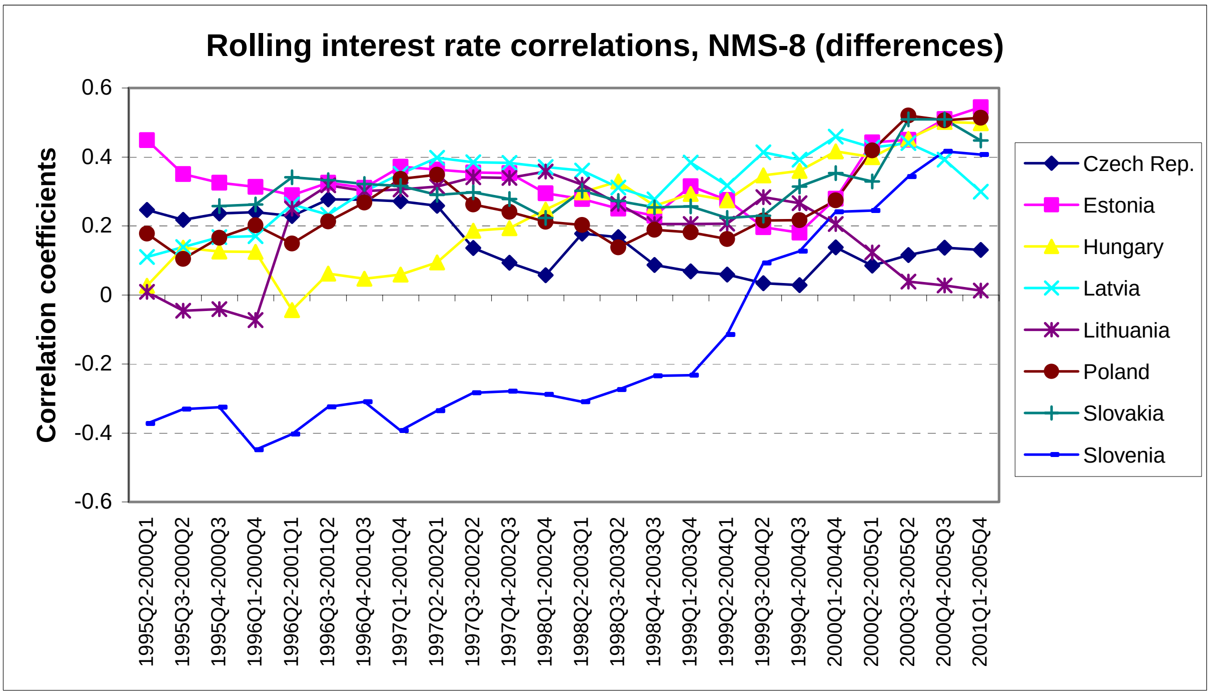
remain below 0.50 but we find only two negative correlations. Now, Estonia and Hungary are among the countries with the highest correlation.[14] Poland and Slovakia still exhibit a relatively large degree of interest rate correlation. Surprisingly, Slovenia's coefficient is now negative. Still, most coefficients tend to rise or remain relatively stable from the first to the second sub-period. They shrink in only two cass, the Czech Republic and Estonia.

To find out more about variations over time, we calculate moving correlation windows of five years length. Figures 4.5 and 4.6 present those in line graphs for both levels and differences. The levels tend to increase strongly over the considered period, ranging across almost the entire spectrum of -1 to 1. Only Hungary and Slovenia stand out with negatively sloped lines. The differences, graphed in figure 4.6, tend to move closer together and range between -0.5 and 0.6. Although most countries experience rising coefficients on the whole, the increase appears less dramatic.

Another way of analysing real interest rate comovement is looking at bilateral differentials. Figure 4.7 presents bilateral differentials of the eight NMS, each paired with the euro area. While most differentials experience enormous fluctuation over time, it seems that some countries achieved more stability since approximately 1999/2000. In particular, the currency boards of Estonia and Latvia seemed to have contributed to this development.

[14] In the case of Estonia, the lack of stationarity in levels may render the corresponding correlation result invalid.





Figures 4.5-4.6: 5-year rolling correlation windows of short-term real interest rates against the euro area, based on quarterly data in levels and first differences, respectively.

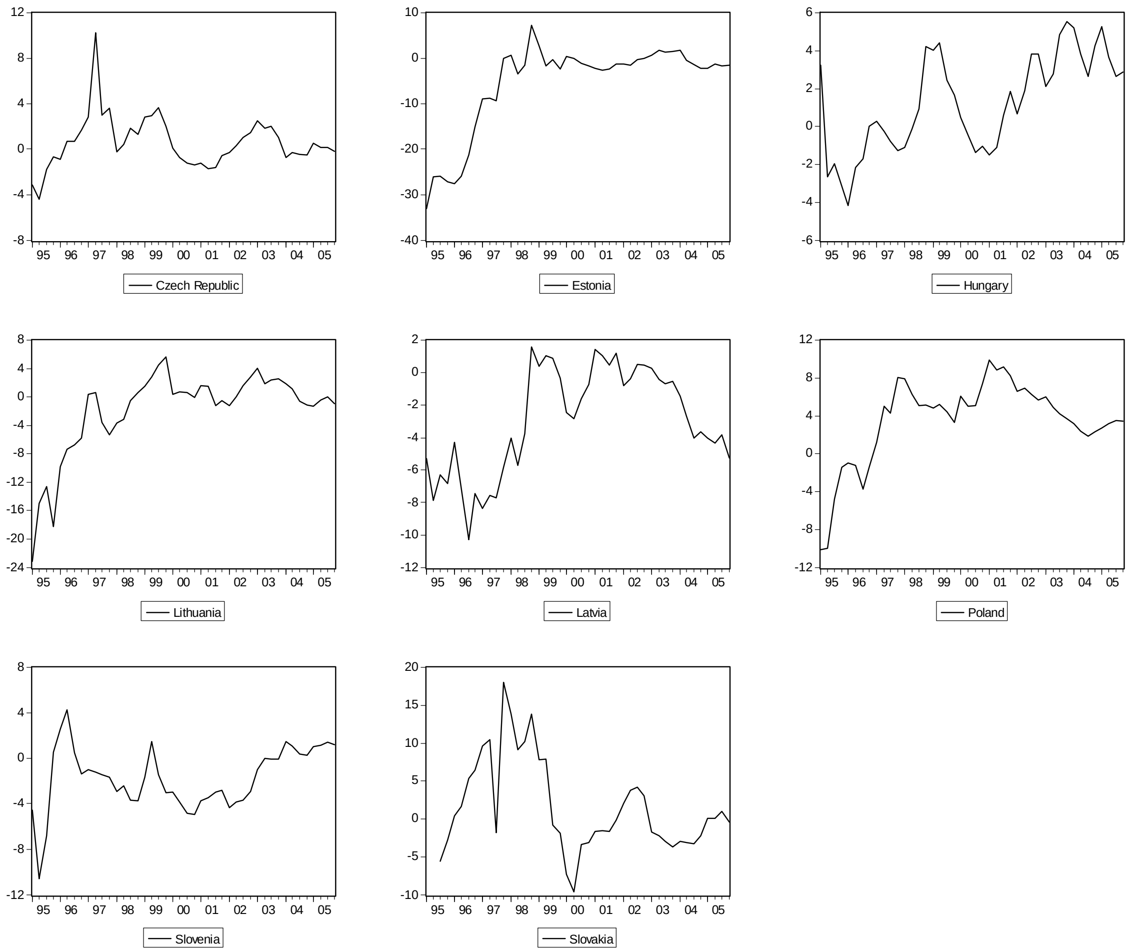


Figure 4.7: Bilateral short-term real interest rate differentials, each country minus the euro area, 1995Q1-2005Q4.

To analyse the variability of real interest differentials further, we calculate the standard deviations for the whole period and for the two sub-periods, 1995-1999 and 2000-2005. The results are displayed in figure 4.8. All NMS are characterised by decreasing variation in their interest rate differentials with the euro area. While standard deviations vary considerably during the first sub-period, they seem to converge to a simliar low level during the second. We regard this as additional indication for increased financial integration.

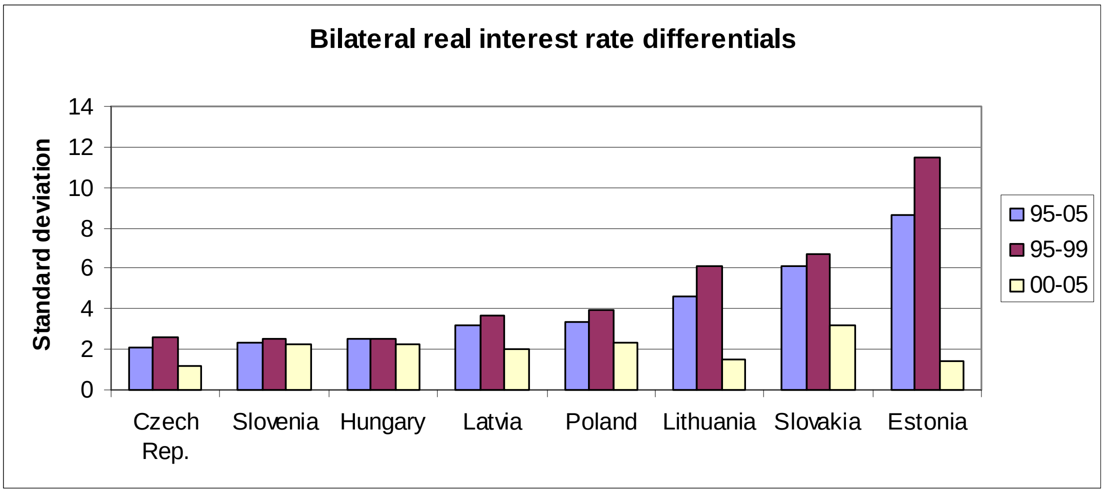


Figure 4.8: Standard deviations of real interest rate differentials vis-à-vis the euro area.

For comparison, we investigate real interest rate correlation and variation of differentials for the EU-15 countries. Now we focus on the pre-EMU period where countries converged towards the benchmark country of those years, Germany. We again split our series into two sub-samples, now ranging from 1981Q1-1989Q4 and 1990Q1-1998Q4.

Table 4.8 provides the correlation coefficients for both levels and differences, each country paired with Germany. Austria, Belgium, Luxembourg and the Netherlands seem to form a core group and experience by far the largest correlation coefficients. This applies for both levels and differences, although the values for the differences tend to be lower on the whole. The smallest coefficients pertain to Greece and Portugal, followed by Ireland and Spain. Almost every country's correlation with Germany increases markedly from the first to the second sub-period, although again the effect is stronger in case of levels. Interestingly, the correlation coefficients for the UK tend to be increasing towards Germany while they go down vis-à-vis the United States which we explore as an additional benchmark.

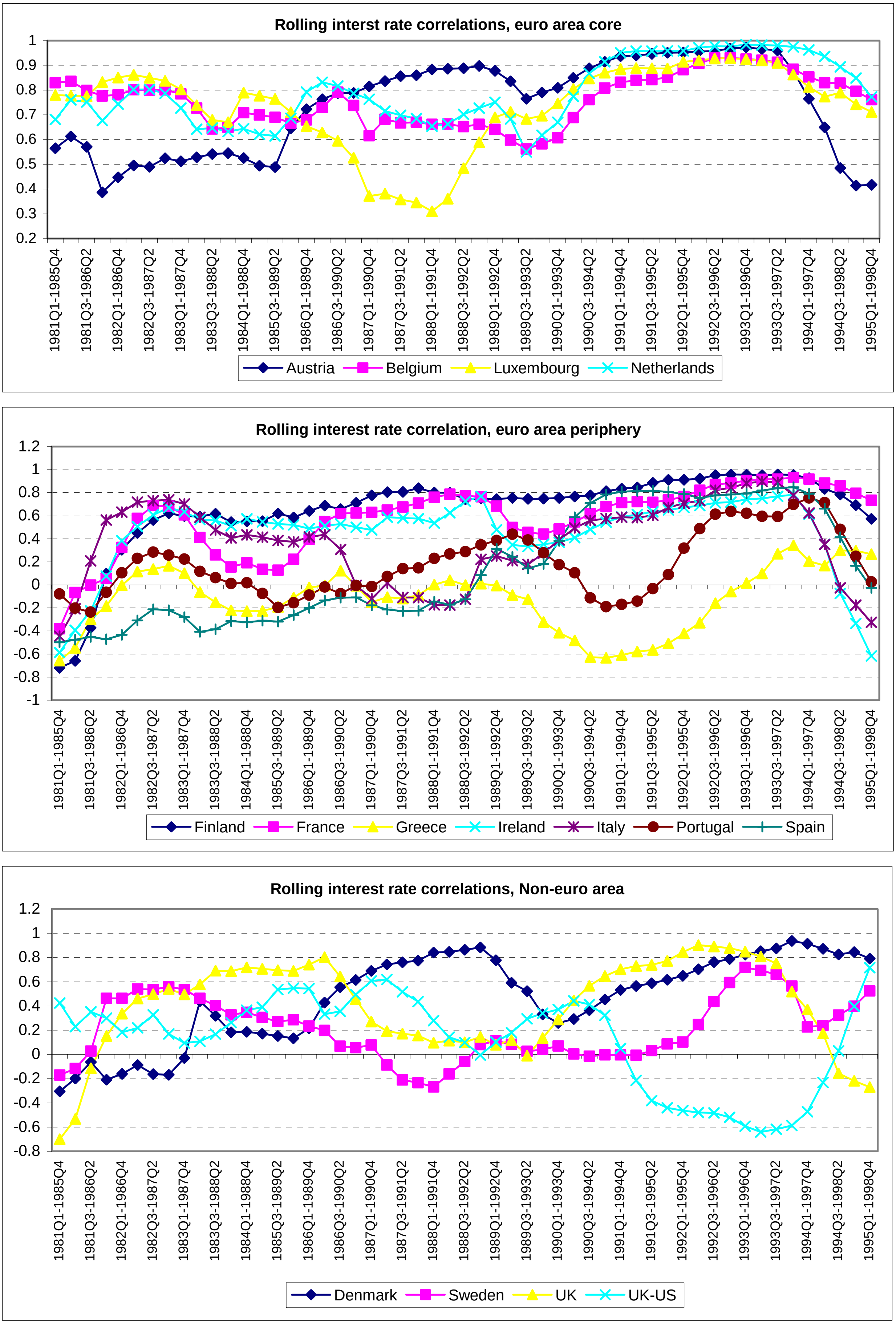
Table 4.8: Real interest rate correlation, EU-15

Country	Levels			First differences		
	81-98	81-89	90-98	81-98	81-89	90-98
Austria	0.78	0.60	0.94	0.44	0.50	0.31
Belgium	0.85	0.79	0.92	0.47	0.48	0.46
Denmark	0.50	-0.20	0.83	0.12	0.11	0.13
Finland	0.42	-0.38	0.91	0.13	-0.15	0.53
France	0.49	-0.13	0.87	0.26	0.28	0.24
Greece	-0.34	-0.31	-0.38	0.02	-0.16	0.18
Ireland	0.17	-0.27	0.70	0.19	0.25	0.10
Italy	0.21	-0.24	0.63	0.28	0.11	0.54
Luxembourg	0.80	0.72	0.94	0.48	0.50	0.41
Netherlands	0.83	0.58	0.97	0.59	0.59	0.62
Portugal	0.00	-0.01	0.33	-0.12	-0.13	-0.11
Spain	0.20	-0.35	0.85	-0.13	-0.25	0.30
Sweden	0.16	0.03	0.24	0.29	0.08	0.50
UK	0.22	-0.28	0.74	0.15	0.12	0.22
UK-US	0.20	0.25	-0.30	0.14	0.19	-0.05

Note: Correlation coefficients of real interest rates vis-à-vis Germany.

Turning to rolling interest rate correlation windows, we split the country sample into three groups to facilitate graphical inspection. For many countries, correlation of interest rates with Germany seems to move in cycles. Figure 4.9 includes those countries with the largest overall correlation coefficients, the euro area core. These are relatively small countries which have maintained close ties to German monetary policy for many years. They tend to experience "correlation booms" during the late 1980s and the mid-1990s, interrupted by downturns around 1990 and in the most recent periods.

The 1990 trough is likely to be due to German reunification which was associated with exceptionally high interest rates in Germany compared to the rest of Europe. On the whole, the core group fluctuates within a relatively narrow band of 0.40-0.90. Figure 4.10 shows the remaining euro area economies. These "periphery" countries show a larger degree of convergence as they all start at negative correlation values and increase drastically from there, some appraoching 0.95 in the mid-1990s. Again, we observe a

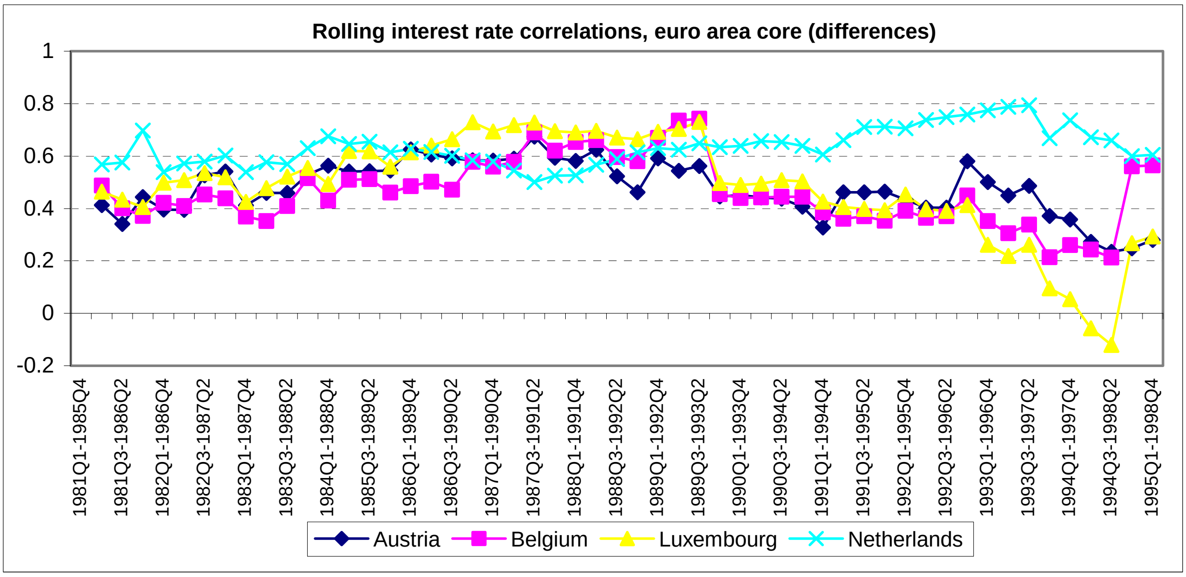


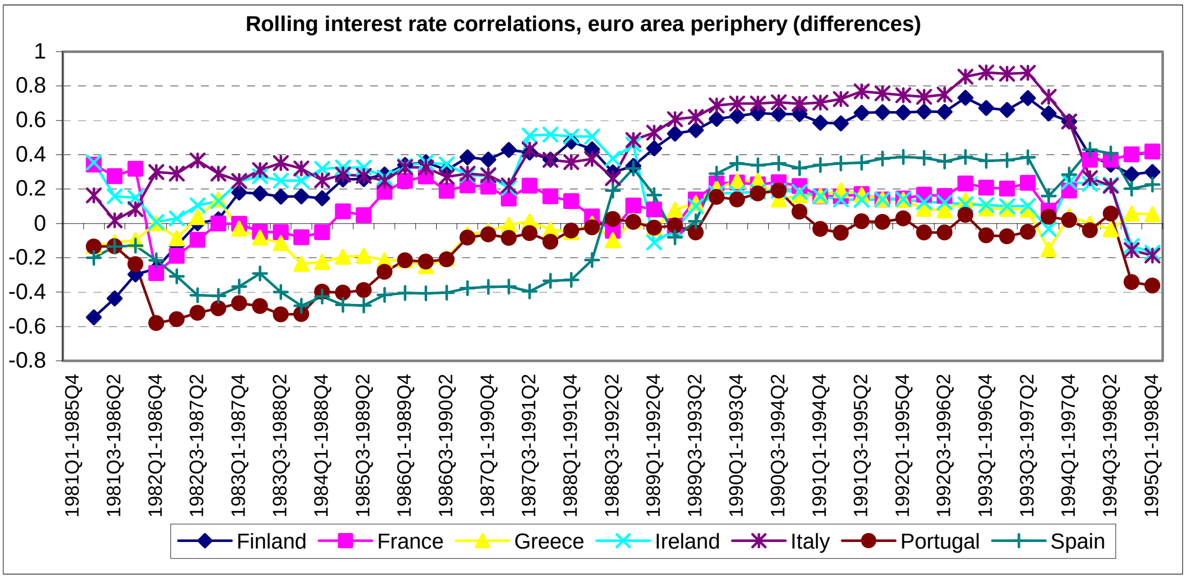
Figures 4.9-11: 5-year rolling correlation windows of short-term real interest rates against Germany, based on quarterly data in levels.

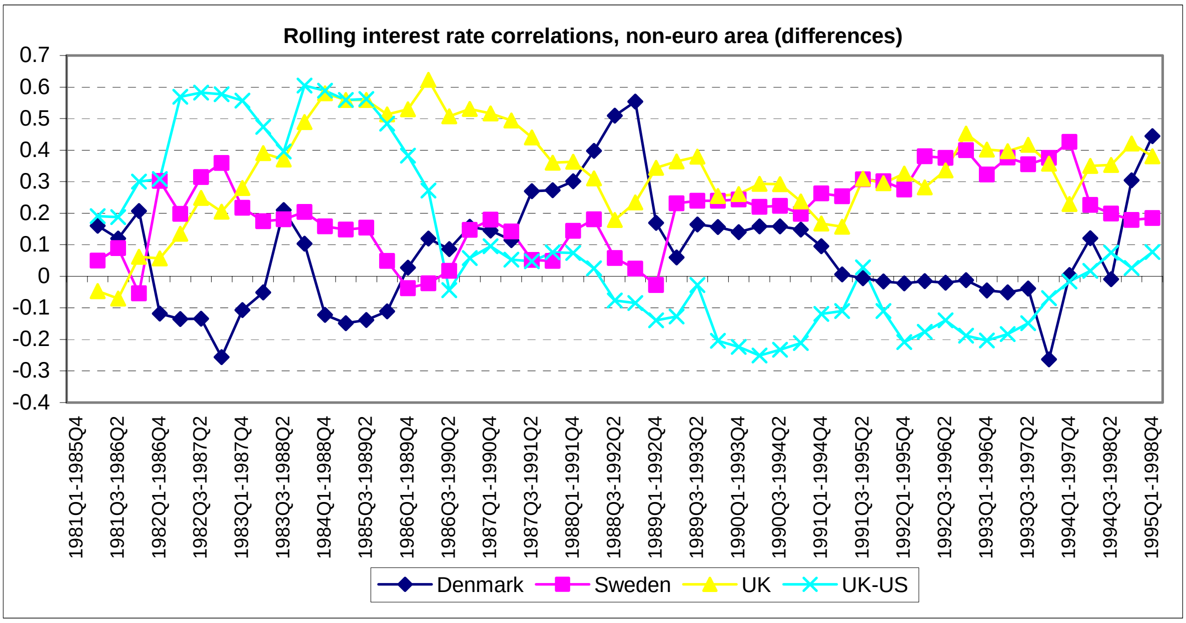
certain cyclical behaviour and a downturn of correlation values towards the end of the sample. Figure 4.11 consists of the three non-euro area countries among EU-15, plus the UK-US relation for comparison. While Denmark and Sweden increased markedly in their interest rate correlations with Germany, the UK pattern against the US seems to mirror that vis-à-vis Germany. During the mid-1990s, UK-German interest rates tend to comove on a high level but turn negative in the most recent period whereas UK-US correlation remains negative during most of the 1990s and picks up towards the end of the sample. Apparently, the UK takes a changing position in terms of financial market integration with Germany and the US. We note, however, that neither for Germany, nor for the UK and the US we can formally reject the unit root hypothesis in levels.

The line graphs of the difference correlations, illustrated in figures 4.12-4.14, tend to follow a roughly similar pattern. Again, the core group fluctuates on a relatively high level against Germany while the periphery countries exhibit a clearer upward trend. Regarding the non-euro area countries, it stands out that the UK-German correlations remain above the UK-US relation at all times since the end of the 1990s. Taken together, the EU-15 countries seem to have followed German real interest rates to an increasing extent during the pre-EMU period which hints at improved financial market integration and policy coordiation in preparation for the euro.

To study the variability of bilateral interest rate differentials between the EU-15 countries and Germany, we first inspect the differentials graphically, see figure 4.15. Although all series seem to include considerable variation, some appear to narrow down in the second half of the sample. Austria, Belgium, Luxembourg and the Netherlands basically even out at a plus/minus one percentage level since around 1993. Other countries, such as France or the UK, tend to remain within a virtual plus/minus two percentage band towards the end of the sample.







Figures 4.12-14: 5-year rolling correlation windows of short-term real interest rates against Germany, based on quarterly data in first differences.

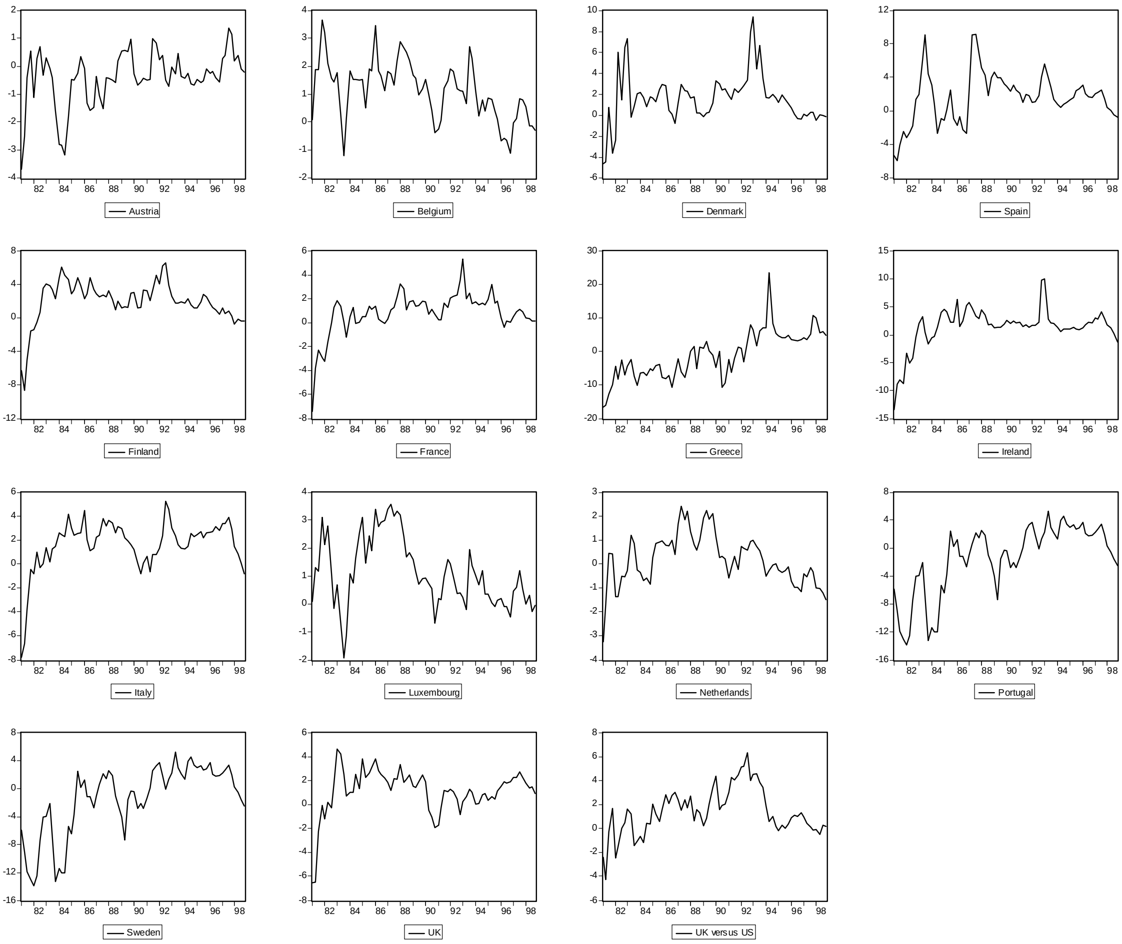


Figure 4.15: Bilateral short-term real interest rate differentials, each country minus Germany, 1980Q1-1998Q4.

Figure 4.16 ranks countries according to the standard deviation of their bilateral interest rate differentials against Germany. The aforementioned euro area core group plus France and the UK lead the list of smallest variations over the whole sample. However, other countries which experienced a high level managed to reduce their variation considerably. As a result, except for Greece all EU-15 countries brought their variation level down to around two standard deviations - which is a similar level as for the NMS versus the euro area since 2000.

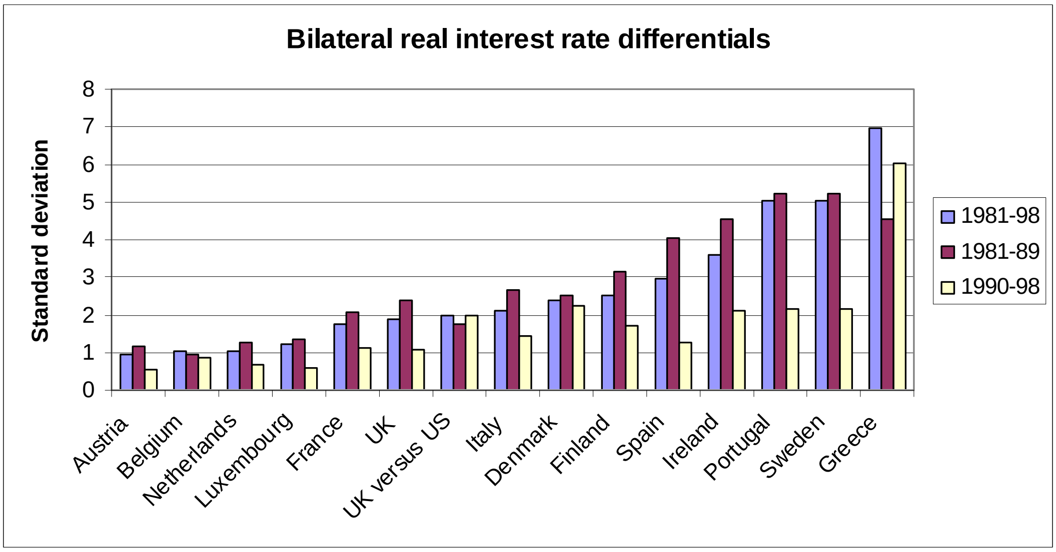


Figure 4.16: Standard deviations of real interest rate differentials vis-à-vis Germany.

Finally, we study the degree of dispersion across EU-15 countries. For this purpose, we calculate the standard deviation across all EU-15 countries at one point in time and repeat this exercise for all periods. Analysing the cross-country dispersion over time has also been known as sigma convergence in the empirical growth literature. Figure 4.17 shows that despite some peaks, the overall dispersion level has clearly decreased between 1981 and 2005. One major spike stands out in 1994 which is due to idiosyncratic developments in Greece. Leaving out Greece (EU-14) delivers an even smoother path of decreasing dispersion.

In summary, our correlation, variability and dispersion evidence suggests that real interest rates have become more similar during the 1980s and 1990s in the EU-15. Although correlations for the NMS tend to be ambiguous, the reduced variability of bilateral interest rate differentials hints at more similar rates as well. We acknowledge that the stationarity analysis of interest rate is subject to limitations and has delivered mixed results. This is, however, in line with the conflicting propostions on stationarity of interest rates in the literature.

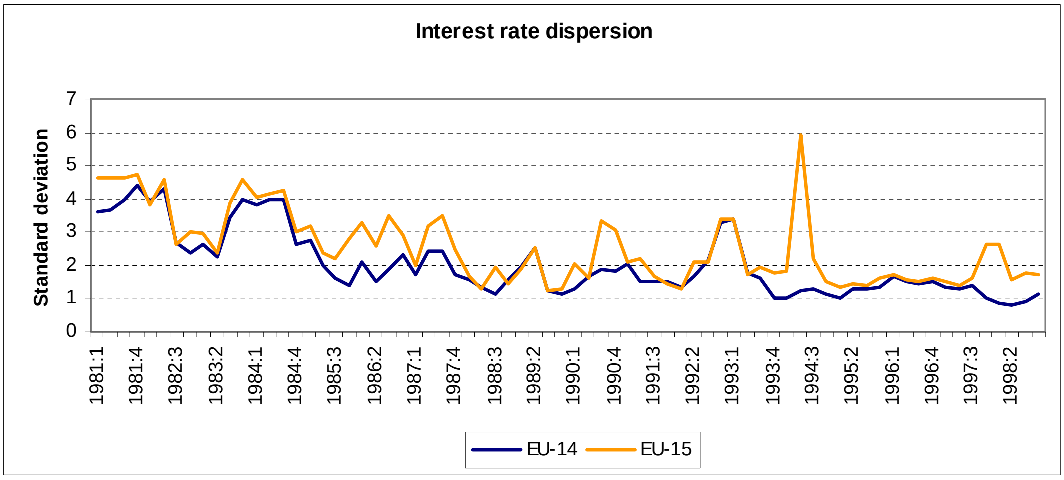


Figure 4.17: Standard deviation of real interest rates across the EU-15 countries (EU-14: excluding Greece), at every point in time.

4.2.3 Interest rate codependence

In the following, we analyse real interest rates across countries employing the codependence technique. Again, as our focus is on the short run, we concentrate on the cyclical part of comovement. We employ the same time and country sample as in the correlation analysis of interest rates, i.e. eight NMS vis-à-vis the euro area during 1995Q1-2005Q4 as well as the EU-15 countries related to Germany during 1980Q1-1998Q4. For comparison, we again consider the relation of the UK to the US. Since the codependence framework incorporates a seasonal adjustment tool, we use annualised month-on-month CPI inflation to calculate "non-adjusted" real interest rates.

First, we look at unit root test results for the non-adjusted data, see tables C.4 and C.5, which are largeley similar to the adjusted data. We note that nearly all countries are stationary in differences while a few seem stationary in levels as well. For some countries, however, we cannot reject the unit root hypothesis either in the levels or in the difference cases.

On the whole, the unit-root results are again subject to debate, as discussed in the previous section, since interest rates are hard to imagine non-sationary in the classical sense. The codependence approach works with differences and almost all countries are stationary at least in differences.

Table 4.9: Interest rate codependenc results, NMS-8

Country	rank	Common features	Codependence Order 1	Order 2	Order 3
Czech Rep.	$m = 1$	16.38***	2.17	4.10	5.73
	$m = 2$	48.36***	19.89**	28.13***	20.64***
Estonia	$m = 1$	23.61***	10.66*	16.68***	11.85**
	$m = 2$	89.00***	30.75***	57.15***	30.72***
Hungary	$m = 1$	1.05	2.57	1.64	2.97
	$m = 2$	36.22***	21.15***	27.23***	21.57***
Latvia	$m = 1$	17.30***	0.99	17.41***	2.00
	$m = 2$	56.89***	31.80***	47.45***	12.99
Lithuania	$m = 1$	14.38***	4.02	20.19***	7.38
	$m = 2$	72.44***	25.48***	53.44***	22.90**
Poland	$m = 1$	3.30*	1.22	0.08	1.37
	$m = 2$	20.17***	4.22	7.79*	16.05**
Slovakia	$m = 1$	5.05	4.03	4.92	5.50
	$m = 2$	30.93***	14.22	20.78**	18.91**
Slovenia	$m = 1$	28.52***	7.97	19.20	12.56*
	$m = 2$	67.49***	35.71***	50.48***	34.60***

*Note: Codependence results of real interest rates of each country vis-à-vis the euro area. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector ($m = 1$) and rejecting a second vector ($m = 2$), we conlude the existence of $n - m = 2 - 1 = 1$ common cycle.*

Table 4.9 provides the codependence results for the NMS. Hungary, Poland and Slovakia seem to exhibit common features, or codependence of order zero, with the euro area. This would mean that their real interest rates have synchronised common cycles which hints at a high degree of financial integration. This evidence matches with the correlation of differences from the previous section where these three countries were among those with the largest correlation coefficients. Some uncertainty, however, remains concerning the autoregressive orders. Intuitively, two countries can only have a common feature if the individual features, i.e. serial correlations, are of equal length. Otherwise, the feature would not cancel out in the linear combination. For the euro area, we found an autoregressive order of $p = 4$ when testing for Q statistics of autocorrelation in the residuals of the autoregressive equations. For Hungary, this criterion yields $p = 1$ although according to the modified Akaike information criterion, lag length 4 would be the optimal choice. The fact that Hungary displays one codependence vector from order zero to three throughout supports the notion that Hungary acutally does qualify for $CD(0)$. For Poland and Slovakia, the cases are less clear. The unit root tests would also allow $p = 4$ but the fact that the codependence tests suggest two codependence vectors for $CD(1)$ renders the case of synchronised common cycles rather unlikely.

The remaining countries have no common feature vectors but all have one $CD(1)$ vector. This indicates common but non-synchronised interest rate cycles which means that the countries would respond to euro area interest rates with a time lag of one quarter. However, the autoregressive orders for the unit root tests are again unclear. Hence, we conclude that the degree of financial integration between the NMS and the euro area is at best intermediate.

Codependence results for the EU-15 countries are provided by tables 4.10 and 4.11. We divide the sample into two the sub-groups 1980Q1-1989Q4 as well as 1990Q1-1998Q4 hoping to learn more about changes in financial integration among the EU-15 over time. During the 1980s, the real interest rates of Austria, France and the Netherlands seem to be synchronised with those of Germany. For France and the Netherlands, however, the autoregressive lag differs from that of the Germany. This does not exclude the possiblity of common features, given the ambiguity of the lag length choice, but it adds uncertainty to the results.

Table 4.10: Interest rate codependenc results, EU-15, 1980-1989

Country	rank	Common features	Codependence Order 1	Order 2	Order 3
Austria	$m = 1$	1.48	1.84	0.04	2.60
	$m = 2$	17.23***	8.47	13.95**	12.30*
Belgium	$m = 1$	1.47	2.18	4.10	3.65
	$m = 4$	33.45***	6.32	30.22***	16.58
Denmark	$m = 1$	18.63***	4.93	11.28**	0.93
	$m = 2$	51.11***	15.38	32.60***	18.44**
Finland	$m = 1$	15.06***	3.24	10.11**	4.34
	$m = 2$	59.68***	19.14**	33.79***	16.49*
France	$m = 1$	12.88*	9.61	10.25	11.09
	$m = 2$	56.29***	37.37***	42.54***	40.21***
Greece	$m = 1$	13.37***	0.38	8.36**	2.15
	$m = 2$	34.52***	11.26	27.79***	14.60*
Ireland	$m = 1$	12.40***	5.94	2.19	4.37
	$m = 2$	71.74***	24.13***	18.14**	24.44***
Italy	$m = 1$	19.26***	3.86	15.56***	5.56
	$m = 2$	64.92***	16.41*	32.23***	32.52***
Luxembourg	$m = 1$	13.08***	1.07	11.06**	3.57
	$m = 2$	35.54***	6.77	26.94***	10.23
Netherlands	$m = 1$	4.44	4.89	6.82	3.78
	$m = 2$	19.35*	19.66*	23.27**	14.91
Portugal	$m = 1$	16.43***	1.42	10.70**	5.60
	$m = 2$	37.15***	7.38	28.15***	21.71***
Spain	$m = 1$	15.91***	1.95	5.35	5.90
	$m = 2$	51.67***	12.72	25.46	24.66***
Sweden	$m = 1$	19.52***	2.62	8.10	2.86
	$m = 2$	46.87***	12.04	27.53***	21.02*
UK	$m = 1$	9.52**	2.57	1.02	6.38*
	$m = 2$	40.21***	13.31	19.08**	21.35***
UK vs. US	$m = 1$	37.28***	16.45***	27.91***	13.86**
	$m = 2$	83.78***	38.07***	46.65***	25.92**

*Note: Codependence results of real interest rates of each country vis-à-vis Germany. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector ($m = 1$) and rejecting a second vector ($m = 2$), we conlude the existence of $n - m = 2 - 1 = 1$ common cycle.*

The UK is a borderline case in which the hypothesis of one common feature vector is rejected with a p-value of 0.02. In addition, results indicate one common feature vector for Belgium. We can rule this out, however, since we were not able to detect stationarity for Belgium's interest rate differences. Codependence of first order is indicated for Finland and Ireland while the latter is disqualified by its unsatisfactory difference-stationarity result. All other countries reveal no signs of codependence vis-à-vis Germany. This holds also true for the UK-US relation.

Turing to the 1990s, we find more favourable results. Austria, Greece, Ireland, Italy, Luxembourg, the Netherlands and the UK have one common feature vector and thus synchronised common interest cycles with Germany. Out of these, only Luxembourg does not fulfill the difference-stationarity criterion. The UK-US relation has one codependence vector for $CD(1)$ and thus shares a common but non-synchronised cycle. Denmark seems to be $CD(1)$ but fails to be difference-stationary. All remaining countries do not exhibit clear results.

On the whole, the financial integration evidence for the EU-15 is not overwhelming but appears to be increasing over time. During the 1990s, more countries seem to share a common interest rate cycle with Germany than in the 1980s, for some even synchronised. This supports the correlation evidence of increasing comovement. However, several aspects remain unclear - for instance, France seems to deteriorate in its financial integration with Germany although these two countries are commonly seen as very integrated. The idiosyncratic impact of German unification in the early 1990s may come into play here but our analysis is not able to isolate such effects. It is remarkable to what a large degree the UK seems to be financially integrated with Germany. Based on this result, the UK may reap a large gain from joing the euro even in the presence of non-synchronised business cycles. For the NMS, it seems that financial integration is still under development but prospects appear good that further economic integration would stimulate financial interactions, suggested by the more favourable results for the EU-15.

Table 4.11: Interest rate codependenc results, EU-15, 1990-1998

Country	rank	Common features	Codependence Order 1	Order 2	Order 3
Austria	$m = 1$	1.87	8.42	14.03**	9.97*
	$m = 2$	21.38**	33.56***	37.59***	35.88***
Belgium	$m = 1$	20.43***	0.94	3.49	5.08
	$m = 4$	47.89***	10.16	23.39***	27.80***
Denmark	$m = 1$	17.40***	3.68	4.99	2.58
	$m = 2$	60.60***	22.76***	15.16*	11.03
Finland	$m = 1$	12.27***	2.08	3.82	4.57
	$m = 2$	32.76***	7.19	14.39*	15.42*
France	$m = 1$	13.48***	3.21	0.51	5.22
	$m = 2$	47.13***	14.83*	7.40	13.93
Greece	$m = 1$	0.56	3.26	11.26**	1.64
	$m = 2$	26.48***	9.144	29.36***	9.10
Ireland	$m = 1$	1.76	0.02	0.01	3.62*
	$m = 2$	29.72***	5.77	8.22*	7.23
Italy	$m = 1$	1.94	0.41	0.02	0.28
	$m = 2$	23.52***	4.72	1.24	4.45
Luxembourg	$m = 1$	4.52	3.02	18.61***	3.86
	$m = 2$	42.25***	17.95*	39.07***	21.02**
Netherlands	$m = 1$	7.76*	2.61	4.78	5.89
	$m = 2$	38.03***	11.94	18.39**	14.82*
Portugal	$m = 1$	0.13	0.01	2.53	2.99*
	$m = 2$	6.62	0.93	8.37*	8.32*
Spain	$m = 1$	1.27	4.32**	0.15	3.76*
	$m = 2$	7.38	15.05***	5.65	10.26**
Sweden	$m = 1$	25.17***	5.84	21.62***	5.64
	$m = 2$	53.33***	19.09*	45.27***	21.53**
UK	$m = 1$	4.73	9.14	14.23**	8.79
	$m = 2$	33.95***	26.91**	30.61***	17.64
UK vs. US	$m = 1$	26.60***	3.33	8.63**	7.88**
	$m = 2$	59.66***	16.22**	26.08***	16.23**

*Note: Codependence results of real interest rates of each country vis-à-vis Germany. Rejection of the null hypothesis of common feature/codependence vectors at the 1 percent level is indicated by "***", the 5 percent level is marked with "**", the 10 percent level with "*". If we find the combination of accepting one vector ($m = 1$) and rejecting a second vector ($m = 2$), we conlude the existence of $n - m = 2 - 1 = 1$ common cycle.*

4.3 Summary and conclusion

This chapter analysed the role of risk sharing and financial integration in the context of the OCA theory and the Mundell II framework. According to Mundell II, countries with less sychronised business cycles benefit most from the risk-sharing properties in a financially integrated currency union. Since a common currency removes exchange rate fluctuations and cross-country risk premia, portfolio diversification is expected to deepen across the currency union and serves as a consumption insurance mechanism because it decouples consumption from national production patterns. This benefit of common currencies has often been overlooked while the cost of currency union membership due to the loss of individual monetary policy has been highlighted alone.

In the present chapter, we investigated the degrees of risk sharing and financial integration in the enlarged EU to explore the case for Mundell II mechanisms for euro area enlargement. In particular, we analysed consumption and real interest rate comovement of the eight Central and Eastern European new member states, each in relation to the aggregate euro area which they are supposed to join in due course. For comparison, we investigated the member countries of the "old" EU-15 in relation to the euro area or, in the case of financial integration, relative to the pre-EMU benchmark Germany. Our main findings are as follows.

Regarding risk sharing, we compare cross-country comovement of consumption with that of GDP. Methodologically, we first look at simple correlation coefficients before we move on to the more sophisticated time-series technique of codependence. From a theoretical point of view, risk sharing would be manifested in internationally diversified consumption patterns so that consumption across countries should be relatively independent of income and hence more highly correlated than GDP. Our results indicate that consumption correlations with the euro area are lower than GDP correlations for most countries under investigation. While this result is, at first glance, in line with the consumption correlation puzzle, we find a number of insightful details. For the NMS, correlations are at far lower levels than for the EU-15 countries while Slovenia stands out with relatively high levels of consumption and GDP correlation. Also, Lithuania and Slovenia display synchronised common GDP cycles as identified by the codependence analysis. Furthermore, rolling correlation windows indicate increasing correlation coefficients for most countries over time. We note that GDP correlations exhibit steeper increases than consumption correlations.

Turning to financial integration, we investigate real interest rate comovement. In addition to the correlation measures, we analyse the variability of bilateral differentials and the dispersion of interest rates across countries over time. We also resort to the codependence framework. While we again look at the eight NMS vis-à-vis the euro area from 1995 through 2005, we consider the EU-15 countries against the pre-EMU benchmark Germany and consider the 1980-1998 period.

We acknowledge a somewhat unclear stationarity situation with interest rates. Theoretically, we would expect interest rates to be associated with consumption growth and hence stationary. However, the unit root hypothesis cannot be rejected in many cases although it is hard to imagine interest rates to be literally non-stationary. High persistence or structural breaks may account for the unit root results. Given this ambiguity, we analyse interest rates both in levels and in differences.

NMS evidence proves mixed. While the correlation analysis delivers partly conflicting results, the codependence exercise suggests common features for Hungary, Poland and Slovakia. In other words, real interest rates of these countries seem to exhibit synchronised common cycles with the euro area. When looking at rolling correlation windows, nearly all NMS seem to increase in their interest rate comovement with the euro area over time. Also, the variability of bilateral interest rate differentials decreases markedly from the mid-1990s until 2005.

For the EU-15 countries, we find more unambiguous evidence of financial integration.

From the 1980s to the 1990s, interest rate correlations with Germany shot up to high levels. Austria, Belgium, Luxembourg and the Netherlands seem to form a core group whose interest rate correlation with Germany fluctuated on high levels whereas the correlation coefficients of most remaining countries started at low levels in the 1980s and experienced stark increases until the late 1990s. The core group of financially integrated countries is confirmed by the variability analysis of bilateral interest rate differentials. Furthermore, the dispersion measure, also known as sigma convergence, indicates a clear downward trend which is even more pronounced when excluding idiosyncratic Greece. Finally, we conducted separate codependence tests for the 1980s and 1990s and found increasing degrees of interest rate comovement between the EU-15 countries and Germany. While only a few countries qualified for synchronised common interest rate cylces during the 1980s, we find common feature evidence during the 1990s for Austria, Greece, Ireland, Italy, the Netherlands and the UK. A number of borderline cases add to this evidence. It is interesting to note that the UK displays high levels of financial integration throughout our analysed indicators.

Taken together, we draw a threefold conclusion from our results. First, we confirm the consumption correlation puzzle established by most empirical literature. Consumption correlations remain below output correlation for most considered countries which contradicts the theoretical proposition. One major reason behind this may be the relatively low degree of financial integration. We confirm this idea at least in the case of the NMS which, to date, seem to be characterised by both little risk sharing and limited financial integration with the euro area.

Second, even though GDP correlation still exceeds consumption correlation for the EU-15 countries, they are both on much higher levels and with a smaller differential than for the NMS. Also, financial integration has improved markedly for the EU-15 countries in the run-up to EMU. Given that these countries have shared a long history of economic integration, we may suspect a similar development for the NMS as integration with the euro area proceeds.

Third, we find that both consumption and GDP correlations increase over time, with the latter more strongly than the former. Also, interest rate correlations tend to rise for most countries over time. Although we did not conduct any causal analysis within the scope of this chapter, these observations may support the hypothesis of Imbs (2006). He analyses a large set of countries and finds that financial integration does not only improve risk-sharing opportunities in the form of cross-country consumption correlation but also boosts, to an even larger extent, business cycles synchronisation across countries. Hence, he argues, the consumption correlation puzzle may not stem from too little risk sharing. Accordingly, we see a widening gap between consumption and GDP correlation not because of low degrees of risk sharing but simply because GDP correlations increase even faster than those of consumption. From our results, we can at least confirm that GDP correlations do indeed increase faster than consumption correlations, and the rising levels of financial integration are not unlikely to play a central part in that.

These propositions hint at further need for research. To shed more light on the dynamics of risk sharing, financial integration and business cycle synchronisation, a comprehensive econometric framework would be desirable. Also, to respond to the prevailing policy question of euro area enlargement and its effects on the new member states and on the euro area, we would welcome more research on these countries. If, as Mundell II argues, countries with relatively asynchronous business cycles benefit most from the risk-sharing opportunities in a financially integrated currency union, the NMS may have far more to gain from euro adoption than previously assumed. This logic applies even more if the euro delivers the enhanced degree of financial integration that some studies suggest.

Chapter 5

Conclusion

The introduction of the single European currency in 1999 was undoubtedly one of the most remarkable innovations in international finance during the last decades. With the enlargement of the European Union by ten states in 2004, and more accession countries in negotiation, the enlargement of the euro area to the formerly communist new EU member states is the next major challenge ahead.

This dissertation addressed a number of fundamental policy questions related to monetary integration in the enlarged EU. Our arguments unfolded along the lines of optimum currency area theory and its three major strands: the classical OCA criteria (Mundell I), the endogeneity of OCAs and the role of risk sharing (Mundell II). We focused our empirical strategy on business cycle synchronisation which has been regarded as a "meta-property" in operationalising the OCA framework. In the following, we give a short overview over the main steps of the analysis and their results before we highlight some policy implications and the need for future research.

In Chapter 2, we tested for common trends and cycles of a number of NMS and accession countries vis-à-vis the euro area. According to the traditional OCA framework, Mundell I, the best-suited candidates for currency union are characterised by a large degree of business cycle synchronisation so that renouncing individual monetary and exchange rate policy would not give rise to major economic costs. Hence, the NMS should condition their adoption of the euro to their business cycles being sufficiently synchronised with that of the euro area.

We investigated this question by applying the integrated cointegration/codependence approach by Engle and Kozicki (1993) and Vahid and Engle (1997) to a decade of European output data. We found that, regarding long-run convergence, the NMS are still in a catching-up process and have not yet reached a steady-state equilibrium with the euro area. Turning to short-run cycle comovement, our results indicate that only Slovenia has a synchronised common business cylce with the euro area. Hungary, Slovakia, Estonia and, as a borderline case, the Czech Republic, show signs of common but not synchronised cycles which points at an intermediate degree of comovement. Thus, our analysis supports the Slovenian euro adoption which is scheduled for 2007 but suggests caution regarding the remaining countries. According to the Mundell I framework, they may incur major costs by renouncing individual monetary and exchange rate policies before having reached a sufficient degree of business cycle synchronisation.

In Chapter 3, we added the endogeneity dimension of OCAs. According to Frankel and Rose (1998) and Engel and Rose (2002), the adoption of a common currency per se may unfold synchronisation dynamics which lead to endogenous trade increase and cycle synchronisation. Hence, even an *ex ante* non-optimal currency area like the EU may turn out to be optimal *ex post*. However, it is still too early to empirically identify a reliable endogenous effect of the euro on the EU economies. Therefore, we followed the approach of Frankel and Rose (1998) by asking which factors are significantly associated with business cycle synchronisation across euro area countries. A positive relation between trade and cycle comovement would then be interpreted as an indication of OCA endogeneity. To test the robustness of the potential determinants, we applied the extreme-bounds analysis by Leamer (1983).

We found that, indeed, bilateral trade has been a robust, positive determinant of business cycle synchronisation across euro area countries over the past 25 years. As we split up our sample to learn more about time-varying effects, our results show that the explanatory power of the trade effect seems to be driven mainly by the earlier sub-period, 1980-1996. Since 1997, the differences in trade structure emerge as a robust determinant of cycle synchronisation. In other words, the degree of intra-industry trade plays an increasing role in business cycle comovement. Given our descriptive finding of a rising degree of intra-industry trade, we interpret our results as a positive indication for both the existing euro area and the prospective entrants.

In addition to the trade-related determinants, we included several policy and structural indicators into our analysis. We found that fiscal policy similarity has had a positive effect on cycle synchronisation up to the EMU preparation phase. Since 1997, we found monetary policy similarity as proxied by real interest rate differentials to emerge as a robust determinant. Furthermore, similar industrial sector size, stock market comovement and similar competitiveness seem to have good explanatory power. In contrast, nominal exchange rate variation, bilateral bank capital flows and differences in labour market flexibility did not turn out as robust.

Chapter 4 addressed a strand of OCA theory which has attracted a lot of attention recently: the role of risk sharing and financial integration in currency unions, known as Mundell II. In a financially-integrated currency union, it is argued, countries with little business cycle synchronisation may benefit even more from adopting the common currency. This benefit is due to new consumption risk sharing opportunities because national consumption patterns can be diversified across the union and are thus less contingent on home output. For the new EU member states, this idea implies that for the countries with asynchronous cycles, the euro would be more, and not less attractive.

To explore the past degree and future potential of risk sharing and financial integration in the enlarged EU, we investigated consumption and real interest rate comovement between the NMS and the euro area. For comparison, we applied the same measures to the "old" EU member states. Methodologically, we resorted to cross-country correlations and the common features/codependence technique. In addition, we employed various variability and dispersion measures. We found that consumption comovement between the NMS and the euro area tends to be below comparable measures of output synchronisation. This result, which is in line with the consumption correlation puzzle, may be due to relatively low degrees of financial integration. For the EU-15 countries, however, consumption and output correlation levels are much higher and the difference, though still often negative, tends to be more narrow. Also, financial integration has increased markedly as the EU-15 countries prepared for EMU. In view of the long common history of economic integration among the EU-15, we may expect similar effects to materialise as the integration of the NMS into the enlarged EU proceeds. Finally, we observed that both consumption and output correlations tend to increase over time, alongside increasing financial integration. Notably, output comovement increased at a faster rate than that of consumption. This experience seems consistent with the hypothesis of Imbs (2006) who suggests that financial integration does not only facilitate risk sharing but also, and even more, boosts business cycle synchronisation. Hence, the consumption correlation puzzle may not stem from too little risk sharing but originates in the often neglected, strong effect on output comovement. In consequence, the benefit that the NMS may derive from early euro adoption, may so far have been underestimated.

In the light of these results, we see a number of interesting policy implications and the need for further research. First, it becomes clear that transition is not yet over in the NMS. Although remarkable progress has been made, business cycles as well as consumption and interest rates are far from being in line with the euro area. Hence, further reform and integration efforts are required that go beyond the scope of the analysis in this dissertation. More research would improve our understanding of how, for example, structural reforms in the banking sector and on the labour markets may improve the functioning of adjustment channels to foster convergence and cycle sychronisation.

The euro itself may, however, turn out to be one means of aligning business cycles. Instead of postponing euro area accession until full synchronisation is achieved, the NMS may want to exploit the endogenous effects of currency union. EU accession pushed trade integration with the EU-15 and monetary union is likely to increase bilateral trade further. Given our results on OCA endogeneity, we may expect increased trade to translate into more closely associated business cycles as the NMS adopt the euro.

One question is, however, whether increased trade will materialise in similar or different sectors. Our evidence for the euro area suggests that the trade structure has become increasingly important in determining business cycle synchronisation.[1] Several studies suggest that trade between the NMS and the EU-15 is increasingly intra-industry[2] which would hint at good prospects for endogenous cycle synchronsiation. However, this trade seems to occur mostly as vertical intra-industry trade. i.e. trade in similar sectors but in different qualities, and driven by wage differentials and foreign direct investment flows. More research is needed to assess whether this kind of specialisation on low-wage production in the same sector leads to more asymmetric cycles, whether the observed foreign ownership patterns tend to bind cycles together, and which effect would dominate.

The role of risk-sharing benfits adds an interesting perspective on euro area enlargement. According to this logic, the NMS with least cycle synchronisation have most to benefit from joining the euro area quickly. Our results confirm that, should the mechanism work reasonably well, most NMS would have substantial gains to reap. Given that this view has not yet received much attention in empirical research and is appears still somewhat vague, we would welcome in-depth studies to substantiate and quantify the potential gains for the NMS. It would be of particular interest to learn more about domestic financial markets and the degree of private stock and equity ownership in the NMS that would facilitate cross-country risk sharing. In addition, the Mundell II view on risk sharing benefits refers mostly to prospective currency union entrants and their

[1] Fidrmuc (2003) reaches a simliar conclusion for the CEECs.

[2] See Caetano and Galego (2006), Gabrisch and Segnana (2003).

gains from joining a common currency. We do not yet know much about the effect on an exisiting currency union, for instance the euro area. How would an acceding country with asynchronous business cycles affect a relatively homogeneous existing monetary union? If it was predominantly the new country that benefits, would the existing union therefore prefer only small countries of relatively little weight to join, as opposed to larger asynchronous economies which may cause more concern for the union as a whole? These questions have remained largely open and are the object of future research.

Although euro area enlargement is ultimately a political question, policy decision makers are in continuous need of answers from empirical economics. With the contribution of this dissertation on business cycle synchronisation, we hope to substantiate economic policy decision making in the enlarged European Union.

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Appendix A

Table A.1: Variables and data sources

Variable Name	Description	Data source
COR	Correlation coefficient of business cycles	European Commission, Ameco Database; own calculations
BTT	Bilateral trade, scaled by total trade	IMF, Direction of Trade Statistics; Ameco; own calculations
BTY	Bilateral trade, scaled by GDP	IMF, Direction of Trade Statistics; Ameco; own calculations
TTY	Total trade of both countries, scaled by GDP	IMF, Direction of Trade Statistics; Ameco; own calculations
ECOPAT	Sum of relative sector shares in total value added	OECD National Accounts Database; own calculations
CD_IND	*Relative shares of industry*	
CD_CNT	*Relative shares of construction*	
CD_FIN	*Relative shares of financial intermediation*	
CD_TRA	*Relative shares of wholesale & retail trade*	
TRADEPAT	Sum of relative sector shares in bilateral exports	NBER World Trade Flows Database, see Feenstra and Lipsey (2005) ; own calculations
CD_FUEL	*Relative shares of mineral fuels*	
CD_MACH	*Relative shares of machinery and transport equipment*	
CD_MANU	*Relative shares of other manufacturing products*	
CD_CHEM	*Relative shares of chemicals*	
BFA, BFL	Bilateral bank flows (assets, liabilities)	BIS, International Locational Banking Statistics, see Papaioannou (2005); own calculations
TOTMKDIFF	Bilateral difference between overall stock market indices	Thomson Datastream ; own calculations
CYSERDIFF	Bilateral difference between stock market indices for cyclical services	Thomson Datastream ; own calculations
IRSCDIFF	Bilateral short-run interest rate differential minus inflation measured by the private consumption deflator	European Commission, Ameco Database ; own calculations
NCIDIFF	Bilateral differences between real effective exchange rates (HICP-deflated)	Calculation
SD_NERE	Bilateral exchange rate variation, defined as the standard deviation of the nominal exchange rates	Bank for International Settlements; own calculations
DEFDIFF	Bilateral difference in fiscal budget deficits	European Commission, Ameco Database; own calculations
TUDDIFF	Bilateral difference in trade union density, defined as the share of organised workers	OECD Olisnet Labour Market Statistics; own calculations
EPADIFF	Bilateral difference in the averaged OECD employment protection indices	OECD Olisnet Labour Market Statistics; own calculations
GEODIST	Geographical distance between national capitals (Bonn for Germany)	International Trade Database, Macalester University; own calculations
POPDIFF	Bilateral difference in national population, scaled by population	European Commission, Ameco Database; own calculations

Note: The fully-detailed description of variables can be found in the text of the paper.

Figure A.1: Business cycle correlation over time

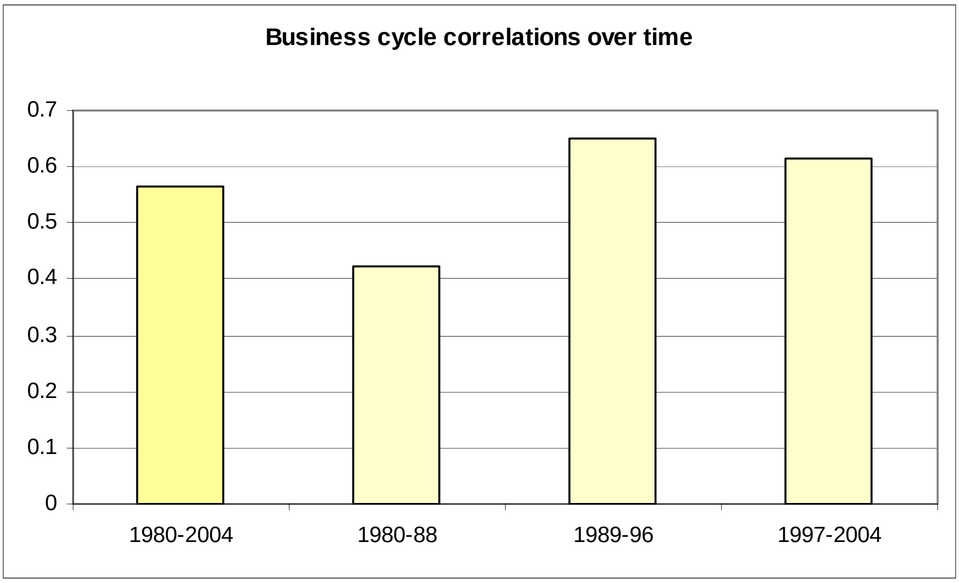


Figure A.2: Business cycle correlation coefficients, 1980 – 1988

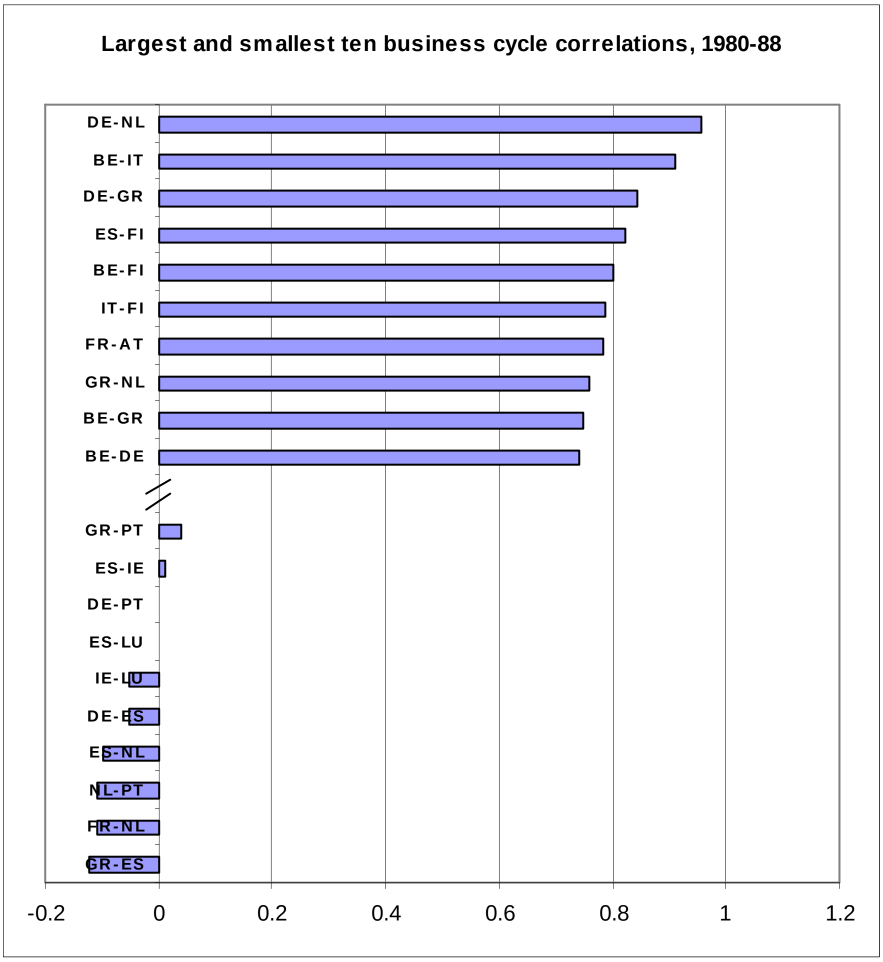


Figure A.3: Business cycle correlation coefficients, 1989 – 1996

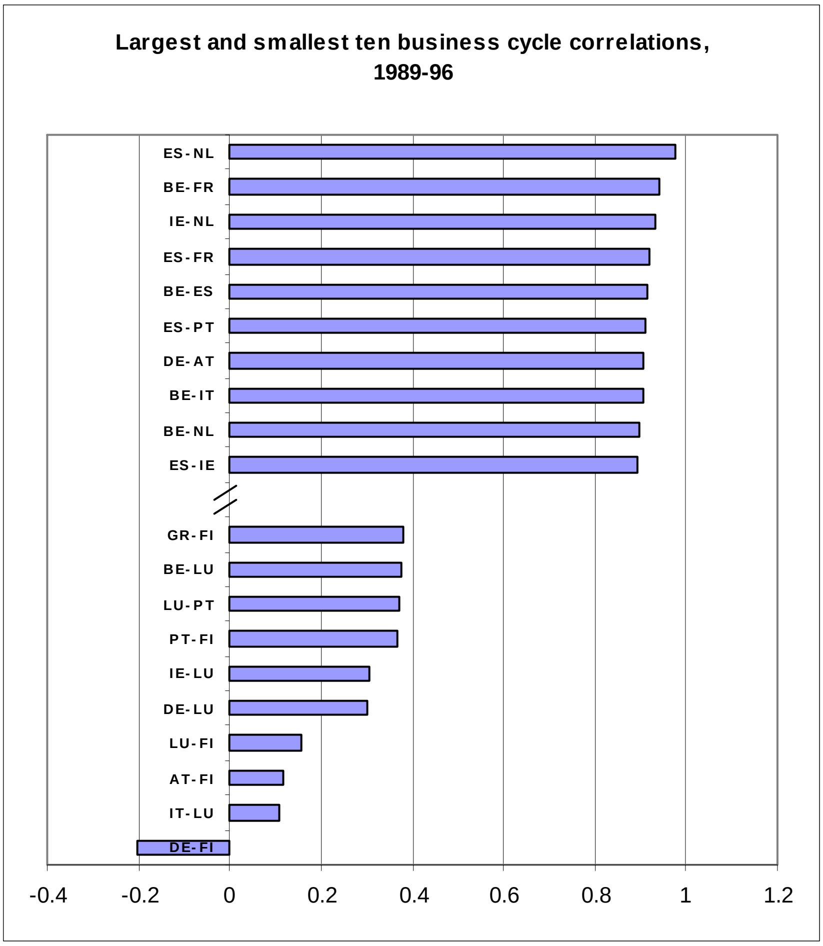
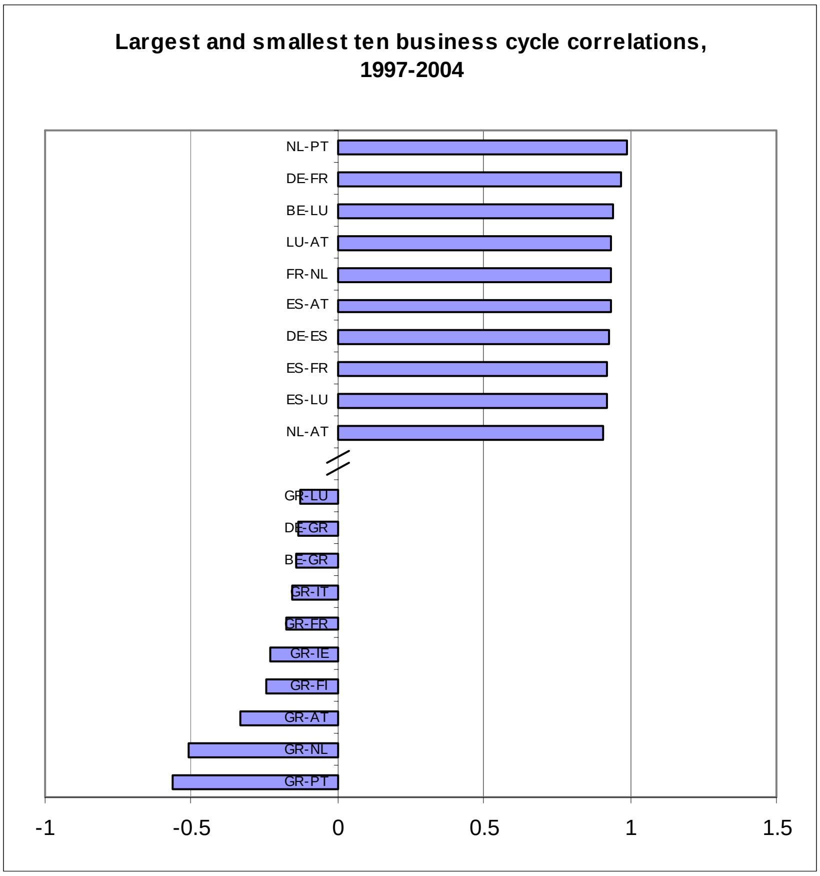


Figure A.4: Business cycle correlation coefficients, 1997 – 2004



Appendix B: EBA estimates

- The results of the extreme-bounds analysis are reported in tables B. 1 to B. 12. For a sample size of 60 (the actual sample has 66 observations), the significance levels for the t-statistics are: 1.671 for the 10% level ; 2.000 for the 5% level ; 2.660 for the 1% level.

- The t-statistics reported in the tables include a Newey-West correction for heteroskedasticity and autocorrelation in the residuals.

- We consider as "quasi-robust" the variables whose coefficients for all equations were significant and of the expected sign, but for which one of the bounds took the wrong sign while remaining around 0, with an absolute value of less than 5% of the relevant coefficient.

Table B.1: Ratio of bilateral trade to total trade (BTT) *W/O geographical distance before 1997*

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Robust	Bivariate		2.065	0.524	3.94	0.18		100%
	High	**3.112**	2.055	0.528	3.89	0.17	TUDDIFF	
	Low	**0.123**	0.956	0.416	2.30	0.40	TOTMKDIFF, NCIDIFF, DEFDIFF	
1980-1996								
Robust	Bivariate		1.872	0.582	3.22	0.12		100%
	High	**3.349**	2.082	0.634	3.29	0.11	SD_NERE, TUDDIFF	
	Low	**0.301**	1.369	0.534	2.56	0.13	TOTMKDIFF, NCIDIFF, TUDDIFF	
1997-2004								
Fragile	Bivariate		4.092	1.456	2.81	0.10		46.3%
	High	**7.269**	4.121	1.574	2.62	0.09	TUDDIFF	
	Low	**-2.660**	-0.830	0.915	-0.91	0.32	TOTMKDIFF, DEFDIFF, GEODIST	

Table B.2: Ratio of bilateral trade to GDP (BTY) *W/O geographical distance before 1997*

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Significant coefficients (%)
1980-2004								
Robust	Bivariate		3.216	1.108	2.90	0.15		100%
	High	**5.393**	3.204	1.095	2.93	0.15	TUDDIFF	
	Low	**0.123**	1.524	0.700	2.18	0.42	IRSCDIFF, NCIDIFF, DEFDIFF	
1980-1996								
Robust	Bivariate		3.111	0.896	3.47	0.11		100%
	High	**5.480**	3.405	1.037	3.28	0.10	SD_NERE	
	Low	**0.682**	2.268	0.793	2.86	0.16	TOTMKDIFF, IRSCDIFF,	
1997-2004								
Fragile	Bivariate		5.893	2.845	2.07	0.09		26.8%
	High	**11.714**	5.895	2.909	2.03	0.07	TUDDIFF	
	Low	**-5.080**	-2.534	1.273	-1.99	0.35	DEFDIFF, TUDDIFF, GEODIST	

Table B.3: Trade specialisation patterns (TRADEPAT)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Fragile	Bivariate		-0.433	0.140	-3.10	0.19		100%
	High	**0.032**	-0.169	0.101	-1.68	0.38	IRSCDIFF, NCIDIFF, SD_NERE	
	Low	**-0.715**	-0.437	0.139	-3.14	0.20	TUDDIFF	
				1980-1996				
Fragile	Bivariate		-0.237	0.157	-1.50	0.04		n.a.
	High	**0.219**	-0.074	0.146	-0.51	0.10	NCIDIFF, GEODIST	
	Low	**-0.586**	-0.246	0.170	-1.45	0.02	SD_NERE	
				1997-2004				
Robust	Bivariate		-1.233	0.293	-4.21	0.35		100%
	High	**-0.022**	-0.469	0.224	-2.10	0.58	IRSCDIFF, DEFDIFF, GEODIST	
	Low	**-2.055**	-1.491	0.282	-5.28	0.40	NCIDIFF, TUDDIFF	

Table B.3a: Trade specialisation in fuels (CD_FUELS)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Fragile	Bivariate		-0.348	0.629	-0.55	-0.01		n.a.
	High	**0.813**	-0.084	0.449	-0.19	0.07	TOTMKDIFF, IRSCDIFF,	
	Low	**-1.555**	-0.655	0.450	-1.46	0.35	DEFDIFF, TUDDIFF, GEODIST	
				1980-1996				
Fragile	Bivariate		0.197	0.628	0.31	-0.02		n.a.
	High	**1.503**	0.245	0.629	0.39	-0.03	SD_NERE	
	Low	**-1.556**	-0.240	0.658	-0.36	0.11	NCIDIFF, SD_NERE, GEODIST	
				1997-2004				
Fragile	Bivariate		-4.943	1.928	-2.56	0.22		92.7%
	High	**0.936**	-0.692	0.814	-0.85	0.76	TOTMKDIFF, IRSCDIFF,	
	Low	**-8.993**	-5.197	1.898	-2.74	0.20	NCIDIFF, TUDDIFF	

Table B.3b: Trade specialisation in machinery and transport equipment (CD_MACH)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.720	0.289	-2.50	0.11		100%
	High	**0.061**	-0.446	0.253	-1.76	0.40	IRSCDIFF, NCIDIFF, SD_NERE	
	Low	**-1.516**	-0.956	0.280	-3.42	0.25	TOTMKDIFF, TUDDIFF	
1980-1996								
Fragile	Bivariate		-0.276	0.337	-0.82	-0.00		n.a.
	High	**0.457**	-0.119	0.288	-0.41	0.09	NCIDIFF	
	Low	**-1.383**	-0.514	0.434	-1.18	0.09	TOTMKDIFF, SD_NERE,	
1997-2004								
Robust	Bivariate		-3.590	0.536	-6.70	0.60		100%
	High	**-0.566**	-1.427	0.431	-3.31	0.78	IRSCDIFF, NCIDIFF, DEFDIFF	
	Low	**-4.680**	-3.680	0.500	-7.36	0.61	TUDDIFF	

Table B.3c: Trade specialisation in other manufacturing (CD_MANU)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.560	0.266	-2.10	0.03		23.8%
	High	**0.707**	-0.062	0.385	-0.16	0.35	TOTMKDIFF, IRSCDIFF, NCIDIFF	
	Low	**-1.376**	-0.808	0.284	-2.84	0.16	IRSCDIFF, SD_NERE	
1980-1996								
Fragile	Bivariate		-0.558	0.319	-1.75	0.03		1.6%
	High	**0.493**	-0.232	0.362	-0.64	0.12	NCIDIFF, SD_NERE, GEODIST	
	Low	**-1.364**	-0.645	0.360	-1.79	0.03	SD_NERE	
1997-2004								
Fragile	Bivariate		0.427	0.725	0.59	-0.01		n.a.
	High	**4.145**	2.098	1.023	2.05	0.26	TOTMKDIFF, DEFDIFF	
	Low	**-2.708**	-1.803	0.453	-3.98	0.74	IRSCDIFF, NCIDIFF	

Table B.3d: Trade specialisation in chemicals (CD_CHEM)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Fragile	Bivariate		-0.285	0.481	-0.59	-0.01		
	High	**1.321**	0.230	0.546	0.42	0.35	TOTMKDIFF, IRSCDIFF,	n.a.
	Low	**-1.278**	-0.510	0.384	-1.33	0.22	DEFDIFF, TUDDIFF	
				1980-1996				
Fragile	Bivariate		0.265	0.731	0.36	-0.02		
	High	**2.235**	0.757	0.739	1.02	0.09	TOTMKDIFF, NCIDIFF,	n.a.
	Low	**-1.498**	0.099	0.799	0.12	-0.01	DEFDIFF, TUDDIFF	
				1997-2004				
Fragile	Bivariate		0.333	0.511	0.65	-0.01		
	High	**4.002**	2.161	0.921	2.35	0.29	TOTMKDIFF, NCIDIFF, DEFDIFF	n.a.
	Low	**-2.336**	-1.616	0.360	-4.49	0.75	IRSCDIFF, NCIDIFF	

Table B.4: Economic specialisation patterns (ECOPAT)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Fragile	Bivariate		-0.499	0.191	-2.61	0.05		
	High	**0.274**	-0.145	0.209	-0.69	0.26	TOTMKDIFF, DEFDIFF,	81.0%
	Low	**-0.980**	-0.604	0.188	-3.22	0.07	TUDDIFF	
				1980-1996				
Fragile	Bivariate		-0.612	0.305	-2.01	0.05		
	High	**0.194**	-0.412	0.303	-1.36	0.13	TOTMKDIFF, NCIDIFF, DEFDIFF	77.8%
	Low	**-1.429**	-0.902	0.264	-3.42	0.16	NCIDIFF, SD_NERE, GEODIST	
				1997-2004				
Fragile	Bivariate		-0.473	0.419	-1.13	0.00		
	High	**1.058**	0.370	0.344	1.07	0.53	TOTMKDIFF, IRSCDIFF,	n.a.
	Low	**-1.284**	-0.497	0.393	-1.27	-0.01	TUDDIFF	

Table B.4a: Economic specialisation in industry (CD_IND)
IRSDIFF: differential between short-term interest rates deflated by the GDP deflators

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Robust	Bivariate		-1.979	0.601	-3.29	0.11		100%
	High	**-0.265**	-1.156	0.445	-2.60	0.44	IRSDIFF, NCIDIFF, DEFDIFF	
	Low	**-3.242**	-2.148	0.547	-3.93	0.13	TUDDIFF	
				1980-1996				
Quasi-robust	Bivariate		-2.048	0.903	-2.27	0.06		100%
	High	**0.126**	-1.482	0.804	-1.84	0.11	TOTMKDIFF, NCIDIFF	
	Low	**-4.462**	-2.692	0.885	-3.04	0.11	IRSDIFF, SD_NERE, GEODIST	
				1997-2004				
Fragile	Bivariate		-1.717	1.088	-1.58	0.03		n.a.
	High	**2.725**	0.519	1.103	0.47	0.17	TOTMKDIFF, IRSDIFF, DEFDIFF	
	Low	**-6.016**	-3.279	1.369	-2.40	0.22	IRSDIFF, NCIDIFF, TUDDIFF	

Table B.4b: Economic specialisation in construction (CD_CNT)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Fragile	Bivariate		5.426	2.530	2.14	0.03		77.8%
	High	**10.862**	6.728	2.067	3.25	0.29	IRSCDIFF, SD_NERE, GEODIST	
	Low	**-1.522**	2.636	2.079	1.27	0.36	TOTMKDIFF, NCIDIFF, DEFDIFF	
				1980-1996				
Robust	Bivariate		11.680	3.584	3.26	0.08		100%
	High	**20.136**	12.476	3.830	3.26	0.10	SD_NERE, TUDDIFF, GEODIST	
	Low	**1.108**	8.986	3.939	2.28	0.15	TOTMKDIFF, NCIDIFF, DEFDIFF	
				1997-2004				
Fragile	Bivariate		-0.953	4.160	-0.23	-0.01		n.a.
	High	**9.161**	4.474	2.344	1.91	0.71	IRSCDIFF, NCIDIFF, GEODIST	
	Low	**-10.915**	-2.919	3.998	-0.73	0.12	TOTMKDIFF, NCIDIFF, DEFDIFF	

Table B.4c: Economic specialisation in wholesale and retail trade (CD_TRA)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.342	0.887	-0.39	-0.01		n.a.
	High	**2.954**	1.180	0.887	1.33	0.24	IRSCDIFF, SD_NERE, GEODIST	
	Low	**-2.267**	-0.621	0.823	-0.75	0.20	DEFDIFF, TUDDIFF	
1980-1996								
Fragile	Bivariate		0.543	0.748	0.73	-0.01		n.a.
	High	**2.676**	1.015	0.831	1.22	-0.00	TUDDIFF, GEODIST	
	Low	**-1.543**	0.103	0.823	0.12	0.10	NCIDIFF, DEFDIFF, SD_NERE	
1997-2004								
Fragile	Bivariate		-5.573	2.060	-2.70	0.12		70.7%
	High	**1.324**	-1.069	1.196	-0.89	0.57	TOTMKDIFF, IRSCDIFF,	
	Low	**-9.594**	-5.742	1.926	-2.98	0.10	NCIDIFF, TUDDIFF	

Table B.4d: Economic specialisation in financial intermediation (CD_FIN)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.450	0.396	-1.13	0.00		n.a.
	High	**0.982**	0.047	0.468	0.10	0.20	TOTMKDIFF, DEFDIFF	
	Low	**-1.429**	-0.901	0.264	-3.41	0.41	IRSCDIFF, NCIDIFF, GEODIST	
1980-1996								
Quasi-Robust	Bivariate		-1.464	0.482	-3.03	0.10		100%
	High	**0.021**	-1.129	0.575	-1.96	0.15	TOTMKDIFF, DEFDIFF	
	Low	**-2.631**	-1.732	0.449	-3.85	0.14	TUDDIFF, GEODIST	
1997-2004								
Fragile	Bivariate		1.045	0.593	1.76	0.01		51.2%
	High	**3.858**	2.062	0.898	2.30	0.22	TOTMKDIFF, DEFDIFF,	
	Low	**-0.439**	0.235	0.337	0.70	0.57	IRSCDIFF, TUDDIFF, GEODIST	

Table B.5: Log of bilateral flows of bank assets trade (LBFA)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		0.038	0.011	3.39	0.16		69.8%
	High	**0.060**	0.039	0.010	3.87	0.13	IRSCDIFF, SD_NERE	
	Low	**-0.023**	0.005	0.014	0.36	0.34	IRSCDIFF, NCIDIFF, DEFDIFF	
1980-1996								
Fragile	Bivariate		0.025	0.019	1.33	0.02		n.a.
	High	**0.088**	0.031	0.028	1.10	-0.03	SD_NERE, TUDDIFF, GEODIST	
	Low	**-0.101**	-0.042	0.030	-1.40	0.21	TOTMKDIFF, NCIDIFF,	
1997-2004								
Fragile	Bivariate		0.025	0.010	2.50	0.12		22.0%
	High	**0.050**	0.028	0.011	2.52	0.12	IRSCDIFF, NCIDIFF	
	Low	**-0.020**	0.000	0.010	0.01	0.31	IRSCDIFF, DEFDIFF, GEODIST	

Table B.6: Real short-term interest rate differential (IRSCDIFF)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.049	0.028	-1.73	0.03		7.3%
	High	**0.175**	0.109	0.033	3.27	0.34	TOTMKDIFF, NCIDIFF,	
	Low	**-0.107**	-0.050	0.028	-1.77	0.03	TUDDIFF	
1980-1996								
Fragile	Bivariate		-0.008	0.018	-0.45	-0.01		n.a.
	High	**0.115**	0.058	0.028	2.05	0.06	NCIDIFF, TUDDIFF	
	Low	**-0.077**	-0.022	0.027	-0.80	0.05	DEFDIFF, SD_NERE	
1997-2004								
Robust	Bivariate		-0.417	0.079	-5.28	0.50		100%
	High	**-0.177**	-0.328	0.076	-4.33	0.58	TOTMKDIFF, DEFDIFF,	
	Low	**-0.753**	-0.596	0.079	-7.59	0.69	NCIDIFF, TUDDIFF	

Table B.7: Nominal exchange rate volatility (SD_NERE)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.301	0.107	-2.80	0.10		36.5%
	High	**0.289**	0.048	0.120	0.40	0.28	NCIDIFF, TUDDIFF, GEODIST	
	Low	**-0.668**	-0.404	0.132	-3.07	0.16	TOTMKDIFF, IRSCDIFF,	
1980-1996								
Fragile	Bivariate		0.006	0.091	0.07	-0.02		n.a.
	High	**0.115**	0.058	0.028	2.05	0.06	NCIDIFF, TUDDIFF	
	Low	**-0.077**	-0.022	0.027	-0.80	0.05	TOTMKDIFF, TUDDIFF	

Table B.8a: Fiscal deficit differentials (DEFDIFF)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Robust	Bivariate		-3.046	0.581	-5.24	0.21		100%
	High	**-0.794**	-1.859	0.532	-3.49	0.43	BTT, IRSCDIFF, NCIDIFF	
	Low	**-4.166**	-3.020	0.573	-5.27	0.20	TUDDIFF	
1980-1996								
Quasi-robust	Bivariate		-1.784	0.573	-3.11	0.07		100%
	High	**0.049**	-1.186	0.618	-1.92	0.13	TOTMKDIFF, IRSCDIFF,	
	Low	**-2.940**	-1.807	0.567	-3.19	0.03	IRSCDIFF, SD_NERE, TUDDIFF	
1997-2004								
Fragile	Bivariate		-7.801	2.056	-3.80	0.12		97.6%
	High	**0.776**	-2.490	1.633	-1.52	0.54	BTT, IRSCDIFF, TUDDIFF	
	Low	**-14.672**	-8.610	3.031	-2.84	0.11	NCIDIFF, TUDDIFF	

Table B.8b: Fiscal deficit differentials (DEFDIFF) *with a dummy for the Germany-Finland pair*

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Robust	Bivariate		-3.003	0.576	-5.22	0.39		100%
	High	**-0.900**	-1.930	0.515	-3.75	0.57	BTT, IRSCDIFF, NCIDIFF	
	Low	**-4.192**	-3.006	0.593	-5.07	0.38	TUDDIFF	
				1980-1996				
Robust	Bivariate		-1.934	0.571	-3.39	0.27		100%
	High	**-0.169**	-1.381	0.606	-2.28	0.33	TOTMKDIFF, IRSCDIFF,	
	Low	**-3.082**	-1.940	0.571	-3.40	0.24	IRSCDIFF, SD_NERE, TUDDIFF	
				1997-2004				
Fragile	Bivariate		-8.043	2.205	-3.65	0.11		97.6%
	High	**0.715**	-2.601	1.658	-1.57	0.53	BTT, IRSCDIFF, TUDDIFF	
	Low	**-14.842**	-8.710	3.066	-2.84	0.09	NCIDIFF, TUDDIFF	

Table B.9: Price competitiveness differentials (NCIDIFF) *W/O geographical distance before 1997*

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Robust	Bivariate		-2.214	0.461	-4.80	0.26		100%
	High	**-0.031**	-1.410	0.690	-2.04	0.38	BTT, SD_NERE, GEODIST	
	Low	**-4.777**	-3.435	0.671	-5.12	0.30	IRSCDIFF, TUDDIFF	
				1980-1996				
Fragile	Bivariate		-0.736	0.409	-1.80	0.04		53.7%
	High	**0.532**	-0.241	0.387	-0.62	0.14	BTT, DEFDIFF, TUDDIFF	
	Low	**-3.159**	-1.781	0.68	-2.58	0.60	IRSCDIFF, SD_NERE, TUDDIF	
				1997-2004				
Fragile	Bivariate		-1.139	3.038	-0.37	-0.01		n.a.
	High	**17.885**	13.791	2.047	6.74	0.70	TOTMKDIFF, IRSCDIFF	
	Low	**-6.979**	-1.190	2.894	-0.41	-0.03	TUDDIFF	

Table B.10a: Total stock market differential (TOTMKDIFF)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Fragile	Bivariate		-0.018	0.011	-1.69	0.0.04		n.a.
	High	**0.010**	-0.003	0.007	-0.47	0.29	BTT, IRSCDIFF, DEFDIFF	
	Low	**-0.037**	-0.021	0.008	-2.56	0.16	IRSCDIFF, SD_NERE, TUDDIFF	
				1980-1996				
Fragile	Bivariate		-0.031	0.011	-2.88	0.05		69.8%
	High	**0.010**	-0.015	0.012	-1.20	0.17	BTT, DEFDIFF, SD_NERE	
	Low	**-0.057**	-0.034	0.011	-2.97	0.03	SD_NERE, TUDDIFF	
				1997-2004				
Fragile	Bivariate		-0.036	0.034	-1.09	0.03		n.a.
	High	**0.035**	0.002	0.017	0.10	0.51	BTT, IRSCDIFF, SD_NERE	
	Low	**-0.108**	-0.038	0.035	-1.10	0.00	NCIDIFF, TUDDIFF	

Table B.10b: Stock market differential for cyclical services (CYSERDIFF)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
				1980-2004				
Robust	Bivariate		-0.008	0.002	-4.70	0.19		100%
	High	-0.001	-0.004	0.001	-2.78	0.40	BTT, DEFDIFF, GEODIST	
	Low	-0.012	-0.008	0.002	-4.97	0.21	TUDDIFF	
				1980-1996				
Fragile	Bivariate		-0.006	0.004	-1.45	0.00		n.a.
	High	**0.007**	0.001	0.003	0.38	0.14	BTT, NCIDIFF, DEFDIFF	
	Low	**-0.015**	-0.007	0.004	-2.02	0.08	IRSCDIFF, NCIDIFF, SD_NERE	
				1997-2004				
Robust	Bivariate		-0.023	0.004	-5.57	0.53		100%
	High	**-0.000**	-0.009	0.005	-2.03	0.76	IRSCDIFF, NCIDIFF, DEFDIFF	
	Low	**-0.032**	-0.023	0.004	-5.72	0.54	NCIDIFF, TUDDIFF	

Table B.11: Differential in trade union membership (TUDDIFF)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.122	0.171	-0.71	-0.01		n.a.
	High	**0.372**	0.077	0.148	0.52	0.34	IRSCDIFF, NCIDIFF, GEODIST	
	Low	**-0.646**	-0.323	0.162	-2.00	0.16	TOTMKDIFF, IRSCDIFF,	
1980-1996								
Fragile	Bivariate		-0.037	0.192	-0.19	-0.01		n.a.
	High	**0.583**	0.168	0.207	0.81	0.05	NCIDIFF, SD_NERE, GEODIST	
	Low	**-0.499**	-0.128	0.186	-0.69	0.03	TOTMKDIFF, IRSCDIFF	
1997-2004								
Fragile	Bivariate		-0.008	0.334	-0.02	-0.02		n.a.
	High	**1.282**	0.500	0.391	1.28	0.34	NCIDIFF, DEFDIFF, GEODIST	
	Low	**-0.783**	-0.434	0.175	-2.48	0.52	TOTMKDIFF, IRSCDIFF	

Table B.12: Geographical distance (GEODIST)

Result	Estimation	Bounds	Coefficient	Stdd error	T Statistics	R^2 adj.	Z control variables	Percentage of significant coefficients
1980-2004								
Fragile	Bivariate		-0.116	0.022	-5.24	0.25		88.9%
	High	**0.026**	-0.040	0.033	-1.21	0.40	BTT, NCIDIFF, DEFDIFF	
	Low	**-0.162**	-0.119	0.021	-5.61	0.23	IRSCDIFF	
1980-1996								
Fragile	Bivariate		-0.045	0.022	-2.05	0.02		20.6%
	High	**0.106**	0.039	0.034	1.16	0.15	BTT, NCIDIFF, DEFDIFF	
	Low	**-0.125**	-0.072	0.026	-2.76	0.00	SD_NERE, TUDDIFF	
1997-2004								
Robust	Bivariate		-0.305	0.083	-3.68	0.30		100%
	High	**-0.005**	-0.081	0.038	-2.13	0.71	BTT, IRSCDIFF, NCIDIFF	
	Low	**-0.496**	-0.321	0.088	-3.67	0.30	TUDDIFF	

Appendix C

Table C.1: *Unit root test results, consumption and GDP, 1995-2005*

Country	Consumption				GDP			
	Levels	Lag	Diff	Lag	Levels	Lag	Diff	Lag
Euro Area	-1.41	4	-1.72*	1	-1.50	4	-1.94*	1
Czech Rep.	-1.83	4	-1.12	1	-3.27**	4	-2.62***	4
Estonia	-2.93*	4	-2.54**	1	-2.77	4	-3.43***	2
Hungary	-1.68	4	-0.95	1	-3.17**	4	-1.55	1
Latvia	-1.08	4	-1.51	1	-1.20	4	-1.49	1
Lithuania	-2.24	4	-1.85*	1	-1.51	4	-2.02**	1
Poland	-0.35	4	-0.31	1	-0.96	4	-1.65**	1
Slovakia	-1.92	4	-2.06*	1	-1.44	4	-1.23	2
Slovenia	-1.76	4	-0.73	4	-0.75	4	-2.36***	1
Austria	-1.65	4	-2.65***	1	-1.24	4	-2.32**	1
Belgium	-1.35	4	-2.60***	1	-2.13	4	-2.90***	2
Denmark	-1.39	4	-1.88*	1	-1.44	4	-2.27**	1
Finland	-1.43	4	-1.96**	1	-1.35	4	-1.87*	1
France	-2.21	4	-3.51***	1	-1.63	4	-1.91*	1
Germany	-0.76	4	-1.76*	1	-1.04	4	-2.05**	1
Ireland	-1.30	4	-0.72	2	-0.53	4	-0.16	4
Italy	-1.95	4	-1.88*	1	-1.36	4	-2.68***	1
Luxembourg	-1.14	4	-2.18**	1	-1.58	4	-1.99***	3
Netherlands	-1.71	4	-0.67	2	-1.83	4	-1.34	1
Spain	-2.08	4	-3.39***	1	-1.61	4	-1.15	4
Sweden	-2.14	4	-1.95*	1	-1.71	4	-2.51***	1
UK	-0.01	4	-1.95*	1	-0.39	4	-2.59***	1

*Note: Results of the DF-GLS unit root test by Elliot et al (1996), in the case of levels including a deterministic trend. The significance levels are indicated as follows: *** = 1%, ** = 5%, *=10%. The consumption and GDP data used in this test are not seasonally adjusted.*

Table C.2: *Unit root test, interest rates, NMS*

Country	Levels	Lag	Differences	Lag
Euro Area	0.35	1	-2.47***	1
Czech Rep.	-1.65*	1	-2.85***	1
Estonia	-0.85	3	-2.07**	2
Hungary	-2.33***	2	-2.25***	1
Latvia	-1.13	1	-2.71***	1
Lithuania	-0.54	1	-1.54	1
Poland	-0.80	1	-1.75*	1
Slovakia	-1.65*	1	-2.48***	3
Slovenia	-3.13***	1	-2.13**	1

*Note: Results of the DF-GLS unit root test by Elliot et al (1996). The significance levels are indicated as follows: *** = 1%, ** = 5%, *=10%. The data used in this test were calculated using year-on-year CPI inflation rates.*

Table C3: *Unit root test, interest rates, EU-15, 1980-1998*

Country	Levels	Lag	Differences	Lag
Austria	-2.12**	3	-1.42	2
Belgium	-1.54	1	-2.32***	1
Denmark	-2.09**	3	-4.77***	3
Finland	-1.03	2	-3.80***	1
France	-0.92	1	-0.67	3
Germany	-1.08	1	-2.83***	1
Greece	-0.53	2	-8.95***	1
Ireland	-0.72	2	-1.59	2
Italy	-0.91	2	-3.71***	1
Luxembourg	-1.48	1	-2.63***	1
Netherlands	-1.35	1	-2.05**	1
Portugal	-2.71***	1	-2.62***	2
Spain	-2.58***	1	-6.04***	1
Sweden	-2.90***	1	-5.59***	3
UK	-1.04	1	-5.44***	1
US	-1.622	2	-1.90*	1

*Note: Results of the DF-GLS unit root test by Elliot et al (1996). The significance levels are indicated as follows: *** = 1%, ** = 5%, *=10%. The data used in this test were calculated using year-on-year CPI inflation rates.*

Table C.4: *Unit root test results, interest rates, NMS-8*

Country	Levels	Lag	Diff	Lag
Euro Area	-0.76	3	-3.68***	4
Czech Rep.	-3.03***	1	-3.39***	4
Estonia	-1.79*	1	-4.78***	1
Hungary	-0.17	4	-3.26***	1
Latvia	-0.80	4	-2.61***	1
Lithuania	-0.70	2	-2.03**	1
Poland	-0.58	3	-2.53***	1
Slovakia	-1.58	4	-2.08**	4
Slovenia	-1.22	4	-3.20***	1

*Note: Results of the DF-GLS unit root test by Elliot et al (1996). The significance levels are indicated as follows: *** = 1%, ** = 5%, *=10%. The data used in this test were calculated using annualised month-on-month CPI inflation rates.*

Table C.5: *Unit root test results, interest rates, EU-15*

Country	1980-1989				1990-1998			
	Levels	Lag	Diff	Lag	Levels	Lag	Diff	Lag
Austria	-2.02**	4	-2.49***	1	-1.53	4	-2.41***	3
Belgium	-2.78***	1	-1.50	4	-0.81	2	-2.23**	1
Denmark	-4.77***	1	-2.19**	4	-0.59	3	-1.08	4
Finland	-1.40	4	-2.91***	1	-1.03	4	-1.75*	1
France	-1.76*	1	-2.80***	4	-1.09	2	-2.00**	4
Germany	-1.40	2	-2.35***	1	-1.39	4	-2.20**	1
Greece	-1.18	4	-3.79***	4	-2.32*	1	-2.77***	1
Ireland	-2.41***	3	-0.83	4	-2.25*	1	-2.98***	2
Italy	-1.80	1	-2.41***	4	-1.08	3	-3.19***	2
Luxembourg	-2.79***	1	-3.35***	1	-0.98	3	-0.84	4
Netherlands	-1.49	1	-1.99**	4	-0.78	4	-4.24***	4
Portugal	-1.81	4	-3.28***	4	-0.74	4	-2.32**	1
Spain	-3.08***	3	-1.73	4	-2.27*	2	-2.76***	4
Sweden	-3.11***	1	-1.71	1	-1.89	1	-2.23**	4
UK	-1.37	4	-3.23***	1	-1.33	4	-2.77***	1
US	-1.61	1	-2.23**	1	-1.25	2	-1.36	4

*Note: Results of the DF-GLS unit root test by Elliot et al (1996). The significance levels are indicated as follows: *** = 1%, ** = 5%, *=10%. The data used in this test were calculated using annualised month-on-month CPI inflation rates.*

CURRICULUM VITAE

UWE BÖWER

MUNICH GRADUATE SCHOOL OF ECONOMICS

Kaulbachstr. 45 ▪ 80539 Munich ▪ Germany

Born 25 January 1977 in Haan/Germany

DOCTORAL EDUCATION

10/2003 – 09/2006	UNIVERSITY OF MUNICH, Munich Graduate School of Economics *PhD in Economics*
04/2006 – 07/2006	UNIVERSITY OF CALIFORNIA, BERKELEY *Visiting scholar*

PRE-DOCTORAL EDUCATION

10/2002 – 01/2004	UNIVERSITY OF MUNICH, Department of Economics *MA in Economics (Distinction)*
09/2001 – 09/2002	UNIVERSITY COLLEGE LONDON, School of Public Policy *MSc in Public Policy (Distinction)*
10/1998 – 08/2001	UNIVERSITY OF MUNICH, Department of Economics *BA in Economics*

RESEARCH AND TEACHING EXPERIENCE

04/2005 – 06/2005	EUROPEAN CENTRAL BANK, DG Economics *Research intern*
10/2004 – 03/2005	UNIVERSITY OF MUNICH, Seminar for Macroeconomics *Teaching assistant*
11/2000 – 07/2001	UNIVERSITY OF MUNICH, Seminar for International Economics *Research assistant*

NON-ACADEMIC EXPERIENCE

08/2000 – 10/2000	MCKINSEY & COMPANY, INC, Financial Institutions Group *Intern*
03/2000 – 04/2000; 09/1999 – 10/1999	DEUTSCHE BANK AG, Private Banking; Corporate Clients and Real Estate *Intern*
08/1996 – 10/1998	DEUTSCHE BANK AG, Private and Business Clients *Vocational trainee*
07/1995 – 07/1996	YOUNG MEN'S CHRISTIAN ASSOCIATION *Community service (Zivildienst)*

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