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Department of Economics

Essays in Empirical Macroeconomics

Tomasz Koźluk

Thesis submitted for assessment with a view to obtaining the degree of
Doctor of Economics of the European University Institute

Florence
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Essays in Empirical Macroeconomics

Ph.D. Thesis

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Introduction

On the 1^{st} of January 1999 eleven European Union (EU) members fixed their currencies to form the Economic and Monetary Union (EMU). In the run-up to the formation of the euro area candidate countries had to fulfil a set of entry criteria laid down in the Treaty of Maastricht. These consisted of limitations on nominal exchange rate volatility, and on the levels of inflation, interest rate, public debt and fiscal deficit. After the successful formation of the common currency, the countries gave up individual monetary policy to the European Central Bank (ECB), which supposedly, at least for some, meant a gain in credibility and a substantial change in monetary policy in general. Finally, following the latest wave of EU enlargement, which took place on the 1st of May 2004, in the next years we can expect new members to be aiming to join the euro. The five nominal criteria will most certainly be applied to future enlargements. Whilst before EU entry practically all candidate countries declared a willingness to join the euro "as soon as possible" the question whether they are ready and suitable to do so, is important.

Moreover, many researchers, such as Gali and Perotti (2003) have expressed fear that the fiscal entry criteria and the Stability and Growth Pact, which did not account for cyclical conditions would cause a decrease in the level of public investment, as it is often claimed that capital spending is less visible, and thus easier to cut than current spending. If so, the restraints imposed by the Treaty and the Pact could lead to an under-provision of infrastructure investment. If this argument were to be supported by empirical evidence the preoccupation with the current inferior levels of public investment could prove to be even stronger in the New Member States. There, the current low levels of public infrastructure are regarded as a limitation to growth prospects while on the other hand in many, especially the larger New Member States fiscal deficits tend to be above or close to the 3% of GDP maximum required to join the euro area. Thus joining the EMU will require fiscal consolidation, and there is a fear that, for the political reasons mentioned above, the cut will fall on public investment spending and thus have an adverse influence on growth prospects.

Next, there are many reasons to suspect that the introduction of the common currency in Europe constituted a major economic change. Tying together 11 and later 12 currencies, replacing them with one common currency, and giving up monetary policy to the European Central Bank gives strong reasons to suspect a structural change in monetary policy, and in the behavior of both the exchange rates against third currencies and prices. Related to this, a good understanding of the link between exchange rates and prices via the exchange rate pass-through, both in the current euro members and in the New Member States is an important factor for monetary policy design and may be a relevant issue once the latter countries attempt to adopt the euro.

This thesis consists of four chapters concerning the empirical assessment of international macroeconomic issues related to the introduction of the European common currency.

The first chapter of the thesis analyzes the potential readiness of the New Member States to join the euro area. It looks at nominal and real convergence indicators, relative to the state of the same markers in the current EMU members in the 1990s. The nominal criteria are a selected set of variables similar to the Maastricht Treaty requirements, while the real criteria are derived from the Optimum Currency Area theory. As most of these criteria are regarded as endogenous, a direct comparison is rather undesirable, and an out-of-time approach is taken, that is the Central Eastern European Countries (CEECs) are compared with the current EMU members within a similar time of potential EMU entry.

The second chapter focuses on the effects of the political horizon, large fiscal cuts and the run-up to the euro area on public expenditure. Here the effect of elections, political cycles and political instability on public spending in OECD members is investigated. Additionally the chapter addresses the question whether the fiscal restrictions related to the introduction of the euro had any effect.

Chapter III analyzes the long run exchange rate pass-through, proposing a new working definition of the long run ERPT which is less arbitrary than the ones present in the empirical literature, and is based on theoretical underpinnings. It then looks deeper into whether individual industries in EMU members can be found to have full pass-through, whether this pass-through was affected by the introduction of the common currency, and if so, in which way had this effect worked.

Chapter IV is intended as a link between Chapters I and III. It deals with the estimation of short run ERPT into import prices in the Central and Eastern European States. Moreover, it extensively reviews exchange rate developments in these countries, finding that the choice of the exchange rate regime does not seem to be a major determinant of the degree to which nominal effective exchange rate fluctuations are passed on to the import price level.

All Chapters are empirical, and use up-to-date econometric tools. In more detail, the first Chapter "CEE Accession Countries and the EMU - An Assessment of Relative Suitability and Readiness for Euro Area Membership" asks the question whether the New Member States look ready to adopt the euro. Eastward enlargement of the euro area will result in transition economies sharing a currency with well-established market economies. There seem to be vast economic and institutional differences between the New Member States and the EMU participants. Therefore, in this paper, I compare the suitability (in terms of the Optimum Currency Area theory) of the candidates relative to the current members at a similar time before joining, as well as their readiness to comply with Maastricht criteria. Using fuzzy clustering and principal components, we assess patterns of convergence, possible inhomogeneities within the future 'enlarged' euro area and create readiness and suitability indexes. We find the CEECs more suitable in terms of OCA criteria and more ready than some of the (southern) current members once were. Moreover, they are not found to follow distinct convergence paths. This analysis was carried out in 2002, and at that point the assumption that all CEECs may aim at an early entry date was plausible. Although, as we know, due to various, quite often exogenous developments, many New Member States have postponed, or have been forced to postpone joining the euro, the main conclusions of the paper still appear to hold. Of the countries that have not decided to join in 2007 the CEEC-4 (Czech Republic, Hungary, Poland and Slovakia) did so mainly because of the lack of political support for early membership, connected somewhat with the unwillingness to reduce rather high fiscal deficits, thus in a sense because of the lack of the drive to fulfil the criteria rather than the technical infeasibility. As for the other 4 New Member States from Central and Eastern Europe, they seemed to aim at an early entry date; however in three cases these attempts were invalidated by the inflation criterion. In the end Slovenia seems likely to be the only country to join the EMU in 2007 however the failure of the Baltic States to satisfy the inflation criterion seems rather the fault of the (currently) exceptionally low value of the benchmark, which was not the case when the euro area was formed.

In the next chapter, "How Easily is Public Investment Cut? A Dynamic Panel Approach" using a dynamic panel of 19 OECD countries over 1971-2004 I analyze the movement of public investment spending. Controlling for downward trends in many countries, I focus on short-term movements and find investment a fairly rigid albeit pro-cyclical component of government spending. This means that on average public investment tends to fluctuate less than current expenditures, however I also find that in severe fiscal restraints it is public investment that is mostly affected. Moreover, the fiscal consolidations related to the adoption of the euro do not seem to differ in this respect. However, this cannot be treated as evidence that the fiscal rules accompanying the introduction of the euro had no effect at all. In a number of countries, strong consolidation of fiscal spending, which as we found usually affects public investment spending very significantly, could be observed during the run-up to the EMU. That is, the urge to fulfil the Maastricht Treaty criteria may have contributed to the reduction of public investment spending because of the need to reduce fiscal deficit in total. As according to the findings of the previous chapter one of the main obstacles for EMU membership of the CEECs will most probably be the fiscal deficit criterion, fiscal consolidation can be expected to occur once these countries make the attempt to join the common currency. From this chapter, it results that one may therefore expect large cuts of public investment spending which may not be desirable in countries with low levels of infrastructure. Moreover, I find a strong link between the "policy horizon" of a government and public investment, in the sense that weaker, more myopic governments prefer to lay even more weight of a fiscal consolidation on investment spending in order to preserve current spending.

The third chapter "Measuring Long Run Exchange Rate Pass-Through" is co-authored with Olivier de Bandt from the Bank of France and Anindya Banerjee from the EUI. Here, we discuss the issue of estimating short- and long-run exchange rate pass-through (ERPT) into import prices in individual industries in the euro area countries and ask the question whether the introduction of the common currency had an effect on the degree of long run pass-through. We show how the measures of this rate proposed by recent papers may be flawed. The underlying theoretical considerations suggest the existence of a long-run relationship in the sense of Engle and Granger (1987) between import unit values, the exchange rate and a measure of foreign prices. This long-run relationship is typically ignored in the empirical analysis and an ad hoc alternative definition of the long run pass-through is proposed. We describe why the Engle and Granger (1987) long-run coefficient is important, how it may be restored to the empirical analysis by taking account of varying lag length and of structural breaks, and most importantly what difference is made to the policy debate surrounding exchange rate pass-through by differentiating between alternative measures of this variable. Our discussion is undertaken using time series and panel data techniques for testing for cointegration in the possible presence of structural breaks. The main findings are that in the long run ERPT is close to unity, especially in commodity sectors over all the countries. Moreover, in the other sectors we are able to reject the hypothesis of zero pass-through in the long run more often then in previous papers. Next, even in the cases of the manufacturing sectors where pass-through equal to one is rejected, the point estimates are generally higher than in previous studies and closer to one. Finally we find that if we allow for the ERPT to change on an estimated break date, in most cases this change occurs close to the introduction of the euro or in the vicinity to the important turn-around in the euro-dollar developments. We find that long run ERPT generally increased, for which we provide possible explanations.

The final chapter focuses on the estimation of the pass-through from the nominal effective exchange rate into import prices in Central Eastern European countries. Applying similar methodology as in the previous chapter I arrive at the conclusion that ERPT was not primarily driven by the choice of the exchange rate regime. In a number of countries characterized by very different exchange rate arrangements I find full or close to full and immediate pass-through into import prices in general and especially into commodity prices. Moreover, I find that even in the short run a large degree of this effect seems to be passed on further to the consumer prices, which may potentially cause some overlooked vulnerability for the countries within the ERM II.

Within the chapters of this thesis I use a wide range of empirical tools. In Chapter I, following the work of Artis and Zhang (2002) I use fuzzy clustering which allows the search of partitions in a multi-dimensional data set. This is a some what non-standard technique and the outcome is rather indicative than rigorous, and therefore it is augmented with the extraction of principal components and the formation of an index based on the strongest of them. The latter technique is sometimes referred to as classical multi-dimensional scaling. In the next Chapter, I use standard static and dynamic panel methods, starting from a pooled OLS estimator, a Least Squares Dummy Variable (LSDV) estimator together with the Anderson and Hsiao (1981) Instrumental Variable estimator and the Arellano and Bond (1991) GMM estimator. In a set of Monte Carlo simulations, I evaluate the performance of the different estimators within panels of similar dimensions and setup and conclude that the LSDV and AB-GMM estimators seem the most appropriate for the purposes of the paper. Finally, in Chapter III and to a more limited extent Chapter IV we apply an entire array of residual based measures in order to test for cointegration - starting from simple ADF tests with different lag selection criteria, then switching to Pedroni (1999) type panel cointegration tests with no cross-sectional dependence and the Banerjee and Carrion-i-Silvestre (2006) test with a common factor type dependence. Finally, these are extended by methods that allow for structural breaks, be it in single equations, as in Gregory and Hansen (1996), or in panels.

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Chapter I
CEE Accession Countries and the EMU an Assessment of Relative Suitability and Readiness for Euro Area Membership

Abstract

Eastward enlargement of the Eurozone will result in transition economies sharing a currency with well-established market economies. We compare the suitability (in terms of Optimum Currency Area theory) of the candidates relative to current members at a similar time before joining, as well as their readiness to comply with Maastricht Criteria. Using fuzzy cluster analysis and principal components, we assess patterns of convergence, possible inhomogeneities within the future Eurozone and create readiness and suitability indexes. We find the CEECs more suitable in terms of OCA criteria and more ready than some of the current members once were. Moreover, they are not found to follow distinct convergence paths.*

JEL Classification Numbers: F33, F0, F15, C6

Keywords: *Economic and Monetary Union. Nominal Convergence. Real Convergence. Optimum Currency Area. Central Eastern Europe.*

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I.1 Introduction

On the 1^{st} of May 2004 ten European countries join the European Union. The enlargement agreement does not allow for any opt-out clause, as in the case of the U.K. and Denmark, thus these countries will be bound to eventually enter the Economic and Monetary Union. Due to the entry requirements, this should not happen earlier than 2007. As there seem to be significant benefits of joining the common European currency, it is often argued that it would be desirable for the candidate countries to adopt the euro unilaterally (see for example Nuti, 2002; Coricelli, 2002; Bratkowski and Rostowski, 2002). This option, however, is strongly discouraged by the EU, and therefore does not seem plausible for the time being. This means, that the candidate countries will have to fulfil the entrance criteria posed by the Maastricht Treaty. This leads one to ask whether these countries are ready to fulfil the requirements? Are they, in general, suitable for common currency area membership? Will this be done through steady convergence, or be a one-time effort, for example by reducing the fiscal deficit exceptionally for one year instead of more thorough reform, yielding perhaps problems in complying with the Stability and Growth Pact or any arrangement that may replace it in the future?

This paper aims to find out in what respect the Central European accession countries fit the patterns distinguished among current members of the EMU. The exercise searches for similarities among the convergence paths towards common currency adoption of the current Eurozone states and that of candidate countries. In order to shed some light on the performance of the accession states, the analysis looks at how they fit in the core and north/south periphery partition found in previous work (see for example Artis and Zhang, 2001; Artis and Zhang, 2002).

Below, the Maastricht Criteria are used to give an idea about readiness and the effort it will take to fulfil the entry requirements, while the Optimum Currency Area (OCA) characteristics serve to judge the suitability of the accession countries relative to current members. Historically, the 8 Central Eastern European (CEECs) enlargement states are former communist countries with centrally-planned economies. Six of them formed parts of other countries, and in fact of other currency unions in the early 1990s. This suggests significant differences from the current members.

If the CEECs can be found to exhibit strong, persistent dissimilarities in comparison to EMU members, this may be an indication that they may be significantly less suitable for EMU, or add to the inhomogeneity of the Eurozone. In this case, entrance will have a strong impact on the conduct of monetary policy and effects. While finding they converged in a similar way as current members did can help predict the impact of a common currency by looking at Eurozone states and their troubles in complying with the SGP, their potential gains and losses from joining the EMU. The idea of this simple experiment is to apply fuzzy clustering to look for partition among a set of current EMU members and accession states, as possible entrants to the euro area. The results are confronted with that of another multivariate analysis tool - principal components, which serves for the creation of 'readiness' and 'suitability' indexes for the candidate countries and relating them to current Eurozone states.

As both the Maastricht Criteria and Optimum Currency Area criteria are often argued to be endogenous, comparing countries already in the EMU with candidates would be problematic and in effect, undesirable. Instead, this paper focuses on an 'out-of-time' analysis, that is, it looks at data at a certain period of time before accession, assuming the earliest possible entry date for the CEECs i.e. 2007, and compares candidate countries with members, within n years before entry. That is, we perform comparative statics looking at the year 2007 − n for CEECs, 2001 − n for Greece and 1999 − n for the other EMU states. This 'out of time' analysis is certainly not flawless, but has the advantage of avoiding the *ex ante* vs. *ex post* problem. Among the drawbacks, the most serious is definitely the fact that changes of all the other characteristics - the so called 'state-of-the-world' are ignored. These include various aspects, somewhat external to our analysis, as the fact of EU membership, ERM participation and generally substantially different monetary regimes, levels of European integration, state of the world economy and integration and technology differences between the two decades. Amid these reservations, the methodology pursued still seems to yield a reasonable trade-off, although conclusions must be drawn with reservation.

As a result of operationalizing OCA theory, we tend to find a concentric core/periphery structure of the potential common currency area, with the CEECs, blending into this pattern. The transition countries converge well enough, so that within 5 years of EMU membership, they become classified as well distributed in the core/periphery pattern, where closeness to Germany seems to govern suitability. In terms of nominal convergence, some CEECs, namely the Baltic states and the Czech Republic exhibit higher readiness to fulfil the Maastricht criteria than most current members did within 5 years of entry. In fact, although often starting from a far away position, according to the methodology used, none of the transition countries are, within 5 years of potential Eurozone membership, less ready to qualify then the EMU outliers were at a similar point.

I.2 The Optimum Currency Area

The OCA theory was developed by Mundell (1961) and McKinnon (1963) and according to the primary view, an optimal currency area is a fairly homogenous region with synchronized business cycles and symmetric response to shocks, flexible prices and factor mobility. A more recent overview of the theory can be found in Tavlas (1993). The author mentions the following characteristics of optimal participants of a common currency zone:

- Synchronization of business cycles and supply/demand shocks - similar cycles, shocks and reactions reduce the need for separate monetary policies,
- Similarity of inflation rates - OCA theory attributes similar levels of inflation to similar preferences on inflation, thus a low cost of joining a common currency,
- Factor mobility - when high, is seen as a substitute for exchange rate movements in promoting external adjustment,
- Price and wage flexibility - less rigidity among or between regions results in a less likely occurrence of the situation when one region is troubled by high unemployment and the other by high inflation because of the lack of scope for real exchange rate adjustment. Therefore, flexibility serves as an asymmetric shock-absorbing mechanism,
- Goods market integration - countries with a similar production structure are less prone to asymmetric shocks, and thus face lower costs of fixing their currencies to each other, and pursuing a common monetary policy,
- Openness and economy size - small, open economies tend to prefer fixed exchange rates, as exchange rate movements have bigger disruptive effects than in relatively closed economies,
- Trade integration - joining a common currency disposes of the exchange rate risk associated with trading, thus is more favorable for countries which trade intensively with each other,

- Degree of commodity diversification - highly diversified economies are less vulnerable, when hit by sector-specific shocks,

- Small need for real exchange rate volatility - historically low exchange rate volatility suggests low cost of fixing the currencies,

- Fiscal integration - a high level of fiscal harmonization between countries, would allow for inter-regional transfers that aim at smoothing out the effects of diverse shocks.

The theory of the Optimum Currency Area has been highly criticized as the sheer fact of joining a common currency area is associated with a major change in the economy of a country. There have been many attempts to operationalize OCA theory (see for example Bayoumi and Eichengreen, 1996 and references therein), but practically all are troubled by the problem of the endogeneity of the criteria themselves. This was sharply posed by Frankel and Rose (1997) and Frankel and Rose (1998). By looking at ex-ante indicators, one cannot draw definite conclusions on the optimality of a currency union. A candidate seeming unfit for a monetary union when looking at historical OCA indicators may well turn out an optimal member once in, as the sole fact of joining a common currency changes the nature of the OCA variables. In fact, a monetary union will most probably foster an increase in trade integration. Business cycle correlation can change as a consequence of this, though the theory is not consistent about the direction of the change. More integration, through Intra-Industry Trade may foster convergence and synchronization of the cycles, but on the other hand may lead to specialization (see for example Krugman and Venables, 1993) and thereby more asymmetric shocks. Although the theoretical predictions are ambiguous, Frankel and Rose (1997) claim to find strong support for the first scenario in their empirical work.

As for the inflation rate similarity criteria, it is important to notice that OCA theory was mainly developed under the belief of the inflation-unemployment trade-off (the Philips curve). Taking the later recognized, vertical long-run Philips curve, suggests that the inflation in a country may not be actually the result of inflation preference, but perhaps of the credibility of the policy makers. Supposedly, entering a monetary union should improve credibility, thus make easier the maintenance of lower inflation.

Even labor mobility can be suspected of being endogenous. Bertola (1989) as cited in Tavlas (1993) proposes a model in which fixing the exchange rate reduces the income risk between the regions, and thus fosters more interregional mobility. Fiscal integration may in fact worsen the response to country specific shocks, as the local policies may be better fine-tuned to deal with them.

Summarizing, the suspected endogeneity of the OCA criteria poses a threat to the credibility of ex-ante analysis, though in fact historical indicators are sometimes the only tools available. This is actually an argument in favor of using similarity and convergence to member countries at the similar stage prior to entry in order to shed some light on possible outcomes for the candidates.

I.3 Accession Countries and the Eurozone

After the accession to the EU, the Central European countries will be required, by the Treaty of Maastricht, to join the EMU 'as soon as they will be ready'. One of the most stressed advantages would be the credibility gain, since the full adoption of the euro, despite historical cases of currency unions' breakdowns, seems a very strong and trustworthy commitment. The loss of the exchange rate as an adjustment mechanism for absorbing asymmetric shocks does not seem a primary concern, as it is not certain whether it serves this purpose or contrarily is a source of disturbances itself (there is some evidence for this from the current EU members, see for instance Artis and Ehrmann, 2006). Therefore, it can be expected that CEECs will not only be obliged to, but also aim for entering the EMU as soon as possible (for an overview of arguments in favor of euroization, see for example Nuti, 2002; Coricelli, 2002; Bratkowski and Rostowski, 2002). The entry conditions discussed below include EU membership and require a two year examination period. Thus, 2007 is probably the earliest plausible date for Eurozone accession, and will be the default date in the analysis conducted.[1]

I.3.1 The Maastricht Treaty Criteria

The Maastricht Treaty of 1992 defined nominal prerequisites of the economy, necessary for EMU membership. Among the requirements to be fulfilled by candidates are:

- for two years prior to entry date:
  - the nominal exchange rate remaining within the ±15% ERM II bounds, without devaluation of the central rate,
- for one year prior to entry date:
  - the inflation rate no more than 1.5% points above the average of the three EU members with lowest inflation,
  - the interest rate on long-term government bonds no more than 2% points above the average of the three low-inflation countries.
  - budget deficit not exceeding 3% of GDP,

[1] This date was the officially declared target date for Lithuania, Estonia and Slovenia. Due mainly to inflation developments the two Baltic States were eventually forced to postpone this date. For more details on this issue and an overview of the developments in the strategies to join the euro we refer the reader to Section 4.4 in Chapter IV of this thesis.

– government debt no higher than 60% of GDP,

At least two issues are worth noting: first of all, the last two of the so called 'Maastricht Criteria (MC)' have been applied somewhat less strictly to current EMU members, as they are accompanied by a clause which allows for higher values if converging or on an exceptional basis. Second, the inflation and interest rate criteria are assessed relative to EU, not EMU members, and thus may, in fact be judged relative to other applicant countries, or even the opt-out countries.

The above criteria have been widely criticized (see for instance Buiter, Corsetti and Roubini, 1993; Bratkowski and Rostowski, 2002), mainly for the arbitrariness of the values and for accounting only for the nominal side of convergence and stability, while ignoring the real side. They do not account for any cyclical adjustments, do not distinguish between various types of public spending, are to a large degree endogenous and take into examination a very short period. The assessment of the appropriateness of the criteria is beyond the scope of this paper, but as most probably forming the obligatory benchmark, they will be used below for the evaluation of the readiness of accession countries for Eurozone membership.

As mentioned, the Accession Countries upon joining the EU will be bound to enter the EMU, as soon as they fulfil the Maastricht Criteria, as no opt-out clause has been allowed.[2] In this paper, we take a look at the performance of the eight CEE candidate countries, according to the Maastricht requirements, and compare their situation to the one of current EMU states within a similar amount of years before entering the common currency.

[2]Admittedly, there is little pressure to do so, and as can be seen from the example of Sweden, which is the only 'old' EU member without an opt-out clause and outside the EMU - the decision to join can in practice be postponed.

I.4 The Data

The EMU candidate CEECs have undergone transition from centrally-planned towards market economies. This process is actually still continuing, but undoubtedly there has been a major structural break in the end of the 1980s and the beginning of the 1990s in the characteristics of these economies. Data produced by statistical offices of the centrally-planned economies is not only itself unreliable, but afterwards, in the first years of transition the countries experienced a landslide, with falling GDP and high levels of inflation, to be followed by a spectacular rebound. Therefore the figures, if at all available, can be expected to be imprecise. Hence any data before, say, 1993 are practically useless, and data for the early years must be treated with extreme caution. This problem flaws the actual choice and construction of the variables for analysis, making them far from ideal.

As mentioned, the analysis takes a specific 'out of time' approach, that is comparing countries within a certain period (years) before joining the currency union. Hence for instance, assuming the CEECs aim for entering the Eurozone in 2007, the analysis conducted for 5 years prior to membership will compare 2002 data for the accession countries, 1996 data for Greece and 1994 data for the other EMU states.

As there is no obvious way to discriminate against each other the criteria used in this analysis, thus all the variables have been standardized by subtracting the mean and dividing by the respective standard deviation.

I.4.1 Real Convergence Variables

As seen before, OCA literature suggests a number of features which make a country more likely to be suitable for common currency membership. The variables chosen for our analysis are: business cycle correlation, real exchange rate volatility, labor market flexibility, trade integration and inflation rate. The first two of the above are measured with respect to Germany, thus we are in a sense assessing the suitability of countries to adopt a common currency together with Germany. The reference core used later for the calculation of the OCA index will be Germany.

The business cycle variable is in fact more of a measure of correlation of industrial production fluctuations, due to the fact that, as explained above, shortness of the sample limits the estimation of business cycles for the CEECs. The reference is Germany, and the correlation is based on smoothed (HP-filter) monthly data over a period of 8 years, thus due to the shortness of the sample does not vary over time - a simplification necessary to avoid the disruptive influence of early 1990s data for the CEECs, and comparability with EMU-11. The real exchange rate volatility against Germany is captured by monthly observations over 2 year moving windows. Labor market flexibility is proxied by a measure of the ease of new job creation - an aggregate index created upon variables such as the duration and complexity of new business registration procedures, as well as the cost of these procedures and minimum capital required relative to GNI. Another suggested proxy was employment protection legislation, which was not used due to the fact that it would be measuring some demand side flexibility - ambiguously related to the capability of the labor market to adjust to shocks. It seems indeed unclear whether more strict employment protection would lead to less severe effects of negative shocks on economy, or contrarily slow down the speed of adjustment and lead to more persistent shocks. The degree of trade integration is measured as the share of trade done with the current EMU members. The Accession Countries, though not formally EU members exhibit relatively high integration with the Euro-11[3] comparable with that of current members. A more precise description of the data sources

[3]By Euro-11 we denote the EMU excluding Luxembourg, for which data are unavailable.

and variables creation can be found in Appendix A. Among the intentions of the real-side analysis is the construction of a Suitability Index - a one dimensional measure capturing the relative suitability of a country, according to OCA theory, for the Eurozone.

I.4.2 Nominal Convergence Variables

The choice of indicators in order to measure the readiness of countries in terms of Maastricht criteria is relatively straightforward. Data for inflation, monthly exchange rate against the ECU and euro, budget deficit and public debt are generally available. The long term interest rate on government bonds poses a minor availability problem in the case of the candidate countries, and thus has been proxied, for both Member States and Accession Countries, by the average market lending rate. In order to introduce reference points, two dummies have been added - $Dummy(0)$ with all the variables set to zero and $Dummy(MC)$ with all variables set to marginally fulfilling the MC. These will form a benchmark for the assessment, as the set of countries does not provide a appropriate reference. Contrary to the case of real convergence criteria where it serves this purpose, Germany is not performing exceptionally in terms of nominal criteria fulfillment, thus the $Dummy(0)$ serves for this purpose, while $Dummy(MC)$ serves as a cut-off value. In other words, being classified far from Germany, in terms of the nominal variables, would not necessarily mean performing worse than Germany. In order to capture countries that score better on the Maastricht criteria than obliged, we adopted $Dummy(0)$ as the reference value, while $Dummy(MC)$ is the furthest away in terms of all requirements, among the combinations still satisfying the treaty.

The details concerning the creation of the variables are presented in Appendix A. One of the results of our nominal analysis is the Readiness Index - intended to be a one dimensional measure to score the readiness of countries to fulfil the Maastricht requirements.

I.5 The Methodology

For each of the two sets of criteria, the analysis consists of two parts and the results are subsequently compared. In the first step, the fuzzy clustering algorithm is applied to search for a pattern in the data. Second, principal component analysis is used in an attempt to simplify the multivariate data set with the intention of creating a relative one-dimensional measure.

I.5.1 Fuzzy Clustering Algorithm

The fuzzy clustering algorithm used to partition the data, can be seen in more detail in Kauffman and Rousseeuw (1990) or in Höppner, Klawonn, Kruse and Runkler (1999). The exact method employed is the c-means method proposed by Dunn (1974) and Bezdek (1974). This follows the work of Artis and Zhang (2002) and Boreiko (2003).

In our case, the data set consists of n countries, and p variables. Each object x_i is characterized by a vector of features ($x_i = \{x_{i1}, .., x_{ip}\}$ for $i = 1..n$), where each variable is standardized with mean zero and unit variance. The dissimilarity index $d(i, j)$ is the Euclidian distance between the two objects x_i and x_j in p-dimensional space:

$$d(i,j) = \sqrt{\sum_{r=1}^{p}(x_{ir} - x_{jr})^2} \tag{1}$$

The objective of the algorithm is to minimize the following objective function G:

$$G = \sum_{k=1}^{m} \frac{\sum_{i=1}^{n} \sum_{j=1}^{n} u_{ik}^2 u_{jk}^2 d(i,j)}{2\sum_{j=1}^{n} u_{jk}^2} \tag{2}$$

subject to the following constraints:

$$u_{ij} \geq 0 \text{ for } i = 1..n,\ j = 1..m$$

$$\sum_j u_{ij} = 1 \text{ for } i = 1..n \tag{3}$$

The outcome of the algorithm is a matrix U_{nxm}, where element u_{ij} is the membership coefficient, or the "degree of belongingness" of object i to cluster j, while m is the number of clusters. Elements in each of the n rows sum up to 1. For each object i, a relatively high, dominating value of one of the membership coefficients u_{ij} allows for assigning i to hard cluster j with high certainty. As for diagnostics, in order to assess how well partitioned the data are, the following Dunn coefficient is proposed:

$$F_m = \sum_{i=1}^{n} \sum_{v=1}^{m} u_{iv}^2 / n \tag{4}$$

and we will apply the normalized version of the coefficient:

$$FN_m = \frac{F_m - (1/m)}{1 - (1/m)} = \frac{mF_m - 1}{m - 1} \tag{5}$$

which after inserting equation (4) into (5) yields:

$$FN_m = \frac{m \sum_{i=1}^{n} \sum_{v-1}^{m} u_{iv}^2 / n - 1}{m - 1} \tag{6}$$

The above takes values from 0 - complete fuzziness, when membership indices have the same value, to 1 - no fuzziness, when each object is assigned to a certain cluster with the membership coefficient of 1. In the latter case we can speak of a 'clear' or 'hard' partition.

The measure of the quality of classification in the hard cluster, the silhouette width of object i is calculated as follows:

$$s(i) = \frac{b(i) - a(i)}{max[a(i), b(i)]} \tag{7}$$

$$-1 \leq s(i) \leq 1$$

where $a(i)$ is the average dissimilarity of i from all objects in the same cluster (in our case the average Euclidian distance $1/(n_{cl(i)} - 1)\sum_{j=1, j\neq i}^{n_{cl(i)}} d(i,j)$) and $b(i)$ the minimum (across all other clusters) of the average dissimilarity of i from all the objects in each single other cluster. When $s(i)$ is close to one, this implies that $a(i)$ is small with respect to $b(i)$, that is, the object is well classified in the appropriate cluster. If $s(i)$ is close to zero, this implies that $a(i)$ and $b(i)$ are approximately equal, thus it is unclear which cluster should i belong to. Negative $s(i)$ implies i is assigned to the wrong cluster. Silhouette width values for clusters and the whole data set indicate the quality of respectively cluster and total partition. In order to choose the optimal number of clusters m average silhouette maximization has been applied.

I.5.2 Principal Component Analysis

Principal component analysis is a multivariate analysis tool, which aims at reducing the number of variables in the data. In fact, in a multi-variable data set, it is often the case that groups of variables move together. This may be a sign of the redundancy of information as variables may be driven by common underlying forces, thus being only a realization of the core structure of the data set. Extracting the principal components allows for a simplification of the data by replacing the variance of a group of variables with a single new one. Each principal component is a linear combination of the original variables, that is the data matrix X with n observations and p variables can be transformed into the Z matrix, where:

$$Z_i = \alpha_{i1}X_1 + \alpha_{i2}X_2 + ... + \alpha_{ip}X_p \qquad \text{for } i = 1..p \tag{8}$$

Principal components have two distinctive features:

- the p components are orthogonal to each other, thus there is no redundancy of information,
- the first component explains the largest percentage of the variation in the original p-dimensional data set (the second principal component explains the second largest percentage and so on).

Although there is not necessarily a pure dimensional gain (there are p principal components, from p initial variables if not perfectly co-linear), often the first few principal components account for most of the variation while the contribution of the rest is negligible. The problem of extracting principal components is basically that of obtaining the eigenvectors and eigenvalues of the data correlation matrix, and arranging the eigenvalues in decreasing order. The highest eigenvalue will distinguish the first principal component and the corresponding eigenvector will contain the variable loadings - that is the α's.

The application of principal component analysis in creating aggregate indexes for multivariable analysis follows the work of Nicoletti, Scarpetta and Boylaud (2000) and is generally a method of classical multi-dimensional scaling. The reduced dimension framework allows for creating a more straightforward index of 'closeness', data patterns presentation, partition and interpretation. In the first step, Bartlett's 2 test is used for finding the dimensionality of the data set, which is then used as guidance for selecting the number of principal components used for creating the index. Second, the principal components themselves are extracted and the old data set is transformed with respect to them. The following requirements are used to find the exact number of components preserved for further analysis:

- cumulatively, they explain at least 60% of the sample variance,
- each of the components is associated with an eigenvalue greater than 1,
- individual contribution of each principal component in explaining overall variance is at least 15%.

Next, upon the previously selected most significant components, the weighted Euclidian distance from the reference values is taken:

$$IND(i) = \sqrt{\frac{\sum_{j=1}^{m}(z_{ij} - z_{Rj})^2 w_j}{\sum_{j=1}^{m} w_j}} \tag{9}$$

where m is the selected number of principal components, w_j is the percent of variance explained by component j, z_{ij} is the value of new variable z_j for country i (see equation (8)). The reference values R are: in the case of Real Convergence - Germany, and in the case of Nominal Convergence - $Dummy(0)$. As the purpose of the component extraction is the formation of the indexes we will not be troubled by the interpretation of the components themselves, which often proves to be problematic.

The main weakness of the principal components approach is the sensitivity to basic data modifications. Revisions, updates or inclusions of other countries affect the variance of the data set, and thus the principal components themselves.

I.6 Results and Discussion

Clustering appears to be an interesting method of analyzing EMU convergence. Appropriately, allowing for a fuzzy partition permits us to make use of a much broader spectrum of information. The main advantage over hard clustering is that observations, in our case countries, are not strictly allocated to single clusters, but given a coefficient of belongingness to each cluster. This allows the determination, not only of similarities between countries inside the clusters, but also the degree of inter-cluster country correspondence, as well as of similarities between clusters. The application of principal components introduces more rigor to the results and allows the formation of suitability and readiness indexes, which though rough, give a clearer idea on how candidates perform relative to each other in terms of OCA membership and Maastricht Criteria compliance.

I.6.1 OCA Criteria

The results of fuzzy cluster analysis applied to the OCA criteria are displayed in Table 1. Due to the fact that two out of five variables are time invariant proxies, only two periods of examination have been taken: 11 and 5 years before potential accession. In both the optimal number of clusters is 5, and the data exhibits quite a high degree of fuzziness - the Dunn's normalized coefficient is in equal to 0.31. This strengthens the argument for using the fuzzy version of the clustering algorithm. Moreover, in both cases the partition appears quite sound - none of the countries is misclassified, and the lowest object silhouette is 0.46. We consider Germany as the default common currency member in the whole of Real Convergence analysis. Thus, within 11 years before potential adoption of the euro, 'the core', that is the countries most suitable to join, lies between cluster I and II. Germany's, and hence 'the cores' coefficients are 49% and 20% respectively. The country with almost

	OCA Criteria 11 years before EMU						OCA Criteria 5 years before EMU					
	Clusters					Country	Clusters					Country
Country	I	II	III	IV	V	Silh.	I	II	III	IV	V	Silh.
Austria	**.77**	.07	.08	.06	.02	.84	**.42**	.25	.25	.05	.03	.50
Belgium	**.76**	.07	.09	.05	.02	.81	**.44**	.26	.23	.05	.03	.46
Finland	.11	**.43**	.11	.27	.07	.89	.11	.13	.08	**.60**	.07	.82
France	**.66**	.16	.09	.07	.02	.76	**.70**	.19	.06	.04	.01	.73
Germany	**.49**	.20	.14	.11	.06	.72	**.47**	.27	.15	.07	.04	.67
Greece	.00	.00	.00	.00	**1.0**	1.0	.00	.00	.00	.00	**1.0**	1.0
Ireland	.12	**.52**	.11	.21	.04	.89	.16	.17	.09	**.53**	.04	.70
Italy	**.86**	.05	.05	.02	.01	.83	.16	**.41**	.29	.11	.03	.58
Netherl.	**.49**	.23	.12	.10	.06	.52	**.57**	.21	.11	.08	.03	.74
Portugal	**.43**	.12	.28	.08	.08	.54	.12	.22	**.58**	.04	.05	.73
Spain	**.66**	.09	.17	.05	.03	.73	.07	.19	**.70**	.02	.02	.43
CzechR.	**.63**	.08	.21	.05	.03	.63	.26	**.42**	.24	.05	.02	.48
Estonia	.10	.24	.20	**.42**	.04	.52	**.47**	.25	.12	.13	.04	.61
Hungary	.12	.10	**.55**	.17	.06	.67	.03	.06	**.89**	.01	.01	.72
Latvia	.05	.25	.12	**.56**	.03	.57	.09	.11	.05	**.72**	.02	.55
Lithuania	.07	.15	.15	**.57**	.07	.68	.05	.05	.03	**.84**	.02	.82
Poland	.10	.07	**.68**	.12	.03	.61	.19	**.40**	.22	.15	.04	.72
Slovakia	.17	.17	**.42**	.19	.05	.46	.15	**.55**	.24	.04	.02	.48
Slovenia	**.65**	.07	.20	.05	.04	.65	.13	.17	**.59**	.04	.06	.76
CL.Silh.	.70	.89	.58	.59	1.0	.70	.62	.56	.66	.72	1.0	.65
Dunn			.3124						.3138			

Table 1: Optimum Currency Area - fuzzy clustering results. Column "Country Silh." gives country silhouette value. Row "CL. Silh." gives cluster silhouette value and "Dunn" gives Dunn's fuzziness coefficient. Bold numbers indicate hard cluster assignment.

identical distribution among clusters is the Netherlands, and thus is the primary candidate for joining Germany in a common currency, exhibiting low inflation and real exchange rate volatility, high trade integration and labor market flexibility, together with an average business cycle correlation. Next is France, with a high business cycle correlation and less flexible labor market. Austria and Belgium are found to be very similar to each other, and moreover very close to the core. Italy is also close to the two, except for a higher inflation level. Further away, yet still in cluster I there are: Portugal and Spain joined by the Czech Republic and Slovenia. Eleven years before accession these exhibit a large degree of similarity. Cluster II with Finland and Ireland, though with coefficients of 43% and 52%,

Hard Cluster	Business Cycle	Real Ex.Rate	Labor Market	Trade Integrat.	Inflation
IT,AUS,BEL,FR, SP,SLN,CZ,GER, NL,PT	Med–Hi	Low–Med	Med-Low	Med–Hi	Low–Med
FIN,IRL	Low	Med	Hi	Low	Low-Med
PL,HU,SLK	Hi	Hi	Low	Med–Low	Hi
LIT,LAT,EST	Low–Hi	Med	Hi	Low	Hi
GR	Low	Low	Low	Hi	Hi

Table 2: OCA 11 years before EMU - hard cluster characteristics.

is characterized by low trade integration and business cycle correlation, high labor market flexibility and medium real exchange rate volatility. Both, but especially Finland, show a high degree of resemblance to the three Baltic States, classified in a hard separate cluster (IV) mainly due to higher inflation and real exchange rate volatility. Greece remains in a separate cluster, which suggests that according to OCA criteria, in 1990, 11 years before acceding to the EMU, Greece was not part of the Germany-based optimal currency area. Cluster III is composed of Hungary, Poland and Slovakia - characterized by high business cycle correlation, exchange rate volatility and inflation, and low labor market flexibility. If we consider the fact that all CEECs excluding Lithuania (15%) and Latvia (12%) have coefficients of belongingness to this cluster higher than 20%, we can interpret this as the Central European periphery. It is worth noting that this cluster exhibits a high degree of similarity to Portugal (28%) and Spain (17%).

Table 2 shows the cluster characteristics, which together with the results in Table 1, allow us to roughly sketch a primary view of the pattern:

- **the core** - Germany, Netherlands, France, Austria and Belgium, followed by Italy and further by Spain, Slovenia, Czech Republic and Portugal.

- **the northern periphery** - Finland and Ireland, with some similarity to the Baltic

Hard Cluster	Business Cycle	Real Ex.Rate	Labor Market	Trade Integrat.	Inflation
FR,NL,EST,GER, AUS	Med-Hi	Low-Med	Med-Low	Med-Hi	Low-Med
SLK,CZ,IT,PL	Hi-Med	Hi-Med	Low-Med	Med	Low-Med
HU,SP,PT,SLN	Med-Hi	Med	Low	Hi	Hi
LIT,LAT,FIN,IRL	Low	Low	Hi-Med	Low	Low-Med
GR	Low	Med	Low	Med	Hi

Table 3: OCA 5 years before EMU - hard cluster characteristics.

States.

- **the 'transition periphery'**- mainly Poland, Hungary and Slovakia, but close to all other CEECs.

The picture changes significantly when we move 6 years forward. The transition economies, further away from the early 1990s' chaotic period, tend to have stabilized their economies, successfully decreased inflation and increased trade integration with current EMU members. Hence, the notion of the transition countries and current EMU states forming diverse clusters, fades away. The level of inhomogeneity is not as intense, and patterns composed of a mix of both CEECs and EU countries emerge. The core again lies between clusters I (Germany's coefficient 47%) and II (Germany 27%). Germany, France, Netherlands, Austria and Belgium show persisting strong resemblance. They are joined in cluster I by Estonia, which lowered its inflation substantially over this time, while accompanied by a strong decrease in real exchange rate volatility - thus became the primary CEEC candidate for the euro in terms of OCA criteria. Though in a separate cluster (II) the Czech Republic and to a slightly lesser extent Poland and Slovakia, are also very close to the core - their fuzzy coefficients exhibit a strong similarity to Germany. Cluster III can be interpreted as a signal of existence of the southern periphery - relatively strong belongingness of Spain, Portugal, Hungary and Slovenia also joined by fairly similar Italy (29% compared to 11% in cluster II) - all with average real exchange volatility, high trade integration and inflation, and low labor market flexibility. Cluster IV indicates the strengthening of the ties between the Baltic States and the northern periphery, with the exception of Estonia which as indicated moved closer to Germany, due to still higher inflation and business cycle correlation, but still exhibits a coefficient of 13% in the northern periphery due to low trade integration and a relatively flexible labor market. Greece tends to form a separate cluster, mainly due to negative business cycle correlation, low labor market flexibility, and very high inflation. Thus the pattern of inhomogeneities changed over the 6 year period, and can be summarized as follows:

- **the core** - Germany, Austria, Belgium, France and the Netherlands joined by Estonia;
- **the southern periphery** - Hungary, Spain, Slovenia and Portugal with significant closeness of Italy;
- **the eastern periphery** - Slovakia, Czech Republic and Poland, surprisingly close to the core and showing similarity with Italy but also Estonia and Slovenia;
- **the northern periphery** - Lithuania, Latvia, Finland and Ireland, with some persisting resemblance to Estonia;
- **the persistent outlier** - Greece, showing some weak similarities with the southern periphery;

Cluster analysis yields the emerging concentric core periphery pattern for real convergence, and the diffusion of the CEECs between the intra-EU peripheries, yet these results lack some rigor. They constitute a starting point and principal components are used to seek confirmation of relative convergence and performance according to the OCA criteria.

OCA	2 first components	
time	11 y.	5 y.
1^{st} var. explained	43.46	49.72
2^{nd} var. explained	27.55	24.84
Bartlett's test $p - value$		
n=2	.01	.02
n=3	.02	.07

Table 4: OCA Analysis - two first principal components: percentage of variance explained by each component and dimensionality test.

The results of the PCs analysis displayed in Table 4 (more details in the Appendix B), are sufficient to allow focusing on the first two components, as in both cases they explain over 70% of the variance whilst meeting other previously stated requirements. Bartlett's test yields the non-rejection of the dimension of the data set equal 2 at 99% confidence level in the case of 11 years prior entry and at 95% level at 5 years prior to entry, thus together with the above information, allows the preservation solely of the first two components for further analysis, without an important loss of information.

The OCA Suitability Index is used to judge relative convergence. It is a transformation into one dimension, hence results in the loss of some information compared to the PC graphs (Figures 2 and 4), and obviously compared to cluster analysis, but facilitates interpretation. The country performance has been presented in Figure 1 (11 years) and Figure 3 (5 years). The first apparent observation is that the OCA criteria ordering is certainly negatively correlated with geographical distance from the core of the common currency (Germany). Eleven years before entry, the CEECs still constitute somewhat of a separate entity - their suitability is certainly lower than of most EU members, though some sort of 'gravity' forces are visible. Five years before membership, the diversity of the former east-block plays a

	OCA 11	OCA 5		OCA 11	OCA 5
Austria	0.62	0.57	Spain	1.01	1.07
Belgium	0.65	0.44	Czech R.	1.25	0.43
Finland	2.62	3.26	Estonia	2.66	0.85
France	0.64	0.36	Hungary	2.93	1.09
Germany	0.00	0.00	Latvia	2.78	1.95
Greece	2.51	2.96	Lithuania	4.26	2.98
Ireland	2.26	2.29	Poland	2.23	0.94
Italy	0.78	0.95	Slovakia	1.98	0.58
Netherlands	0.48	0.49	Slovenia	1.27	1.68
Portugal	1.49	1.50			

Table 5: Suitability Index - columns "OCA 11" and " OCA 5" denote values at, respectively, 11- and 5-years prior to entry date.

much smaller role, and the suggested 'gravity' pattern strengthens. The striking result is the apparent rings formed by EMU candidates - Germany's neighbors, excluding Poland seem most appropriate with OCA index values below 0.6, then followed by the second group - Italy, Poland, Spain, Hungary, and the furthest away geographically - Estonia, all below 1.2. The third group, constitutes a somewhat more peripheral set containing Portugal and Slovenia - below 1.8. Finally, the two last groups from the ring of least suitable according to the OCA criteria - Latvia, Ireland, Lithuania and Finland from the north together with Greece from the south. This, somewhat gravitational pattern suggests that close trade and economical ties govern our criteria. In fact, the amount of trade done especially with Germany, combined with high business cycle correlation and low real exchange volatility exhibited by the countries with close ties to Germany, overwhelm any fading influences of the transition for the CEECs. Further away geographically, these seem to matter less and thus the countries within larger distance form the peripheries.

The comparative statics approach yields:

- a group of **stable optimum currency area members**, consisting of Germany, France, Belgium, the Netherlands, and Austria;

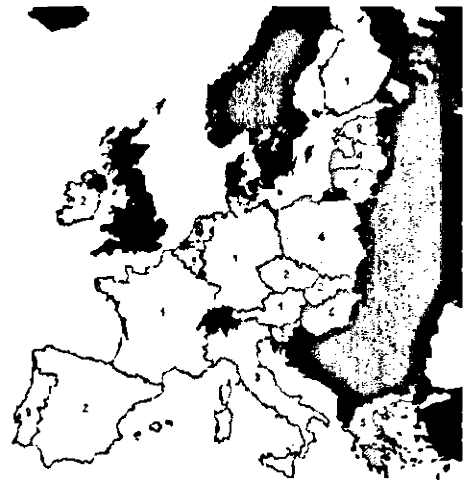

Figure 1: EMU candidates 11 years prior to (potential) entry date - OCA Suitability Index. Lower number (lighter color) indicates 'closer' in terms of OCA criteria. Black - not classified.

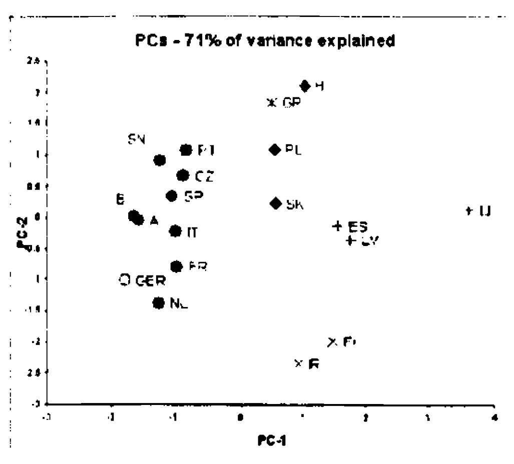

Figure 2: OCA 11 years before entry - clusters and PC's compared. Objects in same hard cluster have identical icons (Germany inverted).

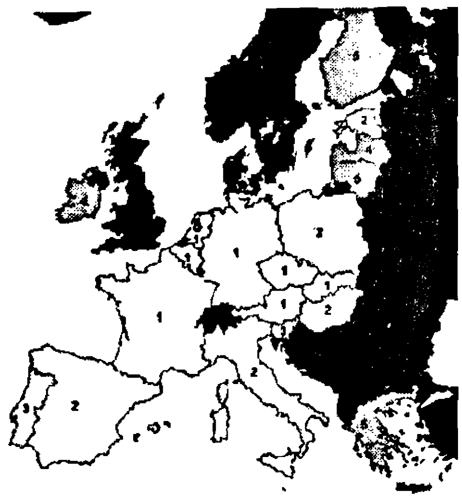

Figure 3: EMU candidates 5 years prior to (potential) entry date - OCA Suitability Index. Lower number (lighter color) indicates 'closer' in terms of OCA criteria. Black - not classified.

- a group of **converging states**, changing fairly rapidly: Czech Republic, Slovakia, Estonia, Poland, Hungary, still rather far Latvia and somewhat further Lithuania;
- a group of **relatively close but stable countries** - the southern periphery - Italy, Spain, Portugal joined by Slovenia;
- a group of **outliers**, which are, according to the criteria used, least optimal for the euro: Ireland, and diverging Greece and Finland.

The fact that principal component analysis allows us the reduction of the data set to two dimensions, preserving over 70% of the variance, permits a more informative illustration of OCA criteria performance. Though perhaps a bit rough, the graphs of the first two of new variables obtained through PCs, allow for a comparison of our index creation methodology and cluster analysis. In Figure 2 (11 years) and Figure 4 (5 years) the different hard clusters are distinguished by different labels. In Figure 2 a clear distinction between the 'core' made up mostly of current EMU members excluding Greece and the far north - Finland and Ireland. The last two countries form a cluster which is distinguishably far from other

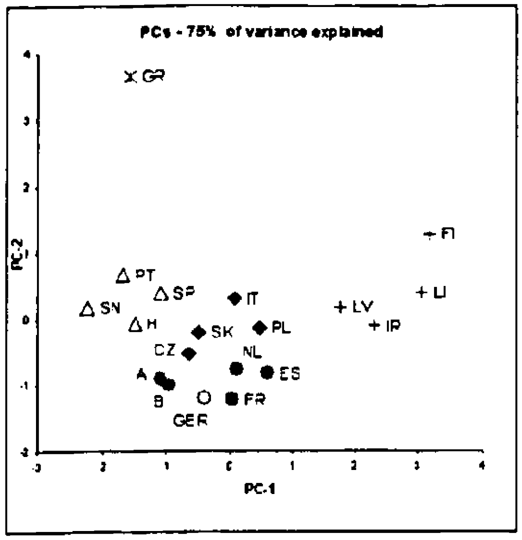

Figure 4: OCA 5 years before entry - clusters and PCs compared. Objects in same hard cluster have identical icons (Germany inverted).

clusters, but closest to the Baltic States. The eastern block countries tend to fit in the medium values of the first component and generally the cluster findings are well confirmed.

The two dimensional plot of the principal components analysis results 5 years prior to EMU membership, is visible in Figure 4. The northern periphery is apparent and similarly the southern periphery - Portugal, Spain, Slovenia and Hungary are plotted fairly close to Italy, and less to the outlying Greece. Hard clusters I and II form the respectively the strict and wider core of the EMU. This somewhat confirms the notion that best fit for foregoing own currency are, according to OCA theory, small, open economies. The CEEC economies are relatively small compared to EU members and the level of openness is on average similar. Additionally, the apparent concentric core periphery pattern, suggests some sort of a gravity model seems to be a next step extension to capture appropriateness for the OCA.

I.6.2 Maastricht Criteria

After constructing the OCA Suitability Index, we turn to assessing nominal convergence. The sheer fact that countries are found suitable to join the EMU according to the selected 'real' Optimum Currency Area variables, does not necessarily mean that they will not have trouble in complying with the Maastricht 'nominal' requirements and thus does not imply actual EMU readiness. In this part we perform a similar analysis to the above, but with respect to the Maastricht criteria.

Table 6 displays the cluster analysis results, for 11, 8 and 5 years before potential EMU accession. As contrary to the OCA exercise, all five of the variables used can be measured yearly, three periods are reported - this contributes to the analysis of the convergence patterns. In the three periods examined the degree of fuzziness if fairly high - between 0.34 and 0.40. The optimal number of clusters is rather high - 7 and 8, but this may be in part attributed to the fact of inclusion of the two 'artificial' dummies that tend to cluster away from other objects. However except for 5 cases, the average silhouettes are above 0.60 and none of the objects is misclassified. In the period of 11 years before EMU membership, we find Germany, France and Austria together with Spain and Portugal. The Czech Republic and Slovakia join this cluster, mainly because of fairly similar inflation, nominal exchange rate volatility and interest rates. This association is weakened by lower debt levels of the two transition economies. Cluster II contains countries with high budget deficit and public debt and relatively low inflation - Ireland, Netherlands, Belgium and to a lesser extent - with a coefficient of 30% - Italy (mainly due to higher inflation). In fact Italy is also partitioned close to Greece (note Italy's coefficient of 21% in cluster VII), which though in a separate cluster, has similarly a very high deficit and high public debt, fairly low exchange rate volatility but much higher inflation and interest rate. Finland clusters somewhat between

Maastricht Criteria 11 years before membership

Country	I	II	III	IV	V	VI	VII	Country Silh.
	Clusters							
Austria	**.60**	.19	.06	.06	.04	.02	.02	.66
Belgium	.10	**.68**	.05	.04	.05	.03	.05	.76
Finland	.20	.09	**.27**	.22	.11	.07	.04	.54
France	**.58**	.10	.16	.07	.04	.02	.02	.67
Germany	**.55**	.15	.15	.06	.05	.02	.02	.67
Greece	.00	.00	.00	.00	.00	.00	**.99**	1.0
Ireland	.08	**.81**	.03	.02	.03	.01	.01	.39
Italy	.16	**.30**	.06	.08	.11	.08	.21	.45
Netherlands	.14	**.71**	.04	.03	.04	.02	.02	.41
Portugal	**.32**	.19	.05	.16	.10	.09	.10	.41
Spain	**.57**	.10	.05	.10	.11	.04	.03	.63
Czech R.	**.37**	.06	.09	.31	.09	.07	.03	.42
Estonia	.15	.06	.08	**.36**	.08	.22	.06	.63
Hungary	.06	.05	.03	.12	.09	**.57**	.07	.76
Latvia	.07	.03	.03	**.44**	.06	.32	.04	.57
Lithuania	.04	.03	.02	.11	.06	**.71**	.04	.56
Poland	.04	.02	.02	.12	.04	**.71**	.04	.52
Slovakia	**.66**	.05	.06	.14	.05	.03	.02	.72
Slovenia	.04	.02	.01	**.85**	.03	.04	.01	.58
Dum(0)	.02	.01	**.95**	.01	.01	.00	.00	.56
Dum(MC)	.01	.00	.00	.00	**.98**	.00	.00	1.0
CL. Silh.	.60	.50	.55	.59	1.0	.61	1.0	.61
Dunn				.3873				

Maastricht Criteria 8 years before membership

Country	I	II	III	IV	V	VI	VII	Country Silh.
	Clusters							
Austria	**.87**	.04	.02	.02	.01	.02	.02	.67
Belgium	.11	.06	**.66**	.05	.03	.06	.05	.76
Finland	.12	.25	.03	**.44**	.05	.09	.04	.54
France	**.65**	.12	.04	.05	.08	.05	.02	.67
Germany	**.76**	.10	.03	.03	.03	.03	.02	.67
Greece	.00	.00	.00	.00	.00	.00	**.99**	1.0
Ireland	**.47**	.11	.16	.08	.07	.09	.01	.40
Italy	.06	.05	**.75**	.03	.02	.05	.21	.45
Netherlands	**.67**	.09	.08	.06	.03	.06	.02	.41
Portugal	.12	.10	.14	.08	.04	**.34**	.10	.42
Spain	**.31**	.25	.08	.11	.03	.20	.03	.64
Czech R.	.10	.24	.05	**.45**	.06	.08	.03	.42
Estonia	.25	**.44**	.05	.08	.08	.08	.06	.64
Hungary	.05	.04	.03	.05	.01	**.79**	.07	.77
Latvia	.07	**.80**	.02	.05	.02	.03	.04	.58
Lithuania	.14	**.38**	.11	.19	.06	.10	.04	.56
Poland	.07	.11	.04	**.51**	.02	.23	.04	.52
Slovakia	.05	.06	.03	.07	.02	**.75**	.02	.72
Slovenia	**.32**	.24	.05	.13	.08	.17	.01	.58
Dum(0)	.00	.00	.00	.00	**.99**	.00	.00	.56
Dum(MC)	.04	.06	.02	**.82**	.01	.05	.00	1.0
CL. Silh.	.61	.48	.81	.62	1.0	.65	1.0	.61
Dunn				.3965				

Maastricht Criteria 5 years before membership

Country	I	II	III	IV	V	VI	VII	VIII	Country Silh.
	Clusters								
Austria	**.77**	.04	.06	.02	.05	.03	.02	.01	.72
Belgium	.30	.09	.13	.07	.17	.11	.05	.07	.57
Finland	.08	.05	**.58**	.02	.09	.15	.02	.01	.25
France	**.53**	.08	.14	.04	.08	.06	.05	.02	.44
Germany	**.47**	.16	.09	.08	.07	.06	.04	.03	.38
Greece	.00	.00	.00	.00	.00	.00	.00	**1.0**	1.0
Ireland	**.30**	.17	.17	.05	.09	.13	.05	.03	.48
Italy	.11	.06	.14	.06	**.36**	.16	.03	.09	.41
Netherlands	**.80**	.04	.05	.02	.04	.03	.01	.01	.73
Portugal	.09	.06	.09	.10	**.49**	.09	.02	.07	.44
Spain	.08	.05	.14	.04	**.47**	.18	.02	.02	.32
Czech R.	.15	.13	**.40**	.04	.08	.13	.05	.02	.69
Estonia	.12	**.40**	.09	.10	.06	.07	.15	.02	.78
Hungary	.12	.05	.12	.07	**.46**	.09	.03	.05	.61
Latvia	.11	.26	**.30**	.05	.07	.15	.04	.02	.54
Lithuania	.13	.22	**.25**	.05	.07	.14	.12	.02	.59
Poland	.07	.07	.21	.04	.14	**.41**	.02	.03	.82
Slovakia	.03	**.83**	.03	.03	.02	.03	.02	.01	.77
Slovenia	.00	.00	.00	**.99**	.00	.00	.00	.00	1.0
Dum(0)	.00	.00	.00	.00	.00	.00	**.99**	.00	1.0
Dum(MC)	.02	.03	.06	.01	.04	**.82**	.01	.01	.80
CL. Silh.	.55	.77	.52	1.0	.45	.81	1.0	1.0	.64
Dunn					.348				

Table 6: Nominal Convergence - fuzzy clustering results. Column "Country Silh." gives country silhouette value. Row "CL. Silh." gives cluster silhouette value and "Dunn" gives Dunn's fuzziness coefficient. Bold numbers indicate hard cluster assignment.

Hard Cluster	Budget Deficit	Public Debt	Nominal Ex.Rate	Interest Rate	Inflation
SLK,AUS,FR,SP GER,CZ,PT	All	Med-Low	Med-Low	Med-Low	Med-Low
IRL,NL,BEL,IT	Hi	Hi	Low-Med	Low-Med	Low-Med
D(0), FIN	Low	Low	Low-Med	Low	Low-Med
SLN,LAT,EST	Low	Low	Low	Hi	Hi
GR	Hi	Hi	Low	Hi	Hi
LIT,PL,HU	Med-Hi	All	Hi	Hi	Hi
D(MC)	Med	Med	Hi	Med	Low

Table 7: MC 11 years before EMU - hard cluster characteristics.

the core cluster I (20%), cluster IV with Estonia, Latvia and Slovenia (22%) and cluster III with the *Dummy*(0) (27%) indicating in fact that it is closest to the zero values and, at least at this point, would have no problem complying with the Maastricht requirements. As emphasized before, the analysis of 11 years before Eurozone entry, is troubled by the somewhat chaotic period of rapid transition for the CEECs. Therefore, it is not surprising that the CEECs exhibit quite a high degree of correspondence, distinguishable from the EU members. Hence, clusters IV and VI, are not only fairly similar to each other, but contain basically all CEECs excluding Slovakia. The latter shows some similarity with cluster IV(14%), as does the Czech Republic (31%). Cluster VI consists of Poland, Lithuania (both 71%) and Hungary(57%) - and is distinguishable from IV because of much higher deficit and debt. The dummies do not contribute much to the interpretation, but one must bear in mind that they also serve a purpose of capturing convergence.

Moving ahead 3 years, we see a strengthening of the current EMU members core in cluster I. Germany, France, Netherlands and Austria become increasingly similar, all with coefficients above 0.65. They are joined by Ireland (47%), which scores also average on all the variables, except for public debt, which although reduced from the previous period, still

Hard Cluster	Budget Deficit	Public Debt	Nominal Ex.Rate	Interest Rate	Inflation
AUS,GER,NL,FR SLN,SP	Med	Med-Hi	Low–Med	Med-Low	Med-Low
IT,BEL	Hi	Hi	Low	Med	Med-Hi
LIT,LAT,EST	Hi-Med	Low	Low	Med-Low	Low–Med
D(MC).PL,CZ,FIN	Med	Low–Med	Hi	All	All
HU,SLK,PT	Med-Hi	Med-Hi	Med-Hi	Hi	Hi
D(0)	Low	Low	Low	Low	Low
GR	Hi	Hi	Med	Hi	Hi

Table 8: MC 8 years before EMU - hard cluster characteristics.

remains high. Spain and Portugal shift away, towards cluster II and VI, because of high inflation, interest rate and budget deficit. The Czech Republic, Poland and Finland join the *Dummy*(*MC*) mainly due to close to 3% of GDP deficits and relatively high exchange rate volatility. The Czech Republic shows also close resemblance (24%) to cluster II, that is Estonia, Latvia and Lithuania, which achieved significant nominal stabilization by fixing exchange rates, reducing inflation, maintaining very low government debt but also relatively high deficits. Belgium and Italy in cluster II still exhibit very high public debt and deficit, again with some (21%) resemblance of the latter to Greece, also with a very high debt and deficit, but clustered separately due to high inflation and interest rates. Cluster VI contains countries with still persistent high inflation and interest rates, and relatively high values of all other variables - these economies are not converging, at least not as quickly and are still characterized by a significant amount of instability. They include Hungary, Slovakia (which actually diverged, mainly due to a jump in inflation and interest rate), and to a lesser degree Poland (23%) and Slovenia(17%). They show some correspondence with the high inflation and interest rate Iberian countries.

Within 5 years of membership, the CEECs managed to achieve further stabilization of

Hard Cluster	Budget Deficit	Public Debt	Nominal Ex.Rate	Interest Rate	Inflation
NL,AUS,FR,GER BEL,IRL	All	Hi–Med	Low–Med	Low–Med	Low–Med
FIN,CZ,LAT,LIT	All	Low	Low–Hi	Low–Med	Low
SLK,EST	Low	Low	Med–Low	Low–Med	Hi–Med
PT,SP,HU,IT	Hi	Hi–Med	Hi–Med	Hi–Med	Hi–Med
SLN	Low	Low	Low	Hi	Hi
D(MC),PL	Med	Med	Hi	Hi–Med	Hi–Med
D(0)	Low	Low	Low	Low	Low
GR	Hi	Hi	Med	Hi	Hi

Table 9: MC 5 years before EMU - hard cluster characteristics.

their economies. The EU members generally ran high deficits, thus the core is joined by Belgium (though only marginally - 30%, because of public debt levels among the highest in the EU) and by Ireland which managed to further reduce its debt burden. Italy, Portugal, Spain and Hungary strengthen their resemblance, thus the idea of a southern periphery seems justifiable. Poland stays close to the Dummy(MC), i.e. the cut-off values, but fails to converge further. Finland, Czech Republic, Latvia and Lithuania seem to remain among the prime candidates in terms of readiness and the three CEECs in cluster III show high similarity with cluster II, that is Estonia and Slovakia, mainly due to low public debt and similar interest rates.

Thus, overall applying cluster analysis to Maastricht Criteria, yields the following pattern:

- **the core** - Germany, France, Austria and Netherlands, with Ireland within reach but still not coping with the public debt criteria and to lesser extent Belgium generally stable with the non-fiscal criteria, but with extremely excessive debt;
- **southern periphery** - not entirely homogenous - Portugal, Spain, Italy joined by

MC	2 first components	
time	11 y.	5 y.
1^{st} var. explained	48.57	49.53
2^{nd} var. explained	34.11	22.73
Bartlett's test $p - value$		
n=2	.00	.05
n=3	.08	.08

Table 10: MC Analysis - two first principal components: percentage of variance explained by each component and dimensionality test.

Hungary with some resemblance to Greece - generally not converging to meet the entry prerequisites. Poland although approaching the Maastricht Criteria marginal values is still outside, and in many ways resembles this periphery;

- **the north/east periphery** of leading qualifiers: Finland with the Baltic States together with Czech Republic and Slovakia of which all but the last steadily qualify according to the Maastricht criteria (in this sense being more 'ready to enter' than the actual core). Slovakia although within reach of meeting the requirements, seems not to follow a steady convergence path, but rather to be fairly unstable;

- **Slovenia** starting off closer than most CEECs, but not improving especially in terms of inflation;

The PCs analysis results displayed in Table 10 (more details in Appendix B), are sufficient to allow focusing on the first two components, as in both cases they cumulatively explained over 70% of the variance in our sample and meeting other previously stated requirements. Bartlett's test suggests 2 dimensions of the data at 95% confidence level in both cases.

The rules of construction of the Readiness Index are exactly the same as in the case of

	MC 11	MC 5		MC 11	MC 5
Austria	1.48*	1.92*	Czech Rep.	1.88*	1.50*
Belgium	2.40	2.42	Estonia	2.76	0.95*
Finland	1.27*	2.06*	Hungary	4.32	2.70
France	1.14*	1.68*	Latvia	3.36	1.78*
Germany	1.12*	1.78*	Lithuania	3.98	1.06*
Greece	4.11	4.05	Poland	3.64	2.48
Ireland	1.66*	1.68*	Slovakia	1.65*	1.39*
Italy	3.06	3.38	Slovenia	2.56	2.46
Netherlands	1.81*	1.93*	Dummy(0)	0.00*	0.00*
Portugal	2.51	2.86	Dummy(MC)	2.21	2.21
Spain	1.96*	2.49			

Table 11: Maastricht Criteria Readiness Index - rescaled for equal Dummy(MC) values at both dates. Columns "MC 11" and " MC 5" denote values at, respectively, 11- and 5-years prior to entry date. * indicates values lower than cut-off point Dummy(MC)

the OCA Suitability Index, thus also the weaknesses are similar. *Dummy*(0) acts as the reference - identically to Germany in the previous analysis. Additionally, the *Dummy*(*MC*) serves as a cut-off value, but rather one way. More precisely, due to the fact of reducing the dimensions to 1, we can only be certain that a value of the index above the one of *Dummy*(*MC*) means problems with complying with the criteria. In the opposite case, when the value is smaller, this does not necessarily mean meeting the requirements, solely that a country is close to fulfilling them - usually the closer, the smaller the index, but it need not be so in every case. Moreover, the MC 5 years index is re-scaled, for the *Dummy*(*MC*) values to be equal - in order to facilitate comparison.

Comparing the indexes over the 6 years yields the following pattern:

- **diverging** - Austria, Finland, France and Germany generally qualifying and Italy, Portugal and Spain not qualifying.
- **converging** - all CEECs with the exception of Slovenia, of which only Hungary and Poland do not qualify.
- **stable** - of which Ireland and Netherlands close to fulfillment, Slovenia, Belgium and

Greece not fulfilling, albeit for different reasons.

The significant stabilization of the CEECs resulted in substantial convergence towards fulfilling the requirements. Within 5 years before the EMU accession most of these countries seem well capable of meeting the entry conditions, the leader being Estonia. Generally the Baltic States, Czech Republic and less stably Slovakia persist in nominal convergence and would seem to have less trouble to qualify than most EU members did. As for the others, the main obstacles are remaining budget deficits and marginally high public debt levels in Poland and Hungary (though still below the 60% of GDP requirement), as well as high inflation in Slovenia. Overall, however, the CEECs do not seem to perform worse than any of the southern EU states. This, suggests that all should be capable of qualifying, though for the three laggards it may require a large effort - whether it will be fiscal contraction in Poland and Hungary or an effort to bring down price inflation in Slovenia. The fact that as far as 5 years before entry, the CEECs seem more ready, may also be a sort of signalling. Being relatively young, developing economies they are generally regarded as less credible and stable. In light of the two facts: (1) the still uncertain result of the ongoing debate whether the enlargement countries should be allowed to enter the EMU as quickly as possible and whether it is optimal from the current members point of view, and (2) assuming significant gains assumed to come from adopting the common currency for the CEECs, and the will to materialize them as fast as possible, the transition economies may be more determined to show(signal) that they are in fact ready for the Eurozone.

As can be seen in Figures 5 and 7, there are clear differences, between the convergence of countries according to Readiness and Suitability Indexes. The 'nominal' structure is not concentric, though does exhibit a strengthening southern periphery. The northern countries, found peripheral in terms of OCA criteria, look most ready in terms of Maastricht

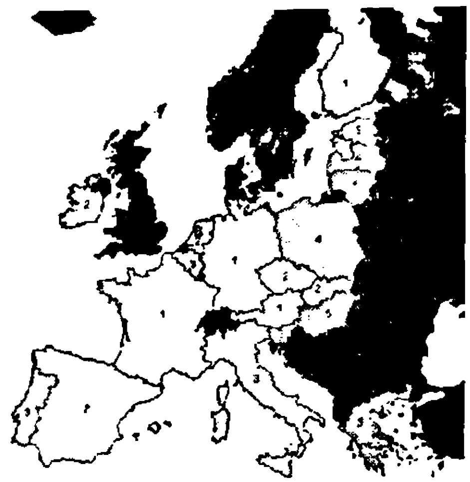

Figure 5: EMU candidates 11 years prior to (potential) entry date - MC Readiness Index. Lower number (lighter color) indicates 'closer' in terms of fulfilling Maastricht criteria. Black - not classified.

requirements. To put it briefly, perhaps not part of the Optimum Currency Area, as defined by theory, nevertheless they should need less effort to fulfil the criteria and qualify. As for the 'core' countries, most are within reach of qualifying, though despite managing to reduce government debt, many remain in excess of the Maastricht requirement.

The Baltic States converge to the northern periphery, eventually outperforming it in terms of Maastricht criteria. The Czech Republic and Slovakia, are, in terms of our index, ready 11 years before accession, and confirm this performance 5 years before. Over the 6 years Poland and Slovenia, but especially Hungary join the southern periphery in terms of the variables used. However, starting from a more unsuitable position Hungary and Poland steadily converge towards fulfilling the criteria, whereas the southern EU members show no such sign within 5 years of EMU membership.

The plot according to the two first principal components, 5 years prior to accession is presented in Figure 8, and should be confronted with Figure 6. Confirming the previous findings, even the core EU countries perform somewhat poorer in terms of nominal criteria than some of the CEECs. Spain, Hungary and Portugal show high similarity, and together

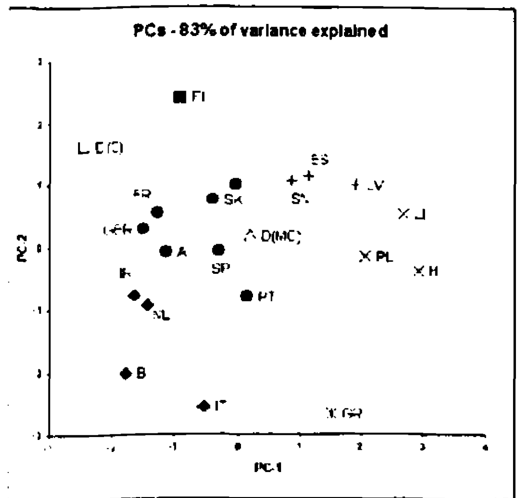

Figure 6: MC 11 years before entry - clusters and PCs compared. Objects in same hard cluster have identical icons.

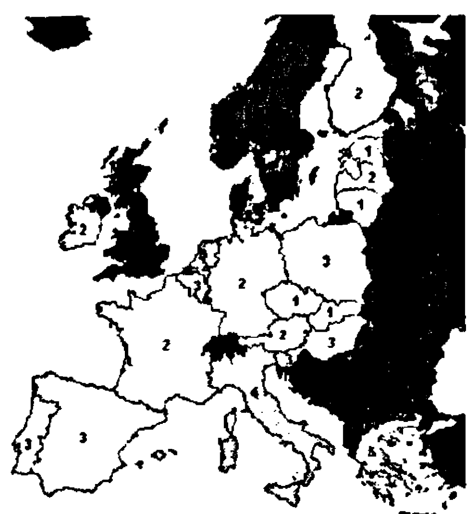

Figure 7: EMU candidates 5 years prior to (potential) entry date - MC Readiness Index. Lower number (lighter color) indicates 'closer' in terms of fulfilling Maastricht criteria. Black - not classified.

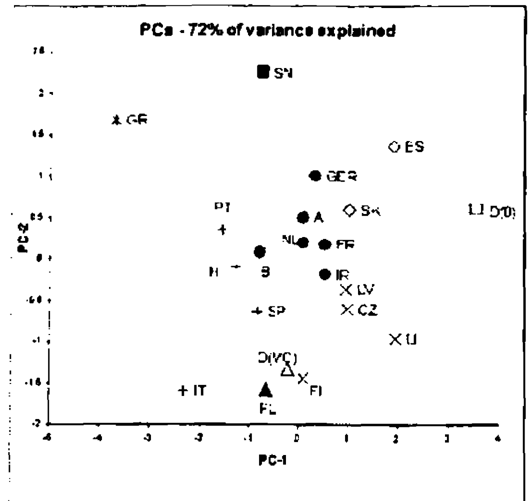

Figure 8: MC 5 years before entry - clusters and PCs compared. Objects in same hard cluster have identical icons.

with Italy and to a lesser extent Greece, seem to constitute the 'nominal' southern periphery - very similar to the 'real' one. Slovenia is also fairly peripheral nominally, but shows less resemblance to the others. The previously found 'real' core and northern periphery countries seem well capable of fulfilling the criteria, especially if we consider that de facto the debt criteria was treated lightly. Poland is not qualifying, but showing significant convergence, and sharing some of the features of the southern periphery. Summing up, in nominal convergence analysis we find:

- **the core** - consisting of Germany, France, Austria, Belgium and Netherlands, joined by Czech Republic and Slovakia, together with **the northern periphery** - Ireland, Baltic States and Finland which actually seem to have less problem to meet the MC;
- **the southern periphery** - consisting of Italy, Spain, Portugal, Hungary to a lesser extent Greece and Slovenia, but also to some extent Poland which is close to MC cut-off values;

It is also worth noting that the findings confirm both nominal and real similarities between

Czech Republic and Slovakia, and also between the Baltic States - which should not be surprising, as these economies not only show resemblance in many fields, but also, not so long ago formed parts of other countries and currency unions.

I.6.3 Results Compared

The comparison with previous work is not straight forward - there have not been many noticeable attempts to judge suitability and readiness of CEECs for the EMU relative to current members. Nevertheless, the results of this seem fairly in line with previous attempts of operationalizing nominal and real convergence for these groups of countries separately. Artis and Zhang (2002) find a similar pattern throughout the current EMU members, and the core-periphery terminology within the EMU candidates has been actually adopted from this paper. In terms of OCA criteria, applied within a shorter period before the EMU, they discover a similar pattern of a core composed of Germany (by default), France, Netherlands, Belgium and Austria, and a southern periphery - Portugal, Spain, Italy and also Greece. The northern periphery found is Finland and Ireland together with Denmark, Sweden and the U.K. which are not part of interest in this paper. Including solely EU countries allows the authors to use better data, especially business cycles correlation and labor market flexibility measures, but the similarity of their results yield support to the above findings. In terms of the Maastricht criteria, their result is slightly different. Overall, however the most appropriate period to relate to our results is the analysis conducted by Artis and Zhang (2002) for 1990-97, as the others (1995-97 and 1997) do not even overlap our sample years. This over-time average analysis is a slightly different approach, but generally they find Germany, France, Austria, Belgium, the Netherlands and Ireland in the core, Spain Portugal and Italy together with Finland in the periphery and Greece as an outlier. The fact that above results are similar but stronger than the ones in our paper can be attributed to the fact that our data set is bigger, much more variant and diverse, and in fact of poorer quality. A cluster analysis of CEECs in search for nominal and real convergence is conducted by Boreiko (2003). Though, an over time average, the results for periods 1998-2001 and 2001 can be compared to the above analysis. Firstly, in terms of real convergence the author finds Czech Republic, Estonia, Hungary and Slovenia as the best performers, sometimes joined by Slovakia. The analysis above confirms the Czech Republic and Slovakia as most suitable according to the OCA, followed by Estonia and Hungary, but fails to find Slovenia in the optimal group, as it exhibits excessive inflation and low labor market flexibility (a variable not used by Boreiko, 2003). As for the Maastricht criteria, Boreiko (2003) finds the Baltic States and Slovenia as the best performing. Our analysis confirms the Baltic States as undoubted leaders, but are joined by Czech Republic and Slovakia instead. It must be noted that we base upon more recent data, which recognizes its closeness, but fails to find Slovenia a leader, mainly due to persistent higher inflation and interest rates. These two criteria cause Slovenia not to converge, but overall it is placed very close, and in fact closer than many current EMU members, to fulfilling the entry requirements. Despite finding Slovakia a good performer in the last period (i.e. 2002) we do notice its unstable path towards the nominal criteria, therefore do not claim its readiness strongly. As for the Czech Republic, it shows persistent convergence, however still maintains an excessive budget deficit. As scoring high on the other criteria, this single violation, seems to matter less in our PC analysis, though consistently with Boreiko (2003) it is ranked lower then the Baltic States. Overall, the results are fairly similar, but it noted that the differences in exact results may arise to different methodology, time horizon and data set.

As for the verification of the indexes created in this paper, we refer to Bayoumi and

Eichengreen (1997), who use exchange rate deviation predictions based upon the estimated historical relation between this variable and standard deviation in real output difference, sum of the absolute differences in the shares of agricultural, mineral, and manufacturing trade in total merchandise trade, the mean of the ratio of bilateral exports to domestic GDP and the mean of the GDP, all relative to Germany, in order to develop an OCA index and rank countries. Therefore, using the following bilateral nominal exchange rate equation:

$$SD(e_{ij}) = \alpha + \beta_1 SD(y_i - y_j) + \beta_2 DISSIMij + \beta_3 TRADE_{ij} + \beta_4 SIZE_{ij}, \qquad (10)$$

on extrapolated independent variables, the authors associate low need for exchange rate deviations with high suitability in terms of OCA criteria. Table 12 displays the comparison between our OCA Suitability Index and an OCA index from Bayoumi and Eichengreen (1997). Although the correlation coefficients do not seem outstandingly high, they can be seen as supportive: firstly the groups of countries found most and least suitable are very similar, and secondly the fact of the actual values correlated with a coefficient above 0.40, despite a use of a very different approach, and not even exactly the same years, is actually encouraging. Overall, the results are consistent with previous findings, but contribute towards a more rigorous and informative assessment of both real and nominal performance of CEECs when approaching EMU membership.

I.7 Conclusions

The comparative statics exercise performed above was intended to find out how Central European EU candidate countries fit into the partitions believed to exist among EMU members. It aimed to explain the convergence paths towards Eurozone accession and assess

	OCA Index (1998)	B&E OCA Index(1997)	OCA Index (1994)	B&E OCA Index (1995)
Correlation (only EMU)	.48		.42	
Most suitable	Netherlands Austria France Belgium	Netherlands Belgium Austria	France Belgium Netherlands Austria	Netherlands Belgium Austria
Least suitable	Finland Greece	Finland Spain	Finland Greece	Finland Spain

Table 12: OCA Index comparison - with Bayoumi and Eichengreen (1997).

the relative suitability and readiness of these states for adopting the euro. The purpose is to yield insight on possible future inhomogeneities and policy pressures in the future union, potential gains and losses from joining and problems associated with fulfilling the entrance criteria as well as complying with the SGP or any other stability agreement that may replace it. Certainly, limitations to the interpretation of the results exist. Among them, the sheer fact that countries not seeming suitable for a common currency, may actually profit most from joining it. Secondly, the data period available for analysis is short, especially for measuring business cycle correlation - not much can be done about this. Thirdly, if we believe the story of the endogeneity of OCA theory, than despite the fact that the methodology used seems to reduce this problem significantly, we may be more reserved to trust the variables used. Nevertheless, the exercise seems interesting and worth the trade-off.

Summarizing, in this paper we find that the CEECs exhibit quite strong convergence towards both fulfilling the nominal requirements as well as to being suitable for the European currency area. The transition economies seem to blend in well with the existing core-periphery partition of the EMU members. When assessing suitability according to

OCA criteria the Central European states fit nicely in the concentric, geographical pattern that can be observed within 5 years before entry. The leaders in real convergence are the Czech Republic, Slovakia and Estonia, which become increasingly like the core. In nominal convergence, the Baltic States converge rapidly, to eventually outperform most EU members, similarly to Czech Republic, which starts of from a more stable economy and Slovakia. However in case of the latter, the convergence is unstable during the period. Hungary is found strongly present in the southern periphery in both real and nominal terms, though its convergence path suggests it may move towards the core, at least in terms of MC readiness. Slovenia starts off from a relatively privileged position in terms of both sets of the criteria, but shows little convergence, similarly to the southern periphery. Poland is converging in terms of OCA criteria, but less in Maastricht criteria and shows some similarity with the southern periphery, especially concerning the fiscal policy stance. Latvia and Lithuania, leading in nominal convergence, in terms of real variables drift strongly towards neighboring Finland and the northern periphery in general. It must be emphasized that at the start of the analysis most CEECs are certainly less prepared and suitable than current EMU members were. However, within the 6 years examined, they become more suitable and ready then southern European countries were upon 5 years before Eurozone entry. In some cases, especially nominal criteria, they manage to outperform current members. Thus, according to the analysis conducted above, the prime candidates suitable for the EMU, that should not have problems being ready to satisfy the Maastricht requirements can be expected to be Estonia and Czech Republic, and less confidently Slovakia. These states converge towards the strict core. Hungary, Slovenia and Poland will require more effort in order to comply with entry conditions, but if successful, the first two, Slovenia especially, should be joining the southern periphery, while Poland should form part of the core. The remaining Baltic

States - Lithuania and Latvia join Finland and Ireland in the northern periphery but should not have significant problems in qualifying for the EMU.

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Appendix A: Data Sources and Descriptions

Nominal Convergence - MC:

1. Budget deficit as % of GDP (Annual) - EMU-10 (ex. Greece) 1990-1994 - IFS (IMF), - 1988-1990 EIU Country Reports - Greece 1990-1996 - Economist Intelligence Unit Country Reports - CEECs 1996-2002 - DB Research;
2. Government Debt as % of GDP (Annual) - EMU-10 (ex. Greece) 1988-1994 - IFS (IMF) - Greece 1990-1996 - Economist Intelligence Unit Country Reports - CEECs 1996-2002 - DB Research;
3. Nominal Exchange Rate Volatility a. against ECU till 1999, against euro 1999-2002 b. monthly data: - all figures from IFS c. 2 year moving intervals, ending on the year reported d. ln(NERt)-ln(NERt-1) e. STDDEV(.)*100;
4. Interest Rate (Annual) a. end of year market lending rate - EMU-10 (ex. Austria), CEEC (ex. 2002) - WDI (WB) - Austria, CEEC(2002) - National Statistics Offices
5. Inflation - CPI annual % change, IFS Dummy(0) - all variables set equal to 0.

Dummy(MC) - all variables set to marginally fulfilling Maastricht Treaty Criteria requirements, i.e.: Fiscal Deficit = -3% of GDP, etc.

Real Convergence - OCA:

1. Business Cycles Correlation: Industry Production Index, Monthly - DataStream a. 8 years, pair wise against Germany: - EMU-10 1986-1994 - Greece 1988-1996 - CEECs 1994-2002 b. rebased at initial year = 100 c. smoothed using HP filter lambda = 14400 d. correlation reported e. TIME INVARIANT
2. Real Exchange Rate Volatility a. against Germany,
b. ER = NER(local/DM)*PPI(local)/PPI(GER), c. sources: - PPI - EMU-11 - IFS, CEECs - DataStream - Nominal Exchange Rate - IFS d. 2year moving intervals, ending on the year reported. e. ln(ERt)-ln(ERt-1) f. STDDEV(.)*100
3. Labor Market Flexibility a. figures -WB Doing Business 2004 Report b. TIME INVARIANT c. aggregated index: - duration and no. of procedures required to setup business - cost and minimum capital required to setup business (%GDP) - quintile ranking 1-5 (1-most flexible) d. data for Estonia missing - proxied by average Lithuania & Latvia
4. Trade Integration with EMU a. (Import from EMU cif + Export to EMU fob)/ (Import total cif + Export total fob) b. World Trade Analyzer - figures for 2002 not available previous year used.
5. Inflation - see nominal convergence.

All variables standardized to have mean=0 and variance=1.

Appendix B: Principal Components - Table with Results

Nominal Convergence - Maastricht Treaty Criteria. Evaluation 11 years prior to EMU accession date. Principal Components Analysis						Relative weights of vars.		Evaluation 5 years prior to EMU accession date					Relative weights of vars.	
Variable	1^{st}	2^{nd}	3^{rd}	4^{th}	5^{th}	$1^{st}PC$	$2^{nd}PC$	1^{st}	2^{nd}	3^{rd}	4^{th}	5^{th}	$1^{st}PC$	$2^{nd}PC$
Deficit	-0.12	-0.13	0.51	0.59	0.6	0.06	0.06	0.48	-0.47	-0.17	-0.53	-0.48	0.24	0.24
Debt	0.7	-0.68	0.15	-0.15	0.01	0.38	0.34	0.35	-0.16	-0.74	0.25	0.49	0.17	0.08
Nominal EX	-0.2	-0.42	-0.79	0.19	0.35	0.11	0.2	0.35	-0.57	0.58	0.39	0.26	0.18	0.29
Interest Rate	-0.66	-0.51	0.31	-0.47	-0.03	0.36	0.25	0.72	0.65	0.16	0.13	-0.12	0.37	0.33
Inflation	0.16	0.29	-0.03	-0.61	0.72	0.09	0.14	0.08	0.11	0.22	-0.7	0.67	0.04	0.06
% Var. Expl. by PC	48.6	34.1	9.94	5.04	2.35			49.5	22.7	16.3	7.46	3.97		
Cummulative	48.6	82.7	92.6	97.6	100			49.5	72.3	88.6	96	100		
Dim. (Barlett)	n=1	n=2	n=3	n=4				n=1	n=2	n=3	n=4			
χ^2	55.9	38.4	9.85	2.83				35.5	16.7	9.85	1.95			
$p-value$	0	0	0.08	0.24				0	0.05	0.08	0.38			

Real Convergence - OCA Criteria. Evaluation 11 years prior to EMU accession date. Principal Components Analysis						Relative weights of vars.		Evaluation 5 years prior to EMU accession date					Relative weights of vars.	
Variable	1^{st}	2^{nd}	3^{rd}	4^{th}	5^{th}	$1^{st}PC$	$2^{nd}PC$	1^{st}	2^{nd}	3^{rd}	4^{th}	5^{th}	$1^{st}PC$	$2^{nd}PC$
Business Cyc.	-0.26	0.56	-0.32	-0.52	0.49	0.16	0.28	-0.12	0.34	-0.55	-0.58	-0.49	0.07	0.18
Real EX	-0.1	0.29	0.68	0.42	0.51	0.06	0.14	-0.81	0.37	0.13	-0.1	0.42	0.47	0.19
Labor Mkt.	0.93	0.35	0.05	-0.11	0.01	0.57	0.17	0.32	0.82	0.41	0.17	-0.17	0.18	0.43
Trade Int.	-0.15	0.13	0.63	-0.6	-0.45	0.09	0.06	0.47	0.17	-0.23	-0.4	0.73	0.27	0.09
Inflation	-0.19	0.68	-0.18	0.42	-0.54	0.12	0.34	-0.02	0.23	-0.68	0.68	0.12	0.01	0.12
% Var. Expl. by PC	43.46	27.6	19.2	6.47	3.36			49.7	24.8	15.8	5.86	3.75		
Cumulative	43.46	71	90.2	96.6	100			49.7	74.6	90.4	96.3	100		
Dim. (Barlett)	n=1	n=2	n=3	n=4				n=1	n=2	n=3	n=4			
χ^2	33.44	22.4	13.9	1.89				36.1	19.1	10.1	0.89			
$p-value$	0	0.01	0.02	0.39				0	0.02	0.07	0.64			

Table 13: Principal components analysis - detailed results.

Chapter II
How Easily is Public Investment Cut? A Dynamic Panel Approach

Abstract

Using a dynamic panel of 19 OECD countries over 1971-2004 we analyze the movement of public investment spending. Controlling for downward trends in many countries, we focus on short-term movements and find investment a fairly rigid, albeit pro-cyclical, component of government spending as it fluctuates less than current expenditures. However, in severe fiscal restraint it is public investment that is mostly affected. We show that governments with a myopic policy horizon did not generally lower public investment, but were more inclined to cut it during fiscal consolidations. Moreover, fiscal consolidations related to the adoption of the euro did not differ in this respect.

JEL Classification Numbers: H50, H62, C23.

Keywords: *Public Investment, Fiscal Adjustment, Political Horizon, EMU, Dynamic Panel Estimation.*

II.1 Introduction

The last half century has seen a notable fall in real government investment expenditures in most industrialized countries. Though government spending was generally increasing in real terms, public investment levels as a share of Gross Domestic Product fell sharply in many of the most developed countries. While undoubtedly this long-term trend decline is partly a matter of changes in demand for public investment it is interesting to ask how much movement in public investment is due to short-term fiscal and political factors.

The objective of this paper is to assess how strongly public spending cuts affect government investment and whether governments with a short horizon, politically weak or facing elections are more prone to cut investment instead of current expenditure, thus breaching the 'golden rule' of investment. Furthermore we ask whether public investment exhibits the features of a Keynesian counter-cyclical tool and whether fiscal requirements associated with the introduction of the euro had an effect on investment in countries involved.

In order to verify the hypotheses, we use several different definitions of fiscal adjustments and of political horizons, as well as different specifications including a selection of cyclical, fiscal and political variables on a set of 19 OECD countries in the years 1971-2004. We apply a set of estimation techniques to a dynamic panel with a lagged dependent variable (LDV) in order to verify a series of hypotheses. In order to be more confident in interpreting the results we perform a short set of Monte Carlo (MC) simulations assessing the performance of the estimation techniques in a panel setup dimensionally similar to the one available. As the Least Squares Dummy Variable (LSDV) and Arellano and Bond (1991) General Method of Moments (AB-GMM) estimators seem to behave well, we use them for the purpose of verifying the hypotheses.

We find that, in the short term, public investment tends to fluctuate far less than total government spending; however it is highly prone to cuts during a major fiscal adjustment. Moreover, we find that during a budgetary consolidation, a myopic government is significantly more willing to pursue investment cuts than a government with a longer policy horizon. On the other hand, we fail to find any noticeable effect of upcoming elections as such on public investment nor any change of public investment behavior due to monetary integration in the EU and the related fiscal constraints such as the Maastricht Treaty entry criteria or the Stability and Growth Pact. The latter part of the analysis may be flawed by the shortness of the sample, but generally though we do find that a significant number of fiscal consolidations took place in the EU countries in the second half of the 1990s, these did not seem to differ, at least in the effect on public investment, from other fiscal consolidations in our sample. Finally we find that public investment, though following a negative trend in many of the OECD countries, was strongly pro-cyclical in the short-term.

The remainder of this paper is organized as follows. Section II.2 reviews the literature on the determinants of government investment, discussing previous approaches and results. Section II.3 presents the estimation setup and data used in the exercise, while the following Section II.4 presents the results, their robustness to changes in specification and focuses on their interpretation and discussion. Finally, Section II.5 concludes, while Appendix C contains results of a Monte Carlo simulation exercise which provides insight on the appropriate estimation method.

II.2 Explaining Public Investment Changes

II.2.1 Public investment in OECD countries

Public Investment in most industrialized countries has seen a decline since the 1960s and 1970s both as a share of GDP and of total government expenditure. In the first case, presented in Figure 1, the decrease is clearly visible for EMU countries excluding Ireland, Greece, Spain and Portugal; for the EU countries outside the EMU; for Norway, Australia, Canada, New Zealand and the US. Public investment levels measured as gross fixed capital formation of the general government where usually between 4-6% of GDP in the 1960s and 1970s while reaching levels around 1-3% with a notable exception of the 1980s' EU accession countries, Ireland and Japan.

Notably, the fall in government investment is even more profound if we look at the share in total government spending. In Figure 2 we see that government investment spending as share of total government disbursements fell gradually reaching less than half the levels of the 1960s in all of the countries apart Ireland, Portugal and Spain. This fall is in fact visible in the gross values of investment, while with the available data we cannot say much about the net values. There is no reason to think that the depreciation rate of public capital should have remained the same through-out the past 40 years, but there is not much work done in this area. Attempts to estimate the real net public capital stock by Kamps (2004) do not show such a severe a drop in the stock, thus perhaps backing some theories of the falling marginal infrastructure demand.

As for the effect of public investment on growth, empirical estimates seem to vary quite widely. A detailed summary of the literature, found in De Haan and Romp (2005), points to the fact that although the empirical results differ in magnitude, significance and even direction, generally public capital stock seems to have a positive effect on growth. The link does appear fragile, depending on the estimation method, data and period used, the exact type of investment (transport, education, health etc.) and the fact that many of the studies can not properly account for the net effect of public investment spending, which arises from the fact that as with any other type of public spending it takes resources away from other

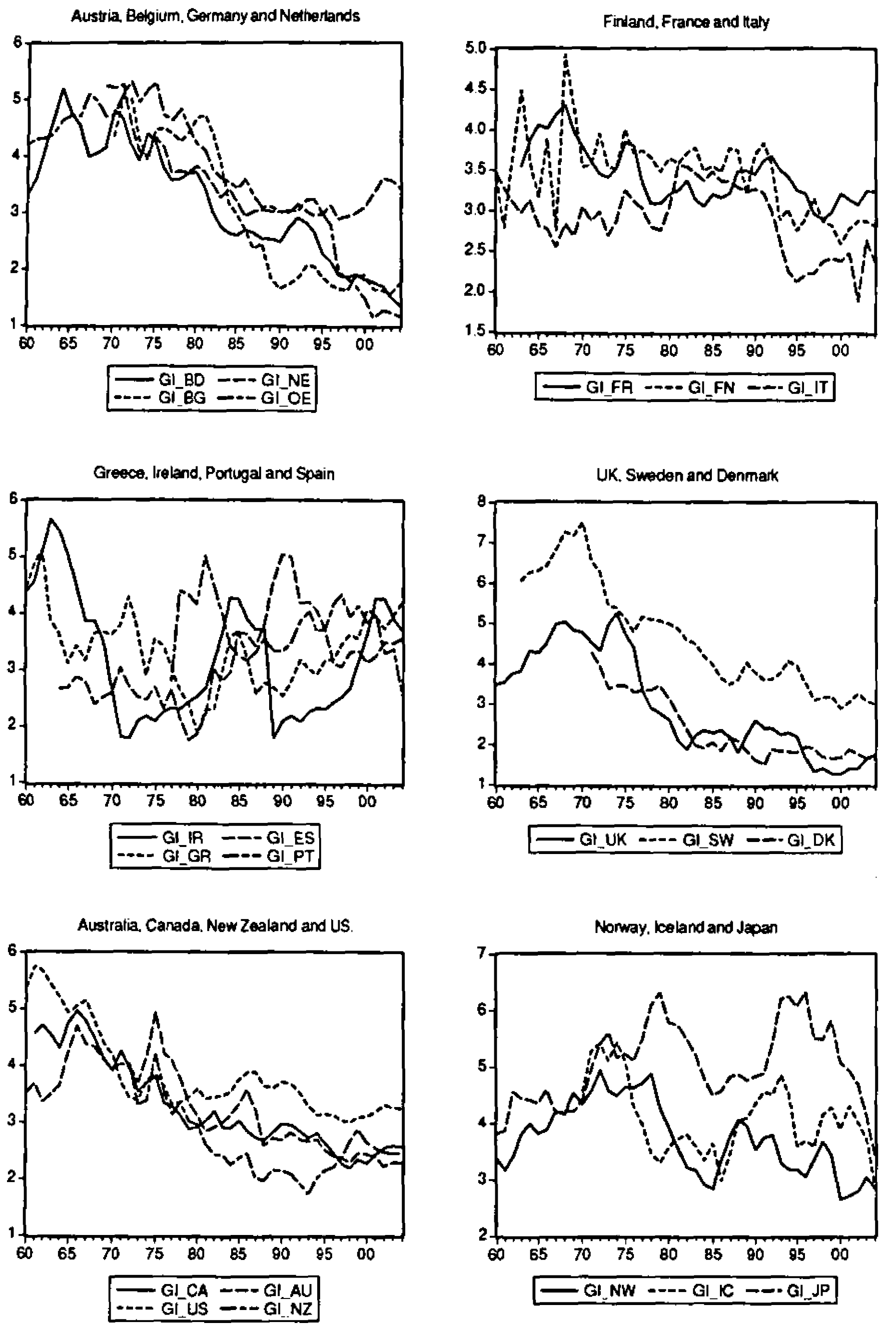

Figure 1: Government Investment as % of GDP.

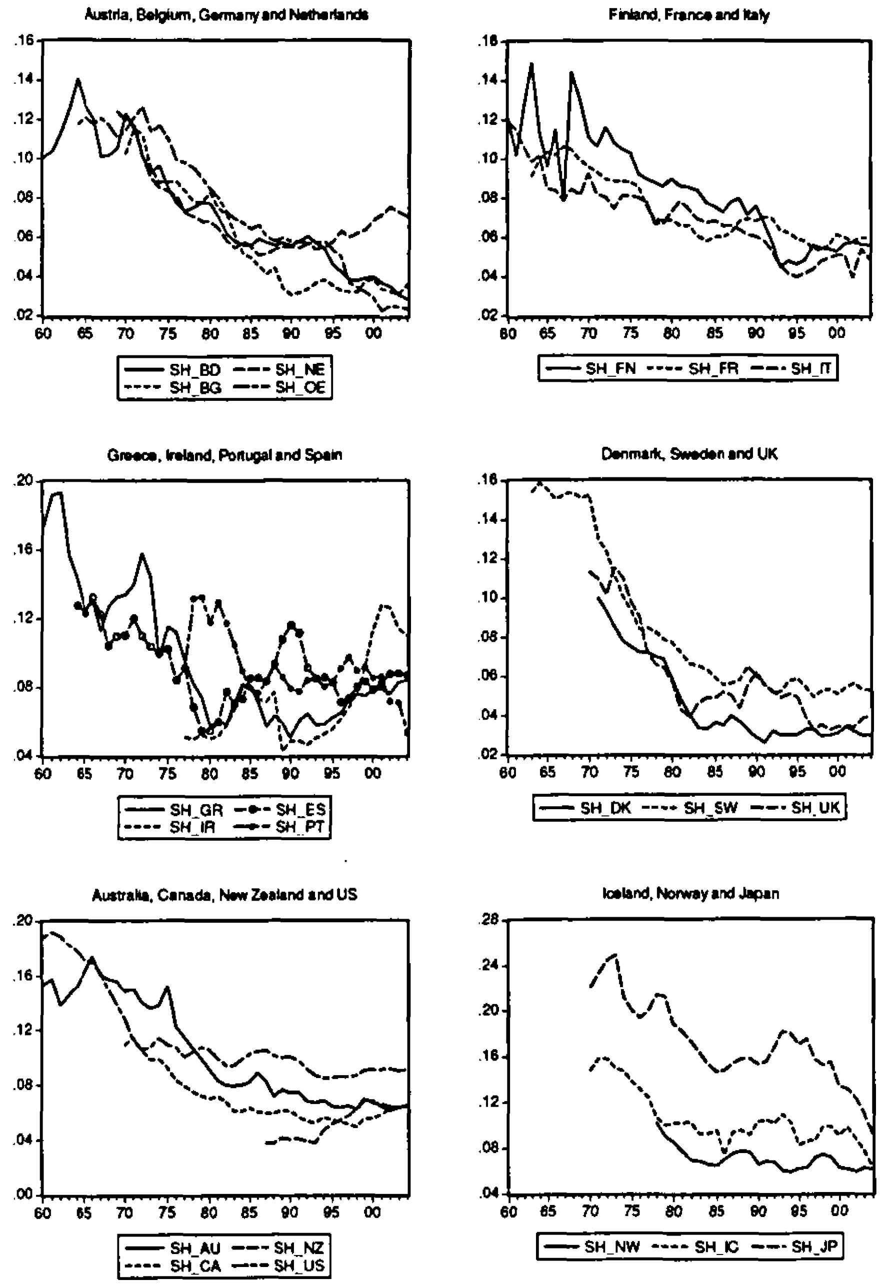

Figure 2: Government Investment as share of Total Government Disbursements.

activities. Overall, the impact of public investment on growth seems to be much lower than the initial estimates of Aschauer (1989), which assessed the productivity of public capital as significantly higher than that of private capital. Nevertheless, practically all studies agree that some levels of public capital are a necessary prerequisite for growth.

In our paper we abstract from evaluating issues such as the causes and consequences of the long-term drop in public capital expenditure, the growth effects, changes in demand for public infrastructure, the possibility that certain types of investment have reached some saturation point, the potential problem of undersupply of public capital and the indirect effect of privatization of certain sectors. The trend decline, if found in some countries, may be a demand phenomenon or a consequence of changes in the depreciation rate, and if so should not have any adverse effects on the level of public capital and moreover on economic growth. On the other hand, if important projects are foregone because of incidental cuts in public expenditure caused by political or other disturbances, the negative effect on development and growth may be noticeable. Hence we focus on the reaction of public investment spending to fiscal adjustments and electoral cycles, yet in order to isolate the effects, take into account the forces determining public investment levels and the different trends in such spending.

II.2.2 Hypotheses

In this paper we attempt to test two main hypotheses:

- (H1) First of all, we ask the question whether public investment tends to be the least rigid component of government spending i.e. does it fluctuate more than other categories of government spending, being more likely to be subject to cuts and increases than current spending?

We define rigidity, as short-term persistence with respect to total public spending movements - that is, for example, during a total spending change of $x\%$, a more rigid component would usually change by less than $x\%$ while a less rigid component would change by more than $x\%$. We are especially interested how strongly do large fiscal consolidation efforts fall upon public investment. That is, whether the above relationship changes when government finances undergo a serious cut - signalling some kind of asymmetry upon the incidence of consolidation.

- (H2) Second, we ask whether governments with a shorter policy horizon tend to be more inclined to cut investment when cutting public expenditures than governments that are more likely to stay in power longer.

This hypothesis originates from the idea that if elections are more frequent and the political situation less stable, governments may be trying to "buy" votes by preserving current spending, which is arguably more visible to the voters, and thus letting public investment bear the burden of fiscal cuts.

The two questions come from the fact that very little empirical research has been done to assess the effect of fiscal adjustments and the political horizon on government investment. On one hand, Oxley and Martin (1991) argue the existence of a *"political reality that it is easier to cut-back or postpone investment spending than it is to cut current spending"* backed by Roubini and Sachs (1989), who claim that *"in period of restrictive fiscal policies and fiscal consolidation capital expenditures were the first to be reduced (often drastically) given that they were the least rigid component of expenditures"*. Intuitive explanations for the first view argue that public investment is less visible than current spending to the voter, at least in the short term, thus a myopic government, when cutting spending should be inclined to cut public investment. The governments' chance of reelection would be more adversely affected if current spending was cut. Therefore, we implicitly assume there exists a trade-off between the 'short-term' policy, as part of which current expenditure is preserved versus investment spending, aiming to please voters and thus at an election success, and 'long- term' policy goals like fiscal consolidation or infrastructure and thus growth prospects. The existence of a temptation to forgo the second in favor of the first because of political weakness and instability is based on the notion that effects of current spending cuts are more instantaneous and direct for individuals, while investment cuts have an effect more prolonged in time, more dispersed and thus individuals do not have a certainty it will affect them. For the sake of clarity, take a simple example - a government can promise to raise teachers wages in the following year, or can promise to build a school. If the government should decide to go back on its promise, postpone its realization or perhaps spread out its realization in a longer period of time, we can imagine different reactions to such a decision in the two different scenarios. Intuitively, in the first scenario (teachers wages) the effect will be felt immediately and clearly by a very specific group, while in the second case (building a school) the effect would not be as instant, not so clear and the affected group not as certain. Therefore in the first case we can expect much fiercer opposition as incentives to lobby are much higher. On the other hand Aubin, Berdot, Goyeau and Lafay (1988) claim that *"investment expenditure is usually more evident than consumption expenditure"* and that it is more acceptable for the wide public. If we believe this than the previous story should not hold.

The second view is often labelled the "golden rule of investment". Discussed for instance in Blanchard and Giavazzi (2004), it originates from the idea that potential benefits from public investment projects will be reaped in the more or less distant future, thus running a deficit and borrowing for investment is justified as sort of an inter-temporal deal - the future beneficiaries will also share the costs.

In addition to our principal line of investigation we try to verify two other questions:

- (H3) Is public investment used as a counter cyclical tool, i.e. can we *ex post* claim that public investment fluctuations followed a counter or rather a pro cyclical pattern?

- (H4) Has the formation of the EMU had a significant effect on public investment in its member countries.

We do the latter by asking whether the burden of budget limits such as the Maastricht Treaty requirements and the SGP affected public investment levels. If we do find evidence for the Pacts' influence on investment, this may be supportive of the literature on SGP reforms, and amendments to account for public investment, as in Blanchard and Giavazzi (2004). This effect may be of special importance in light of the future EMU enlargement, where the CEECs, because of their currently inferior levels of infrastructure may require higher levels of gross investment - as did the southern countries that joined the EU in the 1980s or as did the old members in the 1970s.

II.2.3 Empirical work on public investment

There are two main strands of empirical literature related to our analysis. The first type, literature on the determinants of public investment, ranges from time-series estimations summarized in more detail in Lybeck (1988) or Sturm (1998) to panel data estimates as in De Haan, Sturm and Sikken (1996) and Turrini (2004). A selection of important contributions have been presented in Table 1. We divide the variables used for explaining public investment into four main categories.

Firstly, there are the lagged values of the dependent variable. They are used to capture some persistence dynamics that may be present in public investment figures, arising from features such as the continuity of certain projects or policies.

The second category can be labelled as *Economic Structural Variables* and includes real GDP or its growth rate, GDP per capita, the output gap, unemployment and inflation. These variables are used to account for cyclical factors and test the counter-cyclicality of public investment, that is whether it is used as a stabilization tool. GDP in levels is used to test for Wagners' Law of increasing share of government spending in GDP. Henrekson (1988) suggests that government investment spending in Sweden may have been counter-cyclical but admits that it seems doubtful that government growth theories used in his estimation are applicable to explaining public investment. Aubin et al. (1988) find evidence in favor of the Keynesian stabilization role of French government investment using unemployment rates. De Haan et al. (1996) using the GDP growth rate find evidence that public investment is counter-cyclical, to an extent higher than other government outlays. On the other hand Turrini (2004) uses the output gap to find weak evidence of pro-cyclicality, albeit increasing slightly after the second phase of the EMU. Other variables in this category are the ratio of mean to median income, as a measure of inequality, the real interest rate proxying for the alternative costs of investment and found to have a significant negative effect by Aubin et al. (1988). De Haan et al. (1996) ask the question whether private investment is a substitute or a complement of public investment and suggest there is evidence of complementarity. Among other variables we have government investment inflation, urbanization, trade and current account variables and income distribution.

Fiscal Variables contain various measures of deficit; government receipts and spending; and public debt levels. They are included to take account of the general situation in the public finances and to test more specific hypotheses such as that in periods of restraint public investment is more likely to be cut than other government spending. Most of these variables are found significant if included; worth mentioning are the strongly negative effect of public debt on investment found by Turrini (2004), negative effects of lagged current expenditure and positive of tax revenues in the same work and the strong negative effect of the fiscal stringency dummy in De Haan et al. (1996).

Socio-Political Variables like the degree of unionization, coalition strength, political scene orientation, electoral cycle, public sector employment, political stability etc. are included to test whether political pressure, degree of centralization, government power and orientation, election vicinity etc. impede public investment levels. Henrekson (1988) finds a negative significant effect of weakness of coalition on public investment for Sweden. In the French case, Aubin et al. (1988) discover that central governments tended to raise investment before national elections, while local governments tended to cut investment and increase consumption before local-level elections. De Haan et al. (1996) do not find support for the importance of any of the political variables in yearly data, but manage to find a significant negative effect of the frequency of government changes on both government investment as share of GDP and as share of government outlays on three-year averages.

Finally, we have variables that may be hard to attribute to the other categories, but are included in order to test specific hypotheses. These are for instance EMU dummies and interaction variables, as in Turrini (2004), which serve for testing whether public investment was affected by the fiscal stringency imposed during preparation for the EMU and the budget limit of the SGP.

The second type of empirical analysis relevant to our work is the literature on the composition and characteristics of fiscal adjustments. It tends to focus on how much of the fiscal cut is borne by the reduction of public investment. Balassone and Franco (2000) in a com-

Paper	Countries and Period	Public Investment	Explanatory Variables	Estimation Method
Henrekson (1988)	Sweden 1950-1984	Government Investment deflated by GDP price deflator	Demand: *Urbanization*, Real GDP, Dependency Rate, Income Distr, Unionization, *Deficit*; Supply: *Public Sector*, Unemployment, *Coalition*, Tax Centralization	ML
Aubin et al. (1988)	France, 1961-83	General, Central and Local Gov. GFCF at Cur. Prices - growth rates	*Population growth*, Δ*Unemp*, Δ*Real Interest*, *Local Tax to Spending growth*, *Election Dummies*	OLS
De Haan et al. (1996) revised in Sturm (1998)	19-22 OECD countries, 1980-1992	Government Investment as % of GDP, as % of Gov. Outlays	Economic: *LDV*, Δ*GDP*, *Government Inv. Inflation less GDP Inf.*, *Civil Servants growth*, *Fiscal Stringency*, Δ *Structural Deficit*, *Private Inv. Tax Centralization*; Socio-Political: Government Power, Color, Political Stability, Election year	unbalanced panel WLS, FE
De Haan et al. (1996)	19-22 OECD countries, 1980-1992, 3-year averages	Government Investment as % of GDP, as % of Gov. Outlays	Economic: LDV, *Government Inv. Inflation less GDP Inf.*, *GDP growth*, *Civil Servants growth*, *Political Stability*	unbalanced panel WLS, FE
Turrini (2004)	14 EU countries, 1970-2002	GFCF of government sector as share of potential output	*Per Capita trend real GDP*, *EMU dummy; adjusted tax revenues, lagged current expenditure, output gap, output gap interacted with EMU dummy, lagged debt gross of interest exp., lag of primary CAB, all over trend real* GDP	FE panel and FE IV panel

Italics denote variables significant in at least one specification

Table 1: Determinants of public investment - Literature summary.

parative statics exercise for EU countries in the 1980s and 1990s[1] find positive correlation between fiscal consolidation and reduction of public investment, both as a share of GDP and as a share of primary outlays. The construction of the exercise allows only for rough interpretation as it cannot establish causality, control for or disentangle the various underlying effects or provide a measure of significance. Similarly Jonakin and Stephens (1999) in a comparative statics exercise for 5 Central American economies in the period 1975-1993 find public investment strongly affected by fiscal adjustment. Alesina and Perotti (1996) analyze the composition and longer-term deficit reduction effect of fiscal adjustments in OECD countries to find that fiscal consolidation done through public investment is usually reversed in the near future, while current spending cuts have a more persistent effect. Their finding indicates that consolidation through public investment, apart from being ineffective, results in additional volatility in public finances. Gali and Perotti (2003) ask whether the signing of the Maastricht Treaty had a significant impact on public investment levels in EMU-to-be members. They use comparative statics and fiscal policy rule estimation to discover that a trend decline in public investment commenced long before, in the 1980s and took place in other OECD member states. They find some evidence of public investment being mildly pro-cyclical, but do not find a significant change of this characteristic after Maastricht. Sanz and Velasquez (2003) take a dynamic panel approach and use Arellano and Bond (1991) GMM estimates for 26 OECD countries for 1970-1997 to test the effect of government contraction on various components of public spending. They arrive at the result that a number of spending categories which can be (roughly) associated with public investment, namely housing and economic and public services are affected more-than-proportionally by a decrease in government size, while transport and communication is affected proportion-

[1] with the exception of Greece, Portugal and Spain; Luxembourg is excluded due to data unavailability.

ally. These results are not fully adaptable for our purposes, as it does not target public investment expenditures directly. For a sample of OECD countries Lane (2003) estimates individual regressions over 1960-1998, to find that government investment spending was the most pro-cyclical of the components of government spending, but this characteristic varied among countries as in the UK it was actually the most counter-cyclical component.

Summarizing, there is not an overwhelming amount of empirical research on public investment and the results do not seem very robust. The contradictory evidence on the question of the pro-cyclicality of investment spending and the fact that little support can be found for the political determinants of public investment seem to be just some of the interesting areas for further research.

II.3 Data and model setup

II.3.1 The econometric model

The model that we will estimate is generally a follow-up on the setup proposed by De Haan et al. (1996) - a panel with a lagged dependant variable and country specific intercepts augmented by country-specific trends in order to account for the long term decline in public investment:

$$\begin{aligned} PubInv_{it} = \sum_k \rho_k * PubInv_{it-k} + \sum_s \alpha^s * FisVar^s_{it} + \sum_s \beta^s * PolVar^s_{it} \\ + \sum_s \gamma^s * CycVar^s_{it} + \sum_s \delta^s * OthVar^s_{it} + \phi_i * t + \mu_i + \varepsilon_{it} \end{aligned} \tag{1}$$

where $PubInv$ is the dependent variable; $FisVar$ is the set of 'fiscal' variables, that include total government disbursements, dummies for fiscal cuts and interaction variables; $PolVar$ is the set of 'political' variables including election frequency, election dummies and interaction variables; $CycVar$ is the set of 'cyclical' variables, capturing the effect of the business cycle; $OthVar$ is the set of all other variables, for example monetary integration dummies, private investment and the real interest rate; ϕ_i is the country specific time trend coefficient; μ_i is the country specific intercept.

II.3.2 Variables

A more detailed description of the data sources and construction of the variables is in Appendix A. In our exercise we use the ratio of gross fixed capital formation of the general government to GDP as the indicator of public investment. This is the most common measure of public investment in the literature. One of the convenient features of this measure is that it excludes military and defence spending. It does however, among others, include fixed capital formation in education, health, transport and communication, community housing, social security, electricity and gas supply and general public services. We are reluctant to use estimates of the depreciation rate in order to obtain figures for net investment for a number of reasons. First of all, the figures are not available for most of the countries and years in our setup. Second and more important, these are only estimates, thus depend on very strong assumptions and would add an unnecessary source of error to our investigation. Moreover, in our analysis of the determinants of short-term fluctuations it is gross spending that is more interesting, as it is the actual tool of the government decision.

As for the basic setup it follows directly from De Haan et al. (1996) and other empirical work surveyed above, but is adapted for the purpose of testing our hypotheses and alternative specifications are explored. We can group the explanatory variables in four main categories. For each of the variables we include a small description of why we find it justified to include in the regression. We name all the variables we used, though the individual specifications of the model consist of various subsets of variables. The first is the *lagged values of the dependent variable.*

We estimate the model with a selection of *Fiscal Variables* such as *Net Lending, Primary Balance, Current Disbursements, Total Disbursements and Total Receipts.* However, for the purpose of verifying (H1) we focus on *Total Disbursements* and *Consolidation dummies*, the former are cyclically unadjusted and measured relative to real GDP, while the latter are constructed from other fiscal variables. [1] If public investment spending is a rigid component of total spending, than we should expect it to fluctuate less than total spending - the coefficient on *Total Disbursements*, call it α_{TD}, should be lower than the share of public investment in total spending, call it *share*, in which case an x % of GDP cut in total spending, controlling for other factors, would be accompanied by a $\alpha_{TD} * x$ % of GDP cut in investment spending. This would be lower than *share*x* % of GDP - the amount we would expect if the cut was to affect investment spending proportionately. As for the coefficient on the *Consolidation dummies*, call it α_{CD}, it serves for the verification of whether the previous effect changes during a major fiscal consolidation. In this case, we compare a $\alpha_{TD} * x + \alpha_{CD}$ % of GDP actual cut in public investment with a $share * x$ % of GDP if the cut was to affect investment in the same way it did total spending. Contrary to work such as Turrini (2004) we do not include public debt in our set of variables, though one might think that a large burden of debt causes a decline in public investment to be more likely. However, we are not interested in the causes of cuts in government spending or the primary balance, and the effect of debt should be picked up by the other fiscal variables - either through

[1]The working definitions of fiscal consolidation are based on the idea by Alesina and Perotti (1996). They propose labelling a period in which the cyclically adjusted primary fiscal deficit falls by at least 1.5% of GDP within a year, or by at least 1% of GDP in two consecutive years a fiscal consolidation. We base ourselves on this primary definition, using additional definitions and a wider range of variables for this purpose, such as total disbursements, primary balance and net lending. For more details check Appendix A.

the channel of total disbursements, current disbursements and primary balance, or through fiscal consolidation which, it cannot be excluded, may be induced by the debt burden.

Political Dummies are included to test hypothesis (H2), that is whether elections cause the government to sacrifice public investment for current spending. A confirmation of this statement can be found if the coefficient on *Election dummies* is negative. A negative coefficient on *Election frequency* means that a less stable political situation reduces public investment, while a negative coefficient on *Election Frequency interacted with fiscal consolidation dummy* is evidence that governments in a less stable situation prefer to consolidate through public investment.

For the third hypothesis (H3), cyclical variables like the *Unemployment* rate, the *Output gap*, *GDP growth rate* and *Inflation* are used. Testing implicitly for the stance of public investment policies with respect to the business cycle, in case of a counter-cyclical policy we can expect the coefficients on the *Output Gap*, *Inflation* and *GDP growth rate* to be negative, and the one on *Unemployment* to be positive.

As for the verification of (H4) a negative coefficient on any of the *EMU dummies* or their interaction with fiscal other variables would allow us to argue that monetary integration in Europe had an additional negative impact on public investment. This would mean that the potential consolidations required by the fiscal criteria in order to qualify for the EMU, or in order to adhere to the SGP were different than other consolidations, affecting public investment to a larger extent. Moreover, interaction variables allow us to assess the degree to which the above mentioned relationships changed because of the introduction of the euro.

Other variables include controls such as the *Real interest rate* as a proxy for the cost of investment and *Private Investment as % of GDP*, included to test whether there is a degree of substitutability between the two.

Another important issue arising when government investment is examined is whether to use values of explanatory variables from the same time periods t or the lagged t-1 values to explain the dependent variable at time t. There seems to be no straightforward answer and the empirical literature is not consistent on this matter. We decided to take the values from the same period as the dependent variable, as the decision on government investment expenditures in the budget is made together with other budgetary decisions for a given year. Moreover, we can somewhat disreard the questions of demand for public investment or its desired level in different countries as the identification of those would be cumbersome. The inclusion of the country specific trends roughly takes care of this matter and allows us to focus on our main research questions.

Finally, in order to confirm the stationary nature of our variables, of which most are expressed in percentage of GDP, we refer the reader to Table 4 in Appendix B. Overall, panel unit root tests rather consistently reject the null hypothesis of non-stationarity in the specification with individual intercepts. This allows us to estimate equation (1) in levels.

II.3.3 Estimation techniques

For each specification we use the following techniques to obtain coefficient estimates:

- OLS-Pooled regression with country specific time dummies but common intercept (OLS),
- LSDV - fixed effects panel estimator (LSDV),
- the Anderson and Hsiao (1981) first difference instrumental variable (AH-IV) with second order lags of the dependent variable in levels as instruments,
- Arellano and Bond (1991) GMM estimator (AB-GMM).

Taking the 19 OECD countries sample, over the years 1971-2004, results in an unbalanced panel of N=19, T=34.[2]

If, as we believe, the fixed effect specification is correct the OLS estimator can be expected to be inconsistent even as N and $T \rightarrow \infty$ due to the omitted variable bias,[3] however may be insightful. In this case both the LSDV and AB-GMM estimator should be consistent and unbiased. The AH-IV estimator can be expected to be similar to the previous two but is known for its lower efficiency owed to the fact of not exploiting available information and instrument weakness. We do not expect the finite sample 'Nickell bias'[4] affecting the LSDV estimator to be of a problematic order as it is a noticeable problem only in panels with a large N and small, fixed T while in our case T=34>N. Thus in the finite sample with a similar cross section and time dimension the AB-GMM should have no advantage. Obviously in a macro panel of similar N and T, both relatively small, we should be careful with the asymptotic properties, thus in Appendix C we propose a MC exercise in order to give us some insight on the behavior of the different estimators in similar sample settings. The various estimation methods are used as an illustration of some the problems arising when working with a macro-panel with a narrow cross-section and relatively long time-series.

[2]Due to the unavailability of cyclically adjusted data for the creation of fiscal cut dummies New Zealand and Iceland were excluded. Final number of observations is 575.

[3]In the case of the pooled OLS estimation the country specific effects would be assumed to be uncorrelated with the explanatory variables.

[4]Check Nickell (1981) or Arellano and Bond (1991) for more details. In short, the 'Nickell bias' in dynamic panels with LDV is a finite sample bias resulting from the presence of the country specific effect both in the LDV and in the regression equation. It can be of large and significant magnitude for a small, fixed T, but disappears as $T \rightarrow \infty$.

II.4 Results

We decided to take the approach of estimating different specifications with varying definitions of fiscal stringency, different political variables and separate EMU second and third phase dummies in order to check the robustness of our results. For the purpose of interpreting the coefficients, we report an example specification.[5] The application of all four estimation methods yields fairly similar results, displayed appropriately in Table 2 below. The chosen example specification uses total disbursements cuts as in definition 4 and the fiscal dummy and EMU dummy definition 3.[6]

As for technicalities, the AB-GMM estimator applied in the standard version cannot reject Sargan's over-identifying restrictions test null, thus the application of the technique seems valid. Due to the country set used in our analysis we decided to apply the heteroscedasticity consistent version of the AB-GMM estimator, for which the distribution of the Sargan's statistics is unknown. The Arellano-Bond test hypothesis of lack of second order autocorrelation cannot be rejected at 5% nor 10% significance level, which means our estimates should be consistent. General robustness checks with a shorter time-series dimension can be seen as supportive for our results. As for the choice of fixed versus random effects, the Hausman test rejects the null of the difference between the estimates not being systematic, what justifies the use of fixed effects.

We find the estimated coefficients on the variables of interest consistent throughout different specifications of the model, though varying noticeably in significance, sometimes in magnitude and rarely in sign as different estimation methods are applied.

[5] Appendix B provides results for an entire range of specifications. Generally, the results seem very robust to the changes in measurements of our control variables, as well as different definitions of fiscal consolidations.

[6] Check Appendix A for details of all the definitions. The idea is based on Alesina and Perotti (1996) and De Haan et al. (1996).

Variables	OLS	LSDV	AH-IV	AB-GMM
GGGFCF(-1)	.961***	.865***	.957***	.874***
	(.03)	(.03)	(.13)	(.03)
GGGFCF(-2)	-.149***	-.156***	-.180***	-.154***
	(.03)	(.03)	(.04)	(.03)
OUTGAP	.021***	.032***	.038***	.032***
	(.00)	(.00)	(.01)	(.00)
TOTDIS	-.000	.013***	.051***	.010**
	(.00)	(.00)	(.01)	(.00)
CONS_D	-.111***	-.124***	-.087**	-.127***
	(.02)	(.02)	(.03)	(.02)
PRIVINV	-.002	-.006	-.012	-.008
	(.00)	(.00)	(.01)	(.00)
RI	-.002	-.001	.005	.000
	(.00)	(.00)	(.00)	(.00)
EMU	-.045	.032	-.015	.029
	(.04)	(.04)	(.10)	(.04)
ELEC	-.018	-.019	-.022	-.016
	(.02)	(.02)	(.02)	(.02)
ELEC_FREQ	-.524***	-.459***	-.236*	-.454***
×CONS_D	(.12)	(.12)	(.13)	(.12)
Trends	(-): AU**,BD**, BG**,CN*, DK***,IT*, JP**,OE**, UK***,	(-): AU***,BD***, BG***,DK***, FN**,IT***, JP***,OE***, PT**,SW***, UK***, (+): IR***,		(-): AU***,BD***, BG***,DK***, FN*,IT***, JP***,OE***, PT**,SW***, UK***,
obs	575	575	559	559

Numbers in brackets are standard errors.
*,**,*** denote 10%, 5% and 1% significance.

Table 2: Estimation results, example specification (1) from Table 6 in Appendix B.

II.4.1 Discussion

In this section we review the insight of our estimation results on the hypotheses of interest. We use the example specification of Table 2, but exploit the fact that the overall results are fairly consistent.

As for the general interpretation, the lagged dependent variable seems to influence the current level of government investment strongly, which confirms the necessity to include it in order to account for the multi-year nature of public investment. We decided to include the second order lag, as it is highly significant, but will refrain from searching for a more intuitive interpretation of the dynamics.

Moreover, the country specific trend coefficients are generally of high significance, thus overall the inclusion of country trends in our analysis is justified and the coefficients on other variables can be interpreted somewhat as deviations from the trend. In most specifications significant negative trends are found for 11-12 of the countries, while in some specifications the trend coefficient for Ireland is significantly positive. As emphasized, these trends are included to proxy for the phenomena of generally falling public investment. They may capture issues like demand for public investment, cost changes, or some sort of saturation/decreasing returns issues. In this paper we focus on the short term fluctuations, but include the trends to capture the above-mentioned effects.

Now we turn to the verification of our hypotheses. First, as for (H1) the most interesting results concern *Total Disbursements* as % of GDP.[7] The coefficient on this variable ranges from .01 to .02 in all specifications, generally significant at 5% level. In order to interpret

[7] In the specifications presented in Appendix B *Total Disbursements* is the primary variable of choice as it is of interest for testing hypotheses (H1) and (H2). We have explored other specifications and the results for variables of interest are robust. For example, inclusion of one of the following: *TotalReceipts*, *PrimaryBalance*, *NetLending* does not significantly affect the conclusions on testing the hypothesis, though admittedly including larger subsets of fiscal (as well as cyclical) variables, causes some problems with multicolinearity.

this, we must bear in mind the fact that the share of government investment in total disbursements ranges from .02 to .24 with a mean of .07. This means that if public spending cuts were to proportionally affect all types of spending, including public investment spending, we would expect that a 1 percentage point (of GDP) change in total disbursements, would be accompanied, on average, by a .07 percentage point of GDP change in disbursements on investment. As our estimate is between .01 and .02, significantly lower than .07 this means that a 1% of GDP cut in total disbursements was accompanied with only a .01 to .02 % of GDP cut in public investment. Thus public investment was significantly less affected than other types of spending. In fact, this is an argument against the hypothesis that government investment is the least rigid component of public spending - as it is on average significantly less affected than other categories, in other words, it is rigid according to our definition. But this is not the whole story - the coefficient on the *fiscal consolidation dummy* is significantly negative in most specifications (see Tables 3 and 6) and ranges from -.03 to -.27 (though a majority of values fall between -.09 and -.16). As the average cut in total spending during fiscal consolidation ranged from -1.2 to -2.3 % of GDP (definitions 1-4) and -.16 to -1.1 % of GDP (definitions 5-10), when it comes to fiscal consolidations public investment becomes much more affected than usual. Column (5) of the table presents the estimated average cuts of public investment during fiscal consolidations, which are usually similar or higher, in some cases even significantly higher, than what we would expect if it was to be cut in proportion with other types of spending - the latter values are given in column (4). Thus overall, the behavior of investment spending during fiscal consolidations does not allow us to further claim that public investment is a highly rigid component.[8]

[8]Notably the effect calculated in Table 3 still disregards the effect of the interaction between the election frequency and fiscal consolidation. Adding up the effects, we would obtain an even higher public investment effect of fiscal consolidation, which can partly be seen by comparing the coefficients on fiscal consolidation dummies from Table 6: specifications (1)-(8), where the interaction variable

Hence one of the most important results of our analysis: though public investment spending fluctuates less than other types of government spending, a *severe* fiscal cut falls strongly on this type of spending.[9]

Def. No. (1)	Total Disbursements Change* (2)	Fiscal Dummy coefficient (3)	Expected Proportional Change* (4)	Actual estimated Change* (5)	Std. Err. of col. (5) (6)
1	-2.28	-.14	-.159	-.169	(.04)
2	-1.84	-.16	-.128	-.183	(.10)
3	-1.19	-.094	-.083	-.109	(.04)
4	-1.26	-.124	-.088	-.140	(.03)
5	-1.03	-.134	-.072	-.147	(.03)
6	-1.06	-.106	-.073	-.119	(.05)
7	-0.46	-.023	-.031	-.028	(.03)
8	-0.69	-.105	-.048	-.113	(.03)
9	-0.58	-.107	-.040	-.114	(.05)
10	-0.16	-.088	-.011	-.090	(.03)

*Changes in % of GDP

Table 3: Public investment cuts during fiscal consolidations - column (1) - definition of consolidation (see Appendix A); column (2) - average cut in total disbursements during consolidation; column (3) - estimated coefficient on the fiscal consolidation dummy; column (4)- expected cut if public investment moved proportionately with total disbursements; column (5) - the actual estimated effect through the total disbursements coefficient and fiscal dummy; column (6) - standard error of (5). Based on specification (1) from Table 6.

As for hypothesis (H2), we do not find the election year nor the election year interacted with the fiscal consolidation dummy to have any influence on our dependent variable. This seems to confirm the previous findings of De Haan et al. (1996) - that political variables do not have influence on government investment. However by including an election frequency variable we find that governments with a shorter horizon were tending to expand public investment spending (the positive coefficient on *Election Frequency* is equivalent with the

is included in the specification with (10)-(11) where it is excluded.

[9]The exclusion of lagged dependent variables, as for example in specifications (9) and (12) in Table 6 causes some rise in the coefficient on total disbursements, but it is still significantly lower than the share of public investment in total government spending thus public investment is still found a rigid component of the latter. It also causes an increase in the coefficient on the fiscal consolidation dummy - thus is in line with the finding that public investment gets cut significantly more in large fiscal adjustments. In the end, we decide to include the lagged dependant variables in the equation, but point out that the broad conclusion on the rigidity of public investment spending and its reaction to fiscal consolidation is sustained regardless of this.

fact that myopic governments raise investment). Yet as the coefficient on this variable interacted with the fiscal cut dummy *ElFreq_x_ConsD* is of a larger magnitude and negative, we find that in fact governments with a shorter policy horizon (higher per annum number of elections) did not hesitate to slash investment more than governments with a longer horizon, when pursuing a strong fiscal adjustment. We interpret this as confirmation of the hypothesis that a myopic government would be less reluctant to cut government investment, what can be seen as an attempt to win votes in a politically less stable situation by laying the weight of cuts on less visible spending - i.e. investment spending.

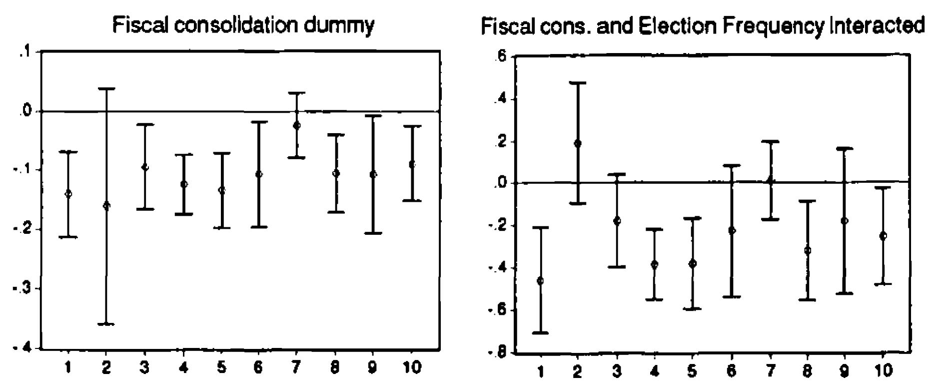

Figure 3: LSDV estimates of the coefficients on the *fiscal consolidation dummy* and *election frequency interacted with fiscal consolidation* with 95% confidence intervals. Specification (1) from Table 6.

As mentioned, in order to check the robustness of our results for (H1) and (H2) we use different definitions of a fiscal consolidation, both per se and in the interaction with the election frequency. The results are quite consistent and in Figure 3 we show the 95% confidence intervals on the coefficients estimated with the specification as in Table 2 but with all 10 definitions of a fiscal cut using LSDV.

In case of the third hypothesis (H3), we find strong support for the pro-cyclical stance of public investment. Out of three proxies used for capturing the influence of the business cycle, the positive coefficient on the output gap and negative on the unemployment rate

- in both cases highly significant. As for the growth rate of GDP, the coefficient is found either insignificant or mildly positive i.e. consistent with our claim of pro-cyclically. Thus overall, we can say that public investment spending exhibited a pro-cyclical pattern.

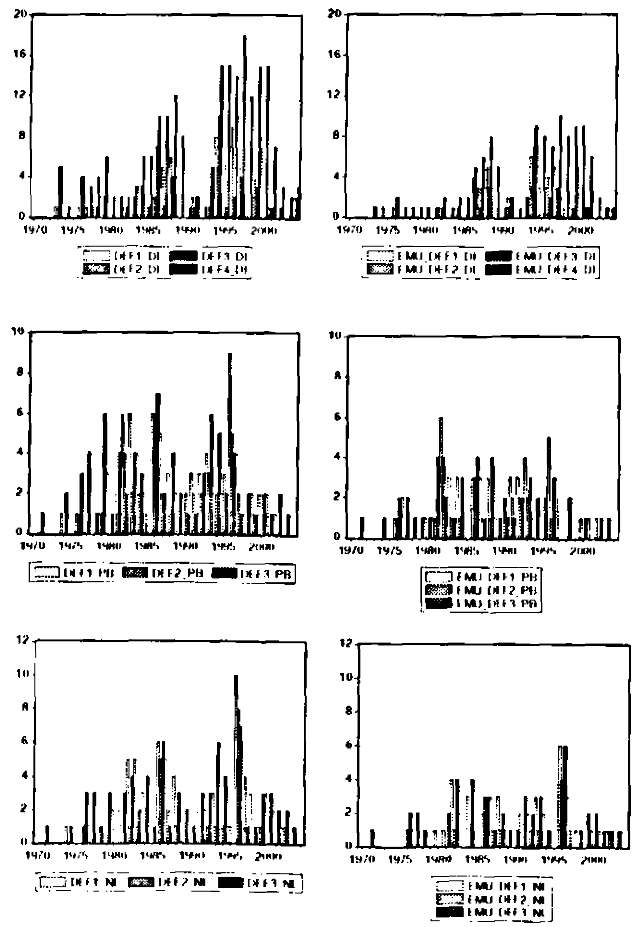

Figure 4: Distribution of fiscal consolidations according to different definitions (top to bottom: based on total disbursements, primary balance, net lending). Right panel: EMU member countries, left panel - entire sample.

Finally, the EMU dummies are not significant in most specifications at the 10% level, and insignificant at 5% for all of the specifications. The slight domination of the negative values in most specifications may be a sign of some negative effect for the EMU members, but the time span of the EMU may be too short to draw any conclusions. Therefore, in the case of the last hypothesis (H4) we do not find any particular change of the behavior of public investment spending due to the EMU. We find no evidence of the effect of public spending cuts on public investment, due to monetary integration being any different from other fiscal consolidations. It is important to note however, that the period of a run up to the euro, that is the second half of the 1990s was flagged by a large amount of fiscal consolidations (see Figure 4) in the countries aiming to join the common currency, in the EU members deciding to opt-out but also, albeit to a smaller extent, in the countries not in the EU. Consolidations in countries aiming to fulfil the Maastricht treaty entry criterion on fiscal deficits, as well as in the other EU members, have caused strong cuts in public investment, though admittedly the effect on public investment spending was similar (higher-than-proportional) as during fiscal consolidations in other periods, or fiscal consolidations in other countries in the 1990s. Notably, a larger-than-average number of fiscal consolidations in 1997-1998 (averaging across our definitions) can be observed in Italy, Finland and Austria, but also in Sweden and the UK.

II.5 Conclusions

In this paper we looked at the general decrease in public investment levels in the OECD countries in the last 30 years. We found that public investment as a share of GDP halved in most countries, following a linear downtrend. The main questions of our interest focused on movements around this trend. According to our results, despite being on average a fairly rigid component of total government spending, public investment expenditures are strongly affected during *large* fiscal cuts, thus confirming the claims of Oxley and Martin (1991). As for the policy horizon and its influence on investment spending, we do not find support for the claim that governments with a higher election turnover invest less, but we do find that such myopic governments lay more of the weight of a large fiscal cut onto investment. We do not find, however, the proximity of an upcoming election having a significant influence on this. Next, we found, similarly to Lane (2003), Gali and Perotti (2003) and Turrini (2004) that the fluctuations in public investment are quite pro-cyclical, in the short-term. This cannot be treated as evidence against the Keynesian cycle-smoothing role of investment, as because of the delayed outcome of public investment, we cannot simply expect that the effect is also pro-cyclical. This means precisely that governments tend to increase investment during booms and reduce it in downturns. As for the influence of the Maastricht Treaty constraints or the SGP we did not find them directly affecting short-term fluctuations in public investment.

As we control for the general downtrend in investment in most industrialized countries, we can say that there may be some danger of political volatility and incidental fiscal consolidation attempts inducing serious one-off cuts in public investment spending. This effect may be important as the potential benefits of public investment are delayed in time, thus weaker governments, or large fiscal cuts may result in the under-provision of some infrastructure in the future. Moreover, as Alesina and Perotti (1996) claim that fiscal consolidation done through public investment is not persistent and is easily reversed, this generally casts some doubt on the purposefulness of such adjustments.

Finally, though there is no evidence of the SGP itself having a contribution to this effect, there are at least two issues worth noting. Firstly, as the time period after its introduction is fairly short, perhaps it is too early to assess the effect. Secondly, we did not find any distinct effect of the SGP or the Maastricht Treaty, but we did find a strong effect of fiscal consolidations as such - thus fulfilling the EMU-related deficit requirements was quite probably no different than other budgetary consolidations.

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Appendix A: Data Description and Definitions

Gross Fixed Capital Formation of General Government (GG GFCF) the figures are cyclically unadjusted, taken as % of actual GDP. *(Source: OECD Economic Outlook)*[10]

Fiscal Variables *(Source: OECD Economic Outlook)*

Current Disbursements (CurDis), Total Receipts (TotRec), Total Disbursements (TotDis), Net Lending (NetLen), Primary Balance (PriBal) all of the General Government. The figures are cyclically unadjusted, as % of GDP.

Other Macroeconomic Variables *(Source: OECD Economic Outlook)*[11]:

Private Investment figures are cyclically unadjusted, as % of GDP.

Output Gap as % of Potential GDP, Inflation, Unemployment Rate

Real Interest Rate - the nominal interest rate on long-term government bonds (10-years*) less the yoy cpi inflation rate, according to the formula: $realrate = (\frac{100+nominalrate}{100+inflation}-1)*100$

Dummy variables:

Fiscal stringency dummies *(Source: OECD Economic Outlook)* idea oringinating from Alesina and Perotti (1996) and De Haan et al. (1996):

Various change indicators used (reduction in GG Total Disbursements, increase in GG Net Lending, increase in GG Primary Balance), all as % of GDP, cyclically adjusted, according to each of the following definitions:

Definition 1: $dumm_{it} = 1$ if fiscal change indicator in country i in year t falls more than 1.5% of GDP

Definition 2: $dumm_{it} = 1$ if fiscal change indicator in country i in year t and $t+1$ falls more than 1% of GDP

Definition 3: $dumm_{it} = 1$ in country i in year beginning a series of years t to $t+k$ that the fiscal change indicator falls more than 0% of GDP, and at least in one year in these series it falls more than 1% of GDP

Definition 4: $dumm_{it} = 1$ in country i and any year t to $t+k$ in the above mentioned series.

Election Frequency: *(Source: www.electionworld.org, http://psephos.adam-carr.net/)*

$ELEC\_D = 1$ if year t was a parliamentary(presidential in case of US) election in country i, 0 otherwise.

EL_FREQ = per annum number of parliamentary elections country i in 5 year window from $t-2$ to $t+2$, when closing in on year 2004 this window shifts to $< t-4, t >$

$EL\_FREQ\_FIS\_CON$ = the interaction of EL_FREQ with the fiscal cut dummy.

EMU dummies

Definition 1: EMU third phase dummy: 1 since 1999 for EMU members (2001 Greece), 0 otherwise;

Definition 2: Maastricht Dummy: 1 in 1998 for EMU members (Greece 2000), 0 otherwise;

Definition 3: Maastricht+SGP dummy: 1 since 1998 for EMU members;

Definition 4: EMU second phase dummy: 1 since 1993 for EMU members;

For Figure 3 the following definitions are used: 1-4 as in **Dummmy Variables** using total disbursements; 5-7 as 1-3 in **DV** using net lending, 8-10 as 1-3 in **DV** using primary balance. Identically for the interaction variable.

Countries used: AU - Australia, BD - Germany, BG - Belgium, CN - Canada, DK - Denmark, ES - Spain, FN - Finland, FR - France, GR - Greece, IR - Ireland, IT - Italy, JP - Japan, NE - Netherlands, NW - Norway, OE - Austria, PT - Portugal, SW - Sweden, UK - United Kingdom, US - United States.

[10] Portugal - Eurostat

[11] Greece - Eurostat

Variable	Test			
	LLC		IPS	
	intercept	trend	intercept	trend
General Government, as % of GDP				
G.F.C.F.	-1.67**	-2.09**	-1.51*	-1.94**
Total Disb.	-4.44***	0.50	-1.59*	1.69
Net Lending	-2.24**	-0.95	-3.32***	-1.70*
Prim. Balance	-2.10**	-1.37*	-4.47***	-2.16**
Total Receipts	-4.60***	0.43	-1.13	2.83
Current Disb.	-4.47***	-0.89	-1.48*	0.98
as % of GDP				
Private G.F.C.F.	-2.42***	-2.44***	-1.56*	-2.56***
% points				
Unemployment	-2.00**	-0.46	-0.78	-1.11
Output Gap[a]	-5.12***	-2.97***	-7.68***	-5.11***
Δ GDP	-16.31***	-17.01***	-16.51***	-16.81***
Interest Rate	-5.25***	-4.74***	-6.47***	-5.24***

[a] - % of potential GDP,

*,**,*** - rejects null at 10%, 5% and 1%

Table 4: Panel unit root tests. Levin, Lin, Chu : H_o: common unit root process, and Im, Pesaran, Shin H_o: individual unit root process. Lag selection - Schwarz BIC. Maximum lags 3.

Definitions 1-10		
Obs	Mean	Std. Dev.
72	-2.2825	1.087359
48	-1.839583	0.923983
62	-1.192258	0.9689058
212	-1.259623	1.067637
72	-1.033056	1.857593
29	-1.055172	1.01282
75	-0.4561333	1.466876
79	-0.6875949	1.791733
43	-0.5758139	1.862389
78	-0.1574359	1.55683

Table 5: Changes in total disbursements indicated by fiscal cut dummies.

Appendix B: Specifications

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
GGGFCF(-1)	.865***	.898***	.877***	.888***	.889***	.892***	.892***	.739***	-	.888***	.735***	-
	(.03)	(.03)	(.03)	(.03)	(.03)	(.03)	(.03)	(.02)		(.03)	(.02)	
GGGFCF(-2)	-.156***	-0.16***	-.161***	-.176***	-.177***	-.176***	-.175***	-	-	-.175***	-	-
	(.03)	(.03)	(.03)	(.03)	(.03)	(.03)	(.03)			(.03)		
OUTGAP	.032***	-	-	0.03***	.029***	.028***	.027***	0.03***	.058***	.027***	.031***	.058***
	(.00)			(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)
ΔGDP	-	.009	-	-	-	-	-	-	-	-	-	-
		(.00)										
UNEMP	-	-	-.037***	-	-	-	-	-	-	-	-	-
			(.00)									
TOT_DIS	.013***	.006	.013***	.014***	.014***	.011***	.012***	.013***	.023***	.012***	.012***	.023***
	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)
CONS_D	-.124***	-.118***	-.093***	-.122***	-.121***	-.122***	-0.12***	-.139***	-.232***	-0.16***	-.179***	-.256***
	(.02)	(.02)	(.02)	(.02)	(.02)	(.02)	(.02)	(.02)	(.04)	(.02)	(.02)	(.03)
PRIV_INV	-.007	.016*	.001	-.006	-.006	-.005	-	-.001	-.023	-	-.002	-.023
	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)		(.00)	(.01)		(.00)	(.01)
R_I	-.001	-.003	.002	-	-	-	-	-	-	-	-	-
	(.00)	(.00)	(.00)									
EMU_D	.033	-.003	-.017	.046	.045	-	-	-	-	-	-	-
	(.04)	(.04)	(.04)	(.04)	(.04)							
ELEC_D	-0.02	-.018	-.016	-.027	-	-	-	-	-	-	-	-
	(.02)	(.02)	(.02)	(.02)								
EL_FREQ xCONS_D	-.459***	-.508***	-.428***	-.428***	-.424***	-.435***	-.437***	-.431***	-.268	-	-	-
	(.12)	(.13)	(.12)	(.12)	(.12)	(.12)	(.12)	(.12)	(.20)			
Trends	17(-): AU***, DE***, BG***, CN,DK***, ES,FN**, FR,IT***, JP***,NE, NW,OE***, PT**, SW***, UK***,US, 2(+):GR, IR***,	15(-): AU***, DE***, BG***, CN,DK***, ES,FN, FR,IT, JP**, OE***, PT*, SW**, UK***,US, 4(+):GR, IR**,NE, NW,	14(-): AU***, DE**, BG***, CN,DK***, ES,FN, IT, JP***, OE***, PT**, SW**, UK***,US, 5(+):FR, GR**,IR, NW,	17(-): AU***, DE***, BG***, CN,DK***, ES,FN**, FR,IT***, JP***,NE, NW,OE***, PT***, SW***, UK***,US, 2(+):GR**, IR*,	17(-): AU***, DE***, BG***, CN,DK***, ES,FN**, FR,IT***, JP***,NE, NW,OE***, PT**, SW***, UK***,US, 2(+):GR**, IR*,	17(-): AU***, DE***, BG***, CN,DK***, ES,FN**, FR,IT**, JP***,NE, NW,OE***, PT**, SW***, UK***,US, 2(+):GR***, IR*,	17(-): AU***, DE***, BG***, CN,DK***, ES,FN*, FR,IT**, JP***,NE, NW,OE***, PT***, SW***, UK***,US, 2(+):GR***, IR*,	16(-): AU***, DE***, BG***, CN,DK***, ES,FN, FR,IT**, JP***, NW,OE***, PT***, SW**, UK***,US, 3(+):GR***, IR*,NE,	16(-): AU***, DE***, BG***, CN***,DK***, FN***, FR*,IT***, JP***,NE***, NW***,OE***, PT***, SW***, UK***,US*, 3(+):ES**, GR***,IR**,	17(-): AU***, DE***, BG***, CN,DK***, ES,FN**, FR,IT***, JP***,NE, NW,OE***, PT***, SW***, UK***,US, 2(+):GR***, IR*,	16(-): AU***, DE***, BG***, CN,DK***, ES,FN*, FR,IT**, JP***, NW,OE***, PT***, SW**, UK***,US, 3(+):GR***, IR*,NE,	16(-): AU***, DE***, BG***, CN***,DK***, FN***, FR*,IT***, JP***,NE***, NW***,OE***, PT***, SW***, UK***,US*, 3(+):ES**, GR***,IR**,

Table 6: Estimation results for different sample specifications.

Appendix C: Simulations

We present a very simple Monte Carlo simulation exercise in order to yield some light on the performance of the different estimation techniques in a macro-panel of similar size. For this purpose, let us use a Data Generating Process (DGP) as follows:

$$y_{it} = \rho_1 * y_{it-1} + \rho_2 * y_{it-2} + \beta * x_{it} + \phi_i * t + \mu_i + \varepsilon_{it} \qquad (2)$$

where ϕ_i is the country specific time trend coefficient, independent, identically distributed (i.i.d.) $\sim$ *negative uniform*$(0,1)$; μ_i is the country specific fixed effect, i.i.d. $\sim N(0,\sigma_\mu^2)$; ε_{it} is the random error term i.i.d. $\sim N(0,\sigma_\epsilon^2)$; $cor(x_{it},\mu_i) = \sigma_{x\mu}$ is the correlation between the country specific fixed effect and the x variable.

In the simulation exercise we experimented with various levels of σ_μ^2, σ_ϵ^2, $\sigma_{x\mu}$, N and T in order to look at the behavior of the finite sample bias and the precision of the four estimation methods used in the applied part of the paper. Each estimation method was repeated 1000 times.

Overall, for all the specifications with ρ_1 in $[0.6, 0.9]$; ρ_2 in $[-0.3, -0.1]$; $\beta = 1$ and N=20, N=35, which as we will see in the following section, resembles our estimates, the following patterns can be observed: on average it is the AH-IV estimator that seems to be close to the DGP, but its standard deviation is of an order of a magnitude higher (especially on the lagged dependent variable) making it inefficient in a single estimation. Both the LSDV and the AB-GMM estimators are very similar, exhibiting on average a low negative finite sample bias for ρ_1, a negative (over estimating) finite sample bias on ρ_2 and estimating the coefficients on β very precisely. They also exhibit relatively low variance. In reasonable specifications the cross-sectional dimension of N=20 proves too low to give an edge to the AB-GMM estimator, which is in theory more precise than the LSDV for a fixed T as N goes to infinity - due to the elimination of the 'Nickell bias'. As for the OLS estimator, as we expected it has a bias due to the correlation of the fixed effect with the explanatory variable, which it does not account for. Obviously if the fixed effects were not correlated with the explanatory variable[12] the LSDV estimator would be still consistent but inefficient but the OLS would be consistent. Generally OLS tends to over-estimate the values of the coefficient on the LDV, as can be expected due to a positive correlation between the fixed effect and the LDV, it exhibits the strongest biases on the explanatory variable X and as could be expected the direction and magnitude of this bias depends positively on the direction of the correlation between the fixed effect and the explanatory variable.

Supposing that our econometric model specification is not degenerate from the assumptions imposed in the MC simulations we proceed to using the LSDV, and the very similar AB-GMM estimators for the purpose of interpretation. Our simulations confirm that the AB-GMM estimators' advantage in micro-panels of long cross sections but short time series does not apply to typical macro-panels.

[12]In which case they would not be fixed effects but random effects.

Calibration	$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx 0.3$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx 0.6$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx 0.9$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx 0.3$ $N=10$, $T=35$			
Statistic	OLS	LSDV	AHIV	ABGMM	OLS	LSDV	AHIV	ABGMM	OLS	LSDV	AHIV	ABGMM	OLS	LSDV	AHIV	ABGMM
$\widehat{\rho_1}$																
mean	0.896	0.743	0.85	0.732	0.892	0.742	0.834	0.732	0.867	0.741	0.834	0.731	0.882	0.738	0.884	0.736
s.d.	0.043	0.039	0.276	0.043	0.041	0.038	0.363	0.041	0.036	0.041	0.312	0.043	0.061	0.057	0.581	0.058
afs bias	11.961	-7.186	6.198	-8.484	11.543	-7.216	4.268	-8.482	8.434	-7.357	4.227	-8.583	10.277	-7.705	10.483	-8.004
rmse	0.105	0.07	0.28	0.08	0.101	0.069	0.365	0.08	0.077	0.071	0.313	0.081	0.102	0.084	0.587	0.087
$\widehat{\rho_2}$																
mean	-0.179	-0.239	-0.2	-0.238	-0.18	-0.237	-0.199	-0.236	-0.188	-0.235	-0.201	-0.234	-0.179	-0.239	-0.199	-0.238
s.d.	0.039	0.036	0.051	0.036	0.038	0.036	0.055	0.036	0.035	0.036	0.051	0.037	0.056	0.051	0.084	0.052
afs bias	-10.49	19.315	0.037	18.84	-9.878	18.694	-0.367	18.146	-6.194	17.643	0.721	16.853	-10.258	19.275	-0.304	19.211
rmse	0.044	0.053	0.051	0.052	0.043	0.052	0.055	0.051	0.037	0.05	0.051	0.05	0.059	0.064	0.084	0.064
$\widehat{\beta}$																
mean	1.163	0.985	1.026	0.981	1.332	0.981	1.014	0.977	1.522	0.994	1.024	0.991	1.159	0.992	1.055	0.99
s.d.	0.147	0.135	0.221	0.14	0.149	0.141	0.242	0.146	0.124	0.135	0.23	0.141	0.206	0.193	0.371	0.196
afs bias	16.254	-1.517	2.594	-1.88	33.232	-1.867	1.449	-2.258	52.166	-0.642	2.431	-0.941	15.875	-0.816	5.527	-1.018
rmse	0.219	0.136	0.222	0.141	0.364	0.142	0.242	0.147	0.536	0.136	0.231	0.141	0.26	0.193	0.375	0.197

Calibration	$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx -0.3$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx -0.6$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx -0.9$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx 0.6$ $N=10$, $T=35$			
$\widehat{\rho_1}$																
mean	0.885	0.742	0.85	0.733	0.864	0.743	0.83	0.734	0.799	0.742	0.823	0.736	0.883	0.74	0.868	0.737
s.d.	0.043	0.039	0.272	0.043	0.041	0.039	0.322	0.041	0.039	0.041	0.268	0.043	0.059	0.056	0.661	0.057
afs bias	10.589	-7.19	6.286	-8.432	8.017	-7.107	3.792	-8.214	-0.111	-7.204	2.932	-7.945	10.384	-7.531	8.52	-7.899
rmse	0.095	0.07	0.276	0.08	0.076	0.069	0.323	0.078	0.039	0.071	0.269	0.077	0.102	0.082	0.664	0.085
$\widehat{\rho_2}$																
mean	-0.181	-0.24	-0.199	-0.239	-0.186	-0.24	-0.2	-0.239	-0.203	-0.242	-0.202	-0.24	-0.18	-0.237	-0.201	-0.237
s.d.	0.041	0.036	0.051	0.036	0.042	0.036	0.054	0.036	0.039	0.036	0.051	0.037	0.052	0.05	0.077	0.05
afs bias	-9.5	19.764	-0.313	19.27	-6.892	20.102	-0.245	19.427	1.413	21.077	0.818	20.053	-10.026	18.523	0.623	18.409
rmse	0.045	0.053	0.051	0.052	0.044	0.054	0.054	0.053	0.04	0.056	0.051	0.055	0.056	0.062	0.077	0.062
$\widehat{\beta}$																
mean	0.834	0.986	1.027	0.983	0.573	0.982	1.012	0.977	0.253	0.991	1.019	0.991	1.309	0.979	1.024	0.978
s.d.	0.147	0.136	0.221	0.14	0.169	0.141	0.23	0.146	0.151	0.135	0.212	0.141	0.214	0.196	0.382	0.198
afs bias	-16.566	-1.453	2.749	-1.733	-42.708	-1.84	1.154	-2.257	-74.709	-0.884	1.887	-0.901	30.904	-2.123	2.367	-2.218
rmse	0.222	0.137	0.223	0.141	0.459	0.142	0.23	0.147	0.762	0.135	0.212	0.141	0.376	0.197	0.383	0.2

Calibration	$\sigma_\mu^2=1, \sigma_\epsilon^2=0.01, \sigma_{x\mu} \approx 0.3$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=0.01, \sigma_{x\mu} \approx 0.6$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=0.01, \sigma_{x\mu} \approx 0.9$ $N=20$, $T=35$				$\sigma_\mu^2=1, \sigma_\epsilon^2=1, \sigma_{x\mu} \approx 0.9$ $N=10$, $T=35$			
$\widehat{\rho_1}$																
mean	1.268	0.793	0.827	0.793	1.225	0.792	0.807	0.791	1.124	0.793	0.863	0.792	0.861	0.74	0.769	0.738
s.d.	0.068	0.013	0.503	0.014	0.053	0.013	0.416	0.015	0.031	0.013	2.095	0.015	0.052	0.056	2.947	0.058
afs bias	58.488	-0.842	3.323	-0.924	53.134	-0.965	0.87	-1.108	40.521	-0.913	7.892	-0.999	7.635	-7.516	-3.851	-7.784
rmse	0.473	0.015	0.503	0.016	0.428	0.015	0.416	0.017	0.326	0.015	2.095	0.017	0.08	0.082	2.946	0.085
$\widehat{\rho_2}$																
mean	-0.403	-0.202	-0.201	-0.201	-0.383	-0.2	-0.2	-0.2	-0.341	-0.2	-0.205	-0.199	-0.188	-0.235	-0.204	-0.234
s.d.	0.053	0.01	0.031	0.011	0.044	0.01	0.027	0.01	0.027	0.009	0.184	0.009	0.05	0.049	0.082	0.049
afs bias	101.521	0.761	0.491	0.601	91.396	0.188	0.208	0.081	70.395	-0.128	2.601	-0.265	-6.182	17.277	2.088	17.186
rmse	0.21	0.01	0.031	0.011	0.188	0.01	0.027	0.01	0.143	0.009	0.184	0.009	0.051	0.06	0.082	0.06
$\widehat{\beta}$																
mean	1.076	0.999	1.013	0.999	1.131	1	1.004	0.999	1.179	0.998	1.03	0.998	1.483	0.977	0.986	0.978
s.d.	0.031	0.014	0.253	0.014	0.03	0.014	0.208	0.014	0.025	0.014	1.047	0.014	0.185	0.197	1.189	0.198
afs bias	7.638	-0.097	1.334	-0.121	13.056	-0.036	0.406	-0.092	17.861	-0.159	3.023	-0.192	48.276	-2.321	-1.388	-2.247
rmse	0.083	0.014	0.253	0.015	0.134	0.014	0.208	0.014	0.18	0.014	1.046	0.015	0.517	0.199	1.188	0.199

Table 7. Monte Carlo simulations: no. of repetitions = 1000 , afsb - average finite sample bias, rmse - root mean squared error

Chapter III
Measuring Long Run Exchange Rate Pass-Through

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Abstract

We discuss the issue of estimating short- and long-run exchange rate pass-through and review some problems with the measures recently proposed in the literature. Theoretical considerations suggest a long-run Engle and Granger cointegrating relationship (between import unit values, the exchange rate and foreign prices), which is typically ignored in existing empirical studies. We use time series and up-to-date panel data techniques to test for cointegration with the possibility of structural breaks and show how the long-run may be restored in the estimation. We also discuss what difference is made to the policy debate surrounding pass-through.[§]

JEL Classification Numbers:F14, F31, F36, F42, C23

Keywords: *exchange rates. pass-through. import prices. panel cointegration. structural breaks.*

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III.1 Introduction

A large number of recent papers (see for example Campa and Gonzalez-Minguez, 2006; Campa, Goldberg and Gonzalez-Minguez, 2005; Frankel, Parsley and Wei, 2005; Marazzi et al., 2005) have investigated the issue of exchange rate pass-through (ERPT) of foreign to domestic prices. Studies of ERPT have been conducted both for the United States and for countries of the euro area, and a focus of some importance has been on its evolution over the past two decades, in response to changes in institutional arrangements (such as the inauguration of the euro area) and to shocks to the monetary system (such as Black Wednesday and the ERM crisis in 1992).

Several economic policy issues hang upon the determination of the rate of pass-through from exchange rates to prices, and its evolution, both in various time horizons as well as in different sectors. These include issues relating to pricing strategies of foreign exporting firms, the persistence of inflation, successfulness of inflation forecasting and the impact of entering into a monetary union. For the European Union countries, more particularly those belonging to the euro area, the issues listed above are of considerable importance. A positive or negative answer to the question of whether pass-through has fallen has an important bearing on inflation persistence and the success of protocols such as the Lisbon Strategy which calls for structural reforms across the European Union.

A notable lacuna in the literature, we argue, is a clear disjunction between the well-worked-out theoretical arguments surrounding the key determinants of pass-through, and the inappropriate techniques used to estimate import or export exchange rate pass-through equations. Thus, while almost all the theories contain a long-run or steady-state relationship in the levels of a measure of import unit values (in domestic currency), the exchange rate (relating the domestic to the numeraire currency) and a measure of foreign prices (unit values in the numeraire currency, typically US dollars), this long run is routinely disregarded in most of the empirical implementations. This may seem surprising for at least two reasons. First, proper determination of the short run ERPT relies on appropriate assumptions about the long run. Second, as monetary policy tends to be medium-term oriented, issues like inflation forecasting and policy actions should in principle look beyond short term developments for a better understanding of the underlying forces.

Since it is commonly agreed that the time series considered are integrated, one way of defining the long run is in the sense of Engle and Granger (1987), henceforth EG, where the long run is given by the so-called cointegrating relationship. The reason for ignoring this long run, and substituting it by an ad hoc measure, is the failure to find evidence in the data for cointegration. The difficulty inherent in such a re-definition of the long run is two-fold, first the contradiction between a theoretical prediction of a steady state that cannot be found in the data, and, second, the ad hoc measure proposed being no more than an extended version of the estimate of the short-run (and, as we shall see below, strongly dominated by the estimated short-run). It is possible that the source of the difficulty is the estimation method used - typically single-equation autoregressive distributed lag (ARDL) models - which may not be powerful enough to verify the theory for the span of data available. Therefore, instead of looking for a new definition of the long-run, a more satisfactory approach is to look for the long run relationship using more appropriate and powerful methods, such as those which allow for changes in the long run or use more powerful panel data methods. This is the route we follow in this paper.

Focusing on a specification of ERPT into import prices from Campa, Goldberg and Gonzalez-Minguez (2005), we argue in particular that: (a) the long run, in the sense of Engle and Granger (1987), is restorable once appropriate testing strategies (including lag length selection) are adopted and proper account is taken of the possibility of breaks in the long-run relationship; (b) the estimate of the 'long run' used in the empirical literature is entirely arbitrary and sensitive to the results of a number of misspecification issues; (c) once the distinction is established between the long run (with a break) in the sense of Engle and Granger (1987) and the definition used in the ERPT literature, it becomes important to investigate the relative magnitudes of these alternative measures and to interpret each differently; and (d) it is important to allow for breaks in the long-run theoretical relationship to take due account of pass-through rates in response to changes in financial regime (such as those following Black Wednesday in 1992 or the ERM arrangements which came into force post 1996.) Not to take explicit account of such changes, which are easily evident in the data, is to make serious mistakes in estimation and inference.

We begin in the next section with a very brief overview of the theoretical framework. We next move to the key empirical issues, since these are the main areas of our concern, and in Section III.3 establish the key ERPT equation in levels and differences. We present the definition of short- and long-run ERPT assumed by the empirical literature and assess its adequacy. Section III.4 presents the data.

Section III.5 proceeds by first looking in more detail at some results reported by Campa and Gonzalez-Minguez (2006), CM hereafter. Here we look more closely at their estimates of pass-through and show that the distinction between the short- and long-run is somewhat confused, once the statistical significance of the coefficients is taken into account. This adds to the uncertainty surrounding the use of the standard measures for short and long run pass-through in the literature. We compare the CM measures with our estimates of the Engle-Granger long run wherever these exist. We continue our analysis of the Engle-Granger long run by allowing for structural breaks in the cointegrating vector using methods developed by Gregory and Hansen (1996), and show that there is strong evidence of cointegration once account is taken of breaks in the deterministic components of the cointegrating regressions (such as the constant) and in the cointegrating vector. Interesting contrasts are drawn between the long-run coefficient under the CM definition and those obtained under the specification of a broken long run. Graphs present the estimates constructed under different assumptions and make the comparisons.

In Section III.6 the analysis of the long run is conducted using panel methods developed by Banerjee and Carrion-i-Silvestre (2006), which are appropriate for looking at cointegration in panels. This is particularly useful in the short-sample analysis where the time series dimension T is small. The tests used allow not only for breaks in the individual units of the panel but also for cross-unit dependence. The results seem to confirm strongly the existence of cointegration, with easily interpretable break dates.

Concluding remarks are contained in Section III.7 where we discuss whether we should reconsider the traditional way of computing the long run pass-through.

III.2 Exchange Rate Pass-Through into Import Prices

By definition[1] import prices for any type of goods j, MP_t^j are a transformation of export prices of a country's trading partners XP_t^j using the bilateral exchange rate ER_t and dropping superscript j for clarity:

$$MP_t = ER_t \cdot XP_t \tag{1}$$

In logarithms (depicted in lower case):

$$mp_t = er_t + xp_t \tag{2}$$

[1]This section is based on Campa, Goldberg and Gonzalez-Minguez (2005), CGM hereafter.

Where the export price consists of the exporters marginal cost and a markup:

$$XP_t = FMC_t \cdot FMKUP_t \tag{3}$$

So that in logarithms we have:

$$xp_t = fmc_t + fmkup_t \tag{4}$$

Substituting for xp_t into equation (2) yields:

$$mp_t = er_t + fmkup_t + fmc_t \tag{5}$$

for each sector in each of the countries.

The literature on industrial organization yields insight into why the effect of the change in er_t on mp_t may differ from one, through mark-up determinants like competitive conditions that exporters have to face in the destination markets. Hence, the estimated pass-through elasticities are a sum of three effects:

- effects of the unity translation effects of the exchange rate movement,
- the response of the markup in order to offset this translation effect,
- the effect on the marginal cost that are attributable to the exchange rate movements, such as the sensitivity of input prices to exchange rates.

Markup responsiveness depends on the market share of domestic producers relative to foreign producers, the form of competition that takes place in the market for the industry, and the extent of price discrimination. Other factors affecting pass-through are the currency denomination of exports and structure and importance of intermediate goods markets.

The empirical setup of CGM is based on (5) which assumes unity translation of exchange rate movements. However, as mentioned above, exporters of a given product can decide to absorb some of the exchange rate variations instead of passing them through to the price in the importing country currency. If the pass-through is complete (producer currency pricing), their mark-ups will not respond to fluctuations of the exchange rates, thus leading to a pure currency translation. At the other extreme, they can decide not to vary the prices in the destination country currency (local currency pricing or pricing to market) and absorb the fluctuations within the mark-up. Thus, mark-ups in an industry are assumed to consist of a component specific to the type of good, independent of the exchange rate and a reaction to exchange rate movements:

$$fmkup_t = \alpha + \Phi er_t \tag{6}$$

Also important to consider are the effects working through the marginal cost. These are a function of demand conditions in the importing country; marginal costs of production (labor wages) in the exporting country and the commodity prices denominated in foreign currency:

$$fmc_t = \eta_0 \cdot y_t + \eta_1 \cdot fw_t + \eta_2 \cdot er_t + \eta_3 \cdot fcp_t \tag{7}$$

Substituting (7) and (6) into (5), we have:

$$mp_t = \alpha + \underbrace{(1 + \Phi + \eta_2)}_{\beta} er_t + \eta_0 \cdot y_t + \eta_1 \cdot fw_t + \eta_3 \cdot fcp_t + \varepsilon_t \tag{8}$$

where the coefficient β on the exchange rate er_t is the pass-through elasticity. Obviously, this is a simple approach, with a highly reduced form representation, where one can have no hope in identifying Φ from η_2. In the CGM 'integrated world market' specification, $\eta_0 \cdot y_t + \eta_1 \cdot fw_t + \eta_3 \cdot fcp_t$, independent of the exchange rate, is dubbed as the opportunity cost of allocating those same goods to other customers, is reflected in the world price of the product fp_t in the world currency (here taken to be the US dollar)[2]. Thus the final

[2]The integrated market hypothesis in CM is based on the assumption that there exists a single

equation can be re-written as follows:

$$mp_t = \alpha + \beta \cdot er_t + \gamma \cdot fp_t + \varepsilon_t \quad (9)$$

which gives the long run relation between the import price, exchange rate and a measure of foreign price.[3]

At this point it is perhaps important to stress two issues. First the exchange rate pass-through literature can be divided in two main streams - with papers which focus on 'first step' pass-through, i.e. ERPT into import prices and those which consider 'second step' pass-through, i.e. into consumer prices. As has been made clear above, for the purpose of this paper we will look at ERPT into import prices.

The second issue is the relevance of ERPT for the design of monetary policy. Since ERPT is a channel linking exchange rates with prices, it is often named as one of the key determinants of monetary policy design. There is a vast literature on optimal monetary policy, starting with models developed for a closed economy, and extended to the open economy (see for example Obstfeld, 2002).

The general consensus is that the optimal design of monetary policy should differ depending on the degree of ERPT. For example, Devereux and Engel (2002) derive a model in which local currency pricing causes relatively high volatility of the exchange rate and low

world market for each good. Therefore, regardless of the origin of the product, on the world market, it has one world price. This price constitutes the opportunity cost of selling to a local market. Thus, in the CM setup for the integrated market and, consequently, in ours, it proxies for the foreign price. The currency denomination does not in fact matter, as long as the exchange rate for the local currency is taken vis-a-vis this 'world' currency. In the CM case the extra-euro area imports into the euro area as a whole(denominated in US dollars) are taken as a proxy for the world price.

[3]It is not uncommon in the literature to insert additional control variables on the right hand side of this equation. For example, Marazzi et al. (2005) use commodity prices, in order to control for changes in marginal costs that producers may face. This seems undesirable in our specifications for at least two reasons. First, we are concerned with ERPT in individual sectors, and thus the appropriate equation for commodity sectors will already contain the commodity price - thus the control variable would be redundant. Second, and more generally any marginal cost effect is assumed to work through the 'world price'.

volatility of prices. In their model exchange rates can fluctuate to a large extent, since due to incomplete pass-through they have little effect on other macro variables.

Adolfson (2001) stresses the understanding of ERPT for monetary policy's response to shocks, and emphasizes that incomplete ERPT yields the exchange rate channel of shock transmission to have less impact, and thus require a lower interest rate response. Moreover, lower pass-through implies less conflict between inflation and output variability and leads to an increase in exchange rate volatility. The result that nominal exchange rate volatility increases as pass-through declines is generated by the fact that low pass-through is induced by large import price stickiness and it is because of price stickiness that the endogenously determined exchange rate must, through larger movements than if prices were less sticky, absorb a part of the relative price adjustment due to a country specific shock. However, this ignores the direct cost of exchange rate fluctuations which affect the variability in the import sector prices and thus offsets the incentive to use the exchange rate channel. On the other hand, Corsetti and Pesenti (2005) propose a model in which the degree of ERPT, that is the degree of exposure of firms mark-ups to exchange rate fluctuations is an important determinant of optimal monetary policy. In their paper, the focus of policymakers on stabilizing both domestic prices and the output gap, combined with the inability of firms to fully pass-through exchange rate fluctuations to prices, may cause an inefficiently high level of prices. As internal stabilization policies may raise the volatility of the exchange rate, this may lead exporters, whose revenues are exposed to such volatility to try to stabilize their profits by charging higher prices.

Consequently, Smets and Wouters (2002) find limited room for an exchange rate stabilization channel in the conduct of monetary policy in the presence of incomplete pass-through and therefore conclude that central banks should refrain from engineering large exchange rate volatility in order to stabilize prices. Thus, in principle, monetary policy aiming at internal stabilization may make better use of the exchange rate channel if ERPT is complete, while being unable to do so in case of (at least partial) local currency pricing.

Importantly, much of the focus of the literature concentrates on short run pass-through, and assumes that pass-through in the long-run is full (see, among others Smets and Wouters, 2002; Adolfson, 2001). This is usually the result of the imposition of staggered price setting, which allows the response to an exchange rate shock with imperfect adjustment in the short run, because of menu costs, and a gradual full incorporation of the change in the long run. On the other hand the literature focusing on price discrimination allows imperfect pass-through in the long run, as part of the adjustment is borne by firms mark-up (this issue is reviewed in more detail in Corsetti, Dedola and Leduc, 2006).

As we will show, there is some evidence that ERPT into import prices, is not always full even in the long run. These results points to the invalidity of the full-ERPT assumption and may have important implications for the proper estimation of the short-run pass-through and consequently the design of monetary policy. Importantly, this finding seems more in line with the price discrimination models as in Corsetti, Dedola and Leduc (2006).

Admittedly, there is a large degree of endogeneity in the observed ERPT and monetary policy. That is, pricing strategies of firms depend not solely on competition conditions on the market, but also on monetary policy, or rather the expected future monetary policy and the policy makers' credibility. The formation of a monetary union, as occurs in the middle of the samples used for the empirical exercise, is thus likely to have an important impact on ERPT (and vice versa) and any estimation method should take account of these changes. This is our guiding motivation for looking at long run relationships with structural change in our study of ERPT.

III.3 ERPT - estimation

Both economic theory and relevant tests lead us to think each of the series (import price, exchange rate and world price) as being characterized by a unit root. However, despite the underlying levels equation (1), CGM claim not to be able to reject the null hypothesis of the non-existence of a cointegrating relationship among the three series. Hence, they proceed by estimating equation (9) in first differences:

$$\Delta mp_t = a + \sum_{k=0}^{4} b_k \cdot \Delta er_{t-k} + \sum_{k=0}^{4} c_k \cdot \Delta fp_{t-k} + \varepsilon_t \tag{10}$$

for a certain type of good i in a certain country j. The superscripts have been omitted for clarity. Next, they define the coefficient b_0 and the sum of coefficients $\sum_{k=0}^{4} b_k$ as the short-run and long-run ERPT respectively.

At this point it is useful to focus on the CM definition of the long run pass-through. Since CM do not find evidence of the long-run in the Engle and Granger (1987) sense, they are forced to propose their own working definition of the long run. We claim that the CM definition of the long run pass-through, which is constructed by summing the estimated coefficients for the first five lags (i.e. lag 0 to lag 4), is somewhat arbitrary, and thus rather inadequate for the purpose of enquiring about the actual long run effect. It is not clear why the five lags are chosen and this rather ad hoc measure does not seem to take into account the significance of the coefficients on the individual lags. Taking for example the estimates for France (see Table 1) we can see that in the majority of cases only the coefficient on lag 0 is significant, while the following four lags are not significantly different from 0. As these coefficients are of relatively large magnitude, the number of lags is rather important - if one summed the first three, four, or six lags, the point estimate of the long run could differ vastly, though potentially would be as justified. The importance of our argument

France

	Lag 0	Lag 1	Lag2	Lag3	Lag4	CM LR
SITC0	0.96	-0.03	-0.01	-0.15	-0.04	0.74
	(0.09)	(0.11)	(0.11)	(0.11)	(0.08)	(0.10)
SITC1	0.01	0.59	-0.49	0.71	-0.41	0.40
	(0.2)	(0.25)	(0.26)	(0.25)	(0.21)	(0.25)
SITC2	0.77	0.16	0.05	0.06	-0.06	0.98
	(0.11)	(0.13)	(0.13)	(0.13)	(0.1)	(0.13)
SITC3	1.06	-0.02	0.1	-0.05	0.07	1.16
	(0.06)	(0.07)	(0.08)	(0.08)	(0.06)	(0.08)
SITC4	1.13	-0.14	-0.31	0.36	-0.08	0.97
	(0.25)	(0.3)	(0.31)	(0.31)	(0.24)	(0.33)
SITC5	0.87	0.08	-0.11	-0.12	0.1	0.81
	(0.17)	(0.19)	(0.19)	(0.19)	(0.15)	(0.26)
SITC6	1.11	-0.26	0.22	0.09	-0.17	1.00
	(0.09)	(0.11)	(0.11)	(0.11)	(0.08)	(0.09)
SITC7	1.12	-0.3	0.25	-0.02	-0.01	1.03
	(0.14)	(0.15)	(0.15)	(0.15)	(0.12)	(0.22)
SITC8	0.95	-0.17	0.11	-0.11	-0.01	0.76
	(0.08)	(0.1)	(0.11)	(0.1)	(0.07)	(0.12)

For each sector first line reports the estimated coefficient, and the second the standard error.

Table 1: Estimates of equation (10) - coefficients and standard errors on the lags of exchange rate - original CM sample 1989-2001. The last column reports the CM long run estimate.

for inference can be illustrated further by taking the coefficients for SITC0 from Table 1 - the CM long run is significantly different from 1, while if we redefine the "long run" as the sum of the first three lags, we could not be able to reject it being equal to 1. With SITC1 the example becomes even more visible - the five-lag CM long run is insignificantly different from 0, while significantly different from 1, whereas the four-lag "long run" would be significantly different from 0, while not differing significantly from 1.

The fact that CM are unable to find a cointegrated 'equilibrium' relationship between the variables in levels may seem surprising in light of the fact that the theoretical underpinning of the ERPT, is in fact a levels relationship, as in equation (1). We proceed by noting that if the cointegrated equilibrium relationship were to exist, the equation to be estimated should contain an error correction term (ECM), as in Engle and Granger (1987), and thus take the following form:

$$\Delta mp_t = a + \sum_{k=0}^{K_1} b_k \cdot \Delta er_{t-k} + \sum_{k=0}^{K_2} c_k \cdot \Delta fp_{t-k} + \lambda \underbrace{(mp_{t-1} - \widehat{\alpha} - \widehat{\beta} \cdot er_{t-1} - \widehat{\gamma} \cdot fp_{t-1})}_{ECM} + u_t \quad (11)$$

while equation (10) would be misspecified.

There are a number of reasons which could lead to a failure to find a cointegrating relationship in series which are suspected to be cointegrated. In particular, as we show below, appropriate lag length selection and proper accounting for a structural break, whether in single equations or more powerful panel methods, can change the inference on the existence of a 'long-run' relationship. This helps to provide a less arbitrary estimate of the long run ERPT and to assess changes to this elasticity following the introduction of the euro. We discuss these issues in Section III.5, following a brief description of the data in Section III.4 below.

III.4 Data

In order to perform our estimations, we use two data sets. The original sample, approximately equivalent to the one used by CM contains data for import unit values (in local currency), exchange rates (relative to US dollar) and world prices (denominated in US dollars) for 1-digit SITC sectors for 11 countries. As noted in the previous section, we concentrate on looking at the integrated market specification, although analogous results may be derived under 'segmented' markets, where the index of world price (or unit values) is constructed as a weighted (by trade shares) geometric average of prices of each country's five largest trading partners.

The CM data set covers the years 1989-2001 and serves mainly to illustrate that the change of methodology would also result in changes in the inference of the original CM paper.

Results of the estimations for this sample are presented in Appendix B. More important for our specific goals we use the sample of 1995-2005, from Eurostat, which has the advantage of extending further beyond the suspected break date related to the introduction of the euro than the previous data set. The construction of the variables follows CM, and is described in the Appendix A.

The indicator we use for import prices, the index of import unit values (IUV) has a series of caveats concerning its use that must be kept in mind. First of all, unit values, as provided by Eurostat are values of kilograms of a certain good. This means we are looking for instance not only at kilograms of food, oil or raw materials, but also kilograms of computers, cars etc. Moreover, following CGM, we consider the 1-digit SITC industries as a reasonable compromise between the informative power of the series and their availability. Using IUVs, means the 'goods' we speak of are not well defined goods as such - they are in fact bundles of goods (of all goods that are traded on the certain month and fall into the specific SITC category) and thus the composition of such bundles may change from month to month (apart from being different from country to country). Additionally, this composition may change precisely because of changes in the exchange rate, as the demand (and supply) and thus the pricing strategy of some specific sub-category goods may be very different especially within categories as wide as SITC 8 Misc. Manufactured goods. Thus the part of the adjustment to the exchange rate change that will go through quantity and not price, will affect the implicit weight of the good in our 1-digit SITC basket.

Further following from our discussion in Section III.2, it is important to emphasize that there are a number of reasons why we expect there may be a change in the long run ERPT within our sample.

Firstly, on the 1^{st} of January 1999 11 European countries fixed their exchange rates by adopting the euro.[4] This constituted a change in monetary policy, especially for countries were it was previously less credible. The perceived stabilization of monetary policy, especially in countries with previously rather less successful monetary policy, may have induced the producers to change their pricing strategies, and thus have an influence on the ERPT. We expect the formation of the euro area to have caused a change in long run ERPT, though this change may have commenced both before the exact adoption date, for instance upon joining the ERM, as well as after, when the euro became a well established currency.

Anticipating to some extent our future results, on left hand side of Figure 1 we show the errors from the estimation of the levels equation (9), for which as we will see in Section III.5.1 it is not easy to reject the null hypothesis of non-stationarity. On the right hand, we have the residuals from the same equation once we allow for a break - these seem to appear more stationary. The substantial changes in the behavior of the residuals commence, as may be noted in the figure, in the run-up to the euro. Similar figures may be constructed e.g. for France which again shows significant change around the end of 1998. This goes somewhat ahead of our argument, to which we will return to it in more detail in Section III.5.2, but serves for the purpose of illustrating that not accounting for a structural break in the relationship may lead us to the failure of finding a long run, although we must be constantly vigilant that what we classify as a 'break' is not a data artifact. We have good reasons for believing this not to be the case.

Moreover the adoption of a common currency has changed the competitive conditions, by increasing the share of goods denominated in the (new) domestic currency. Turning to Table 2 we can roughly assess the importance of imports originating from outside the euro area in overall imports for each single country and industry. Overall, the share of the

[4]Greece failed to fulfil the Maastricht Treaty criteria, and therefore joined 2 years later, effective 1^{st} of January 2001.

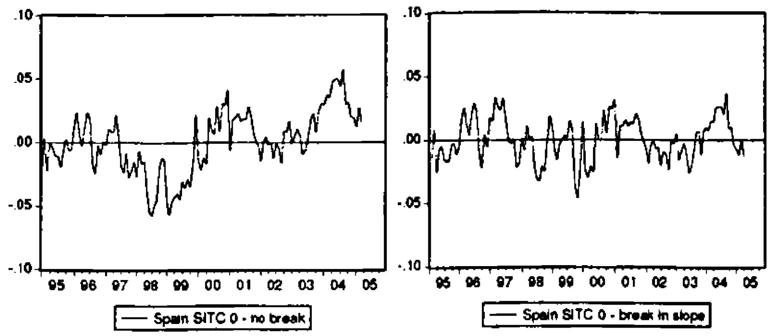

Figure 1: Residuals from the estimation of equation (9) without a break (left) and with a single estimated break (right) on the series for Spain, SITC0.

value of imports from outside the euro area is below 50% in most cases. Moreover it has slightly decreased in the past decade. There are however noticeable exceptions: the share of extra euro area imports is high and often above 90% in case of SITC 3 (Mineral fuels) and relatively high though substantially lower (closer to 60-70%) in the case of SITC 2 (Crude materials). As for countries we immediately notice the high share of extra-euro imports in Ireland, which still does the majority of its trade with the UK and to a smaller extent Finland which trades intensively with other Scandinavian countries and with Russia. Therefore in most cases, joining the euro shifted the composition of on average 50% of imports from foreign currency denominated to home currency denominated, and thus immune to exchange rate fluctuations.

These cautions having been stated, it remains the case that we are constrained in our investigations by the quality of the publicly available data. While there may be numerous doubts about using IUVs as a proxy for import prices, the lack of alternative measures (especially at a sectoral level) forces us to use what is available. This has the advantage that we can make comparisons with the CM or CGM estimates which are based on similarly constructed data.

Finally, looking at the exchange rates of current euro area currencies in Figure 2 we see

Imports shares

		France	Belg.	Nether.	Germ.	Italy	Ireland	Greece	Port.	Spain	Finland	Austria
SITC_0	1995	0.4	0.3	0.43	0.36	0.35	0.8	0.26	0.45	0.53	0.61	0.24
	2000	0.35	0.32	0.44	0.36	0.33	0.7	0.26	0.35	0.5	0.57	0.21
	2005	0.33	0.32	0.45	0.35	0.37	0.74	0.34	0.29	0.51	0.54	0.27
SITC_1	1995	0.24	0.19	0.39	0.31	0.2	0.59	0.47	0.34	0.57	0.38	0.23
	2000	0.25	0.17	0.42	0.34	0.17	0.47	0.44	0.41	0.55	0.37	0.27
	2005	0.25	0.12	0.32	0.32	0.15	0.71	0.4	0.31	0.36	0.33	0.25
SITC_2	1995	0.57	0.54	0.69	0.6	0.6	0.79	0.69	0.69	0.65	0.81	0.48
	2000	0.53	0.55	0.65	0.59	0.63	0.73	0.7	0.61	0.66	0.86	0.52
	2005	0.48	0.43	0.68	0.52	0.6	0.81	0.79	0.49	0.64	0.78	0.46
SITC_3	1995	0.86	0.31	0.89	0.68	0.91	0.95	0.93	0.87	0.92	0.91	0.8
	2000	0.88	0.31	0.85	0.72	0.96	0.92	0.96	0.78	0.92	0.97	0.67
	2005	0.75	0.36	0.77	0.71	0.93	0.97	0.97	0.75	0.9	0.96	0.48
SITC_4	1995	0.24	0.38	0.59	0.52	0.43	0.71	0.39	0.2	0.47	0.62	0.13
	2000	0.24	0.26	0.59	0.54	0.35	0.66	0.4	0.11	0.56	0.66	0.08
	2005	0.3	0.18	0.63	0.48	0.38	0.72	0.58	0.1	0.56	0.67	0.12
SITC_5	1995	0.38	0.36	0.42	0.41	0.34	0.72	0.31	0.23	0.31	0.49	0.24
	2000	0.38	0.37	0.47	0.41	0.34	0.73	0.29	0.21	0.33	0.5	0.27
	2005	0.39	0.26	0.45	0.36	0.34	0.7	0.29	0.19	0.34	0.45	0.3
SITC_6	1995	0.27	0.51	0.39	0.45	0.44	0.77	0.37	0.24	0.3	0.57	0.25
	2000	0.29	0.57	0.41	0.49	0.48	0.72	0.43	0.25	0.32	0.59	0.3
	2005	0.28	0.56	0.48	0.47	0.52	0.75	0.49	0.25	0.37	0.58	0.29
SITC_7	1995	0.4	0.39	0.56	0.51	0.36	0.83	0.37	0.27	0.3	0.6	0.25
	2000	0.43	0.4	0.67	0.59	0.39	0.76	0.46	0.28	0.3	0.59	0.37
	2005	0.42	0.41	0.69	0.58	0.4	0.72	0.47	0.22	0.34	0.59	0.35
SITC_8	1995	0.48	0.37	0.56	0.63	0.59	0.81	0.33	0.24	0.42	0.64	0.27
	2000	0.49	0.47	0.59	0.67	0.61	0.8	0.42	0.2	0.45	0.68	0.31
	2005	0.47	0.48	0.59	0.68	0.63	0.84	0.42	0.16	0.5	0.64	0.33

Table 2: Share of extra-Eurozone imports value in total imports value, by importing country and industry, 1995 and 2005. Source: COMEXT.

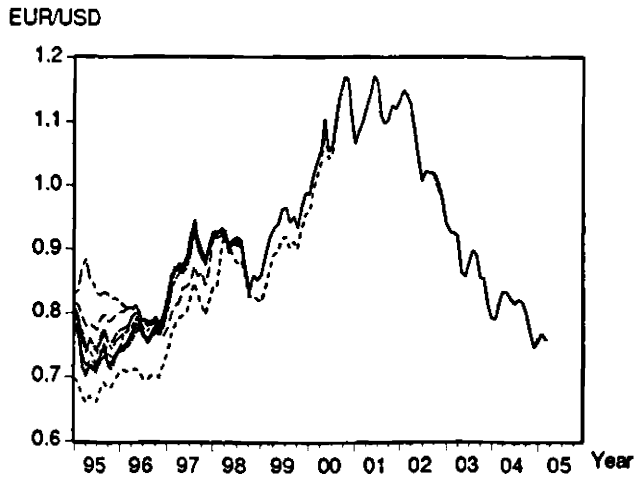

Figure 2: Monthly index of exchange rates of euro area currencies versus the USD. 1995-2005.

that in virtually all the countries the currencies were depreciating against the US dollar in the period 1995-2000, and especially since 1996. Moreover, after a short period of a stable euro dollar exchange rate, the euro currency(ies) started appreciating, till the end of our sample. This asymmetry of exchange rate developments may have different implications for the ERPT, as obviously for an imported good of a fixed dollar price, depreciation of the euro vis a vis the dollar would mean the increase of the price of the good on the euro area market, while the appreciation of the euro, a decrease of the price, leading to possibly different behavior of the producers margin.

III.5 Results

III.5.1 Single equations - without breaks (importance of lag length selection)

Simple augmented Dickey-Fuller tests for cointegration in single time series for individual country/industry combinations (see Tables 3 and 4) do not support the CM view about the lack of cointegration between the series. The results concern the more recent sample (1995-2005) yet by switching to automatic lag selection criteria we manage to obtain rejections of the null of no cointegration for the majority of the series (at 5% level). Moreover, as we see in Table 11 in Appendix B, adopting an information criteria chosen lag length when testing the null on the 1989-2001 CM sample, leads to the rejection of the null of no cointegration for most of the series. Therefore we can say there is some evidence that a long run, levels relationship, in the Engle and Granger (1987) sense, exists between our variables.

III.5.2 Single equations, with structural breaks

In order to pursue the issue of looking for cointegrating relationships further, we propose the use of the Gregory and Hansen (1996, GH hereafter) algorithm which allows for testing the null of no cointegration against the alternative of cointegration with a (estimated) structural break. We test two alternative versions of the model proposed in equation (9). First, a break in the constant, thus a levels shift:

$$mp_t = \widehat{\alpha} + \widehat{\alpha}_1 * d_s + \widehat{\beta} \cdot er_t + \widehat{\gamma} \cdot fp_t + \varepsilon_t \tag{12}$$

Second, a break in all the cointegrating equation coefficients, thus a slope shift:

$$mp_t = \widehat{\alpha} + \widehat{\alpha}_1 * d_s + \widehat{\beta} \cdot er_t + \widehat{\beta}_1 \cdot er_t * d_s + \widehat{\gamma} \cdot fp_t + \widehat{\gamma}_1 \cdot fp_t * d_s + v_t \tag{13}$$

In both cases d_s is a dummy variable equal to 0 if $t < s$ and equal to 1 otherwise. The GH allows for the estimation of the break point s positioning it where the ADF test on errors from the estimated levels equation yield the strongest evidence for the rejection of the null hypothesis of no cointegration.[5]

It is an issue of considerable interest to decide which formulation of the model to adopt. We provide evidence below to show that it is the second of the two formulations that we would tend to choose. Generally, as mentioned earlier, upon the introduction of the euro,

[5]Brief details of the procedure are contained in Appendix C.

H_0: Unit root (no cointegration)

Country	SITC0	SITC1	SITC2	SITC3	SITC4	SITC5	SITC6	SITC7	SITC8
France	-2.75***	-2.69***	-3.06***	-3.11***	-1.85*	-5.51***	-5.47***	-3.65***	-4.12***
Netherlands	-2.51**	-3.15***	-3.3***	-2.97***	-2.61***	-3.14***	-2.1**	-3.42***	-6.47***
Germany	-1.51	-2.29**	-2.46**	-4.94***	-3.77***	-3.79***	-2.21**	-6.65***	-5.7***
Italy	-1.93*	-1.69*	-2.45**	-2.72***	-4.11***	-3.41***	-2.3**	-2.8***	-1.92*
Ireland	0.2	-2.13**	-4.98***	-5.26***	-2.21**	-9.31***	-2.98***	-5.06***	-4.21***
Greece	-1.82*	-1.93*	-2.28**	-2.73***	-2.06**	-4.29***	-2.91***	-3.47***	-3.08***
Portugal	-2.48**	-1.82*	-2.9***	-4.31***	-1.59	-9.13***	-2.16**	-4.24***	-2.66***
Spain	-1.95**	-2.27**	-3.65***	-3.8***	-2.3**	-4.02***	-2.15**	-3.62***	-5.55***
Finland	-0.76	-7.76***	-2.2**	-2.62***	-3***	-3.91***	-3.47***	-2.75***	-6.36***
Austria	-6.49***	-2.26**	-3.53***	-4.65***	-2.43**	-8.14***	-2.9***	-2.63***	-6.44***

For each sector first line reports the ADF t-statistic, and the second p-value.

Specification: no constant, no trend. Maximum lags number = 12. Lag selection: Akaike(AIC).

Table 3: ADF tests on the errors from the OLS regression of the "long run" equation (9). Sample: 1995-2005. For the equivalent result for the CM 1989-2001 sample we refer the reader to Appendix B.

H_0: Unit root (no cointegration)

Country	SITC0	SITC1	SITC2	SITC3	SITC4	SITC5	SITC6	SITC7	SITC8
France	-2.75***	-2.59**	-3.64***	-2.71***	-1.32	-5.51***	-3.95***	-3.65***	-4.12***
Netherlands	-2.87***	-2.55**	-3.08***	-2.97***	-4.09***	-3.14***	-2.88***	-3.42***	-2.97***
Germany	-1.51	-2.29**	-3.09***	-4.94***	-2.34**	-3.79***	-2.39**	-3.67***	-2.13**
Italy	-1.93*	-1.69*	-1.5	-2.57**	-5.07***	-3.41***	-2.3**	-1.99**	-2.87***
Ireland	0.2	-6.47***	-2.84***	-3.97***	-2.21**	-4.27***	-2.98***	-3.19***	-4.21***
Greece	-1.82*	-3.67***	-2.28**	-2.73***	-2.59**	-3.96***	-3.12***	-3.47***	-3.08***
Portugal	-2.48**	-1.82*	-2.85***	-4.31***	-1.59	-9.13***	-2.4**	-4.2***	-2.56**
Spain	-2.24**	-2.19**	-2.17**	-2.36**	-2.3**	-1.93*	-2.35**	-2.91***	-1.79*
Finland	-0.95	-3.7***	-2.76***	-2.91***	-3.86***	-3.91***	-3.47***	-2.97***	-2.13**
Austria	-4.72***	-2.27**	-4.13***	-2.27**	-2.43**	-8.14***	-3.15***	-3.99***	-2.88***

For each sector first line reports the ADF t-statistic, and the second p-value.

Specification: no constant, no trend. Maximum lags number = 12. Downward t selection.

Table 4: ADF tests on the errors from the OLS regression of the "long run" equation (9). Sample: 1995-2005. The downward-t selection method essentially starts with a general specification with the maximum lags and deletes lags one-by-one, re-estimating and selecting the number of lags where the last lag is significant(t-statistic).

Industry	"Long run" exchange rate pass-through coefficients											
	France		Netherlands		Germany		Italy		Ireland		Greece	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
SITC_0	(.)		(.)		(.)		(.)		0.52	6/03	0.29	2/02
									(0.08)	+***	(0.03)	+***
SITC_1	(.)		0.65	7/98	(.)		0.15	10/00	0.09	11/00	(.)	
			(0.05)	-***			(0.06)	+***	(0.09)	-***		
SITC_2	1.08	1/97	(.)		(.)		1.15	3/97	0.6	9/03	(.)	
	(0.04)	-***					(0.03)	-***	(0.05)	+***		
SITC_3	0.99	1/01	(.)		1.1	8/96	(.)		0.6	6/00	1.26	10/02
	(0.01)	-***			(0.02)				(0.06)	+***	(0.06)	+***
SITC_4	0.9	2/00	1.24	1/99	0.95	*12/00*	0.21	10/96	(.)		(.)	
	(0.09)	+***	(0.04)	-***	(0.05)	-***	(0.12)	-**				
SITC_5	0.4	2/01	0.56	12/02	0.52	*1/03*	0.68	4/97	0.46	12/97	0.48	2/03
	(0.03)	+***	(0.03)	-***	(0.02)	-***	(0.04)	-***	(0.04)	+	(0.02)	-***
SITC_6	0.81	4/01	0.95	11/01	0.75	*3/98*	(.)		0.59	5/98	0.7	*4/02*
	(0.02)	+***	(0.02)	-***	(0.01)	+***			(0.03)	+***	(0.03)	+***
SITC_7	0.55	7/98	0.89	9/98	0.51	1/01	0.45	12/97	(.)		(.)	
	(0.02)	-***	(0.03)	-***	(0.02)	+**	(0.02)	-***				
SITC_8	0.72	7/96	0.86	9/01	(.)		(.)		0.61	7/96	(.)	
	(0.01)	+***	(0.03)	-***					(0.04)	+***		

Value not reported (.) if the hypothesis of unit root (no cointegration) cannot be rejected at 10%
(Italics) if the hypothesis of unit root (no cointegration) can be rejected at 10%, but cannot be rejected at 5%;
For each country/industry combination columns: (1) first row estimated coefficient on ER, second row standard error, (2) first row estimated break date, second row direction of change in constant
+ or − indicate change in constant is positive or negative,
,,*** indicate this change is significant (t-stat of $\widehat{\alpha}_1$) at 10%, 5% and 1%*

Table 5: Estimated ERPT coefficients and break dates extracted from GH. Specification: break in constant - equation (12). Statistic: ADF*. Sample: 1995-2005.

Industry	"Long run" exchange rate pass-through coefficients							
	Portugal		Spain		Finland		Austria	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
SITC_0	0.64	6/03	(.)		0.79	2/03	0.69	4/98
	(0.06)	+***			(0.05)	-***	(0.04)	+
SITC_1	(.)		0.73	9/03	0.74	1/03	(.)	
			(0.06)	+***	(0.03)			
SITC_2	0.68	*1/99*	0.97	3/98	(.)		0.72	1/98
	(0.05)	-***	(0.03)	-***			(0.02)	+***
SITC_3	0.99	*10/02*	(.)		(.)		0.87	12/97
	(0.03)	-***					(0.03)	+***
SITC_4	1.32	12/02	(.)		0.13	9/97	(.)	
	(0.16)	+***			(0.09)	+***		
SITC_5	0.7	3/98	0.74	6/97	0.56	1/03	0.4	4/98
	(0.07)	-**	(0.04)	-***	(0.03)	-***	(0.05)	-***
SITC_6	0.73	*8/02*	0.86	11/01	(.)		0.44	9/00
	(0.03)	-***	(0.02)	-***			(0.02)	+***
SITC_7	0.36	6/03	(.)		(.)		0.25	3/02
	(0.02)	-***					(0.02)	-***
SITC_8	0.78	4/99	0.86	11/98	0.35	6/97	0.57	9/97
	(0.06)	-***	(0.03)	-***	(0.03)	+***	(0.03)	+**

Notes: See previous page for explanations.

Table 5: continued.

Industry	"Long run" exchange rate pass-through coefficients											
	France			Netherlands			Germany			Italy		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.83	1	6/00	0.72	0.51	2/99*	(.)	(.)		(.)	(.)	
	(0.03)	(0.11)	-	(0.03)	(0.1)	+						
SITC_1	0.51	1.96	8/02***	0.82	0.5	1/98**	(.)	(.)		0.13	0.91	10/00***
	(0.05)	(0.37)	-***	(0.08)	(0.12)	+				(0.06)	(0.27)	-***
SITC_2	(.)	(.)		(.)	(.)		0.72	0.9	2/99***	0.83	1.11	11/97***
							(0.03)	(0.03)	-	(0.07)	(0.03)	+**
SITC_3	1.03	1.09	1/01*	(.)	(.)		(.)	(.)		(.)	(.)	
	(0.01)	(0.03)	-***									
SITC_4	1.04	0.48	2/00***	0.96	1.34	4/99***	0.95	0.79	12/00	-0.25	0.71	7/99***
	(0.11)	(0.12)	+**	(0.06)	(0.06)	-***	(0.06)	(0.11)	+	(0.22)	(0.15)	-***
SITC_5	0.35	0.9	1/01***	0.52	1.22	9/02***	(.)	(.)		0.42	0.75	9/98***
	(0.05)	(0.14)	-***	(0.04)	(0.25)	-***				(0.1)	(0.06)	+
SITC_6	0.82	0.83	*7/01*	(.)	(.)		(.)	(.)		(.)	(.)	
	(0.02)	(0.04)	+									
SITC_7	0.38	0.56	7/98**	0.96	0.9	9/98	0.47	1.1	10/00***	0.41	0.44	1/98
	(0.08)	(0.02)	-**	(0.13)	(0.04)	+	(0.02)	(0.11)	-***	(0.05)	(0.02)	+*
SITC_8	0.74	0.72	7/96	0.87	1.33	8/01**	(.)	(.)		0.4	0.86	1/98***
	(0.11)	(0.01)	+	(0.05)	(0.2)	-**				(0.05)	(0.03)	-***

Value not reported (.) indicates H_0: unit root (no cointegration) cannot be rejected at 10%
(Italics) indicate H_0 cannot be rejected at 5%, but can be rejected at 10%;
For each country and industry combination columns: (1) first row coefficient on ER before break date, second row standard error, (2) first row coefficient on ER after break date, second row standard error, (3)first row: estimated break date, second row direction of change in constant + positive, - negative,
,,*** indicate the change of the long run pass-through (first row) or constant (second row) on the date of the break is significant (t-stat of $\widehat{\beta}_1$ or $\widehat{\alpha}_1$ respectively) at 10%, 5% and 1%.*

Table 6: Estimated coefficients and break dates extracted from GH. Specification: break in cointegrating vector. Statistic: ADF*. Sample: 1995-2005.

Industry	"Long run" exchange rate pass-through coefficients											
	Ireland			Greece			Portugal			Spain		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.44	0.89	4/02*	0.24	0.27	8/01	0.57	1.68	7/02***	0.67	1.28	9/99***
	(0.06)	(0.25)	-***	(0.03)	(0.2)	-	(0.05)	(0.3)	-***	(0.03)	(0.1)	-***
SITC_1	-0.13	-0.43	2/03	(.)	(.)		0.49	2.47	1/00***	(.)	(.)	
	(0.07)	(0.82)	+				(0.13)	(0.56)	-***			
SITC_2	0.58	0.75	9/03	(.)	(.)		0.82	0.58	*1/99***	0.96	0.96	3/98
	(0.05)	(0.35)	-				(0.08)	(0.06)	+	(0.06)	(0.03)	-
SITC_3	0.87	0.83	5/03	1.12	-0.18	4/00***	0.86	1.03	11/98**	(.)	(.)	
	(0.05)	(0.43)	-	(0.09)	(0.13)	+***	(0.07)	(0.03)	-			
SITC_4	(.)	(.)		0.68	1.08	12/00	1.39	0.78	12/02	(.)	(.)	
				(0.11)	(0.23)	-	(0.17)	(0.46)	+			
SITC_5	0.27	0.47	12/97	0.46	0.5	11/98	0.32	0.76	8/98**	0.46	0.73	2/97
	(0.14)	(0.04)	-	(0.04)	(0.03)	+	(0.15)	(0.08)	-***	(0.29)	(0.04)	-
SITC_6	0.56	0.93	8/02***	(.)	(.)		(.)	(.)		0.85	1.18	11/01***
	(0.03)	(0.11)	-***							(0.02)	(0.06)	-***
SITC_7	(.)	(.)		0.35	0.37	*1/98*	0.4	0.15	1/02	(.)	(.)	
				(0.12)	(0.03)	-	(0.03)	(0.27)	+			
SITC_8	0.47	1.04	*5/00***	(.)	(.)		0.57	0.93	*5/99***	0.99	1.09	12/99
	(0.1)	(0.23)	-***				(0.11)	(0.08)	-**	(0.07)	(0.1)	-

Notes: See previous page.

Table 6: continued.

Industry	"Long run" exchange rate pass-through coefficients					
	Finland			Austria		
	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.86	0.78	7/99	0.69	0.79	2/98
	(0.04)	(0.14)	+	(0.05)	(0.12)	–
SITC_1	0.72	0.95	9/01	(.)	(.)	
	(0.04)	(0.2)	–			
SITC_2	(.)	(.)		0.77	0.72	*1/97*
				(0.06)	(0.03)	+*
SITC_3	(.)	(.)		0.75	0.88	12/96
				(0.1)	(0.03)	–
SITC_4	-0.17	0.14	9/96	(.)	(.)	
	(0.33)	(0.1)	–			
SITC_5	0.47	0.72	4/02	0.21	0.46	2/98*
	(0.03)	(0.27)	–**	(0.1)	(0.07)	–**
SITC_6	0.72	0.75	12/99	0.43	0.51	9/99**
	(0.02)	(0.04)	–***	(0.02)	(0.03)	–
SITC_7	(.)	(.)		0.27	0.23	*5/01*
				(0.03)	(0.25)	+
SITC_8	0.59	0.49	4/99	(.)	(.)	
	(0.06)	(0.1)	|			

Notes: see previous page

Table 6: continued.

we would expect the fixed component of the mark-up (denoted by the coefficient α) to rather fall than increase - due to potentially improved competition in the market arising from increased price transparency. Table 5 (for the GH single equation tests) shows clearly that in the specification of break only in constant, the fixed component in the mark-up tends to rise roughly in as many cases as it tends to fall. However as the specification from equation (12) is much more restrictive than the one based on equation (13), not allowing for a possible break in the other variables would tend to cause the estimate of α_1 to be biased. Table 6 shows that the more flexible specification of a break in slopes and constant lead to the majority of the estimates pointing to a decrease or insignificant change in the fixed mark-up component.[6] In more detail, when we allow for the more general break as in equation (13), for the GH single equations for 22 out of 60 series α_1's are negative and significant, leaving only 5 significantly positive.

Comparing the results for the two alternative specifications (with breaks in constant and with breaks in constant and slope) we see that in a handful of cases the rejection of the null of no cointegration was possible when the alternative did not allow for a break, while not possible when the alternative accounted for a break. This tends to suggest, that in these cases the evidence for the existence of a break is weak.[7] Overall the most important

[6]Significance is judged with respect to conventional critical values. A more general consideration would require bootstrapped critical values and is the subject of on-going investigations.

[7]If for a certain series we are able to reject the null of no cointegration against an alternative of cointegration without a break (ADF), but unable to do so against an alternative with a break (GH), this may be evidence that there is no break in the cointegrating relationship. The reasoning is as follows. We treat rejection of the null of no cointegration (ADF) as evidence of existence of a cointegrating relationship between the variables, as in equation (9). In this case, imposing a break, i.e. a dummy variable as in equation (12), or a dummy variable with interaction variables as in equation (13), would mean adding variables of no explanatory power (insignificant) and should not in principle affect our statistic. However, the critical values of the GH tests are higher in absolute value than those of the standard ADF test, in order to guarantee appropriate test sizes for the null of no cointegration against of an alternative of cointegration with a break. Thus a more or less unaffected test statistic and a higher critical value may result in the failure to reject the null in the case when imposing a break is not justified.

outcome is that in a relatively short sample, there are only about 12 out of 90 series for which we are unable to reject the null of no cointegration in any of the three specifications (no break, break in constant, breaks in slopes). We treat this as strong evidence of the presence of the theory backed levels relationship in our data, which changes in response to key economic events.[8]

A selection of the single equation results for the "long run" ERPT are presented in Figures 3 to 6. As indicated in the notes to the figures, they present the point estimate and the 95% confidence interval for both the CM-defined long run (estimator (1) in all the figures) i.e. the sum of five lags, as well as the EG long run in 5 different specifications. Noticeably, apart from yielding different values of the pass-through, the EG estimates are more precise which allows for more definite conclusions regarding the rejection or acceptance of the hypotheses of ERPT being equal to 0 or 1. The narrower confidence intervals are an immediate consequence of the superconsistency of the OLS estimator in a cointegrating relationship. The coefficients obtained from the levels estimation of equation (9) when allowing for a structural break in the entire cointegration vector (observations (4) and (5) for the GH estimated break and (7), (8) for the imposed 1998/1999 break) are however more imprecise, especially if the estimated break happens to lie towards the beginning or end of the sample.

There is some country- and industry-specific variety in long run pass-through, where commodity sectors (SITC 2 and SITC 3) tend to have a higher (closer to 1) pass-through than manufacturing sectors, and with very few exceptions we can strongly reject zero rates of pass-through. A glance at the tables and figures also suggests, if anything, an increase in

[8]The changes are modelled here as discrete breaks in constant or slope and is a limitation of our framework. A richer alternative to consider would be allow for non-linearities, which may in fact pick up evidence for more change which is gradual. This is unfortunately precluded in our study by the shortage of data.

the pass-through rates in most countries and most industries, with some exceptions. Not all of these changes are significant, but the tendency is nevertheless rather clear cut.

Overall, tests for cointegration, be it without a break, with a break in the constant, or in the entire "equilibrium" relationship allow us to reject the null of no cointegration therefore providing support for the existence of a long run relationship as in equation (9) in our data. This stands somewhat in contrast with the CM conclusion that no cointegrating relationship exists, and allows us to switch from an arbitrary definition of long run ERPT as a sum of five (mostly insignificant) coefficients on the lags of the exchange rate to the long run in the EG sense.

The evidence gathered above, by looking at individual sectors within each country, can be strengthened even further by using several recently developed panel-based tests for cointegration. Dealing with single time series, albeit with about 110-120 observations, we still have a time span of only about 10 years of data. However, by looking at the evidence from all the sectors and countries together (if the number of sectors in each country, is 9 and there are 10 countries in our data set, a panel-based test could use up to $9 \times 10 \times 110$ observations) and allowing for heterogeneity, we should in principle obtain a far clearer idea of the common trends underlying the series and hence the existence of the long run. In the spirit of the discussion above, any such estimation procedure in panels would of course need to allow for structural change. In addition it would also need to allow for dependence among the units of the panel. We turn now to a consideration of these issues.

III.6 Panel cointegration tests

There are essentially three ways of proceeding in order to construct panels from the data sets - (1) creating country panels of industry cross-sections, (2) industry panels with country cross-sections and (3) a pooled panel in which every country and industry combination

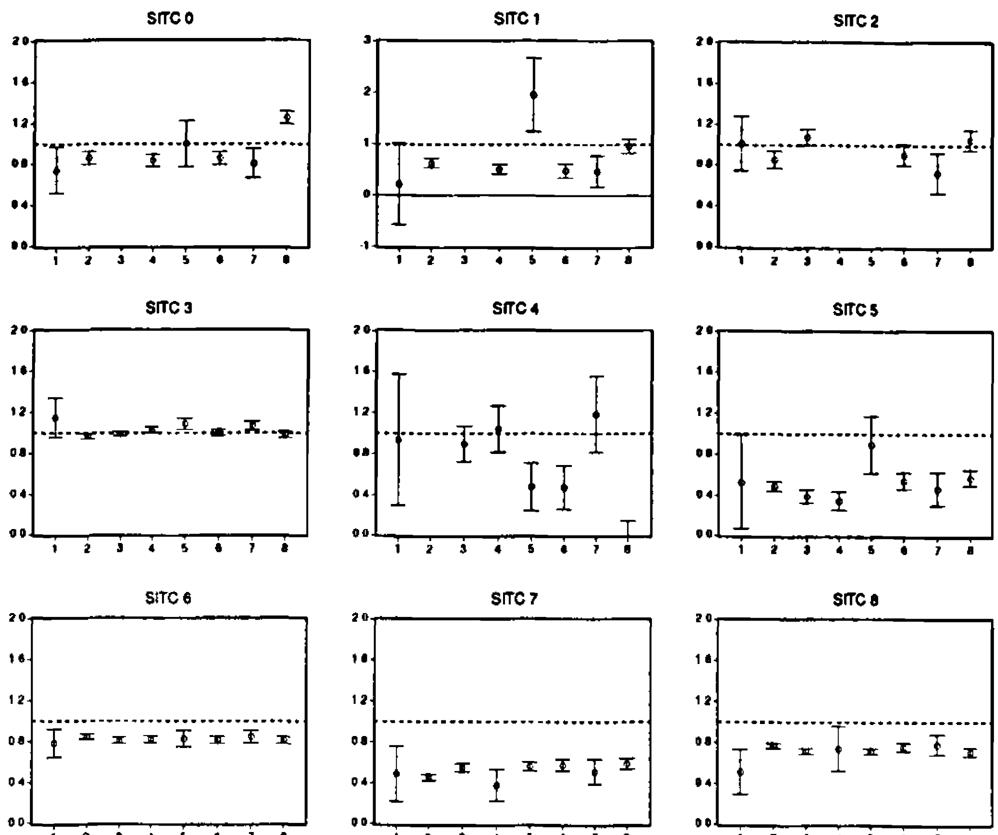

Figure 3: France - "long run" exchange rate pass-through estimates with confidence intervals (95%). Individual industries, sample: 1995-2005, entire sample analysis. The estimators are presented in the following order: (1) CM long run, no cointegration, no break, (2) cointegrating long-run, no break, (3) cointegrating long run, break in constant (GH), (4) cointegrating long run before break in slope(GH), (5) cointegrating long run, after break (GH), (6) cointegrating long run, break in constant (1998/99), (7) cointegrating long run, before break in slope (1998/99), (8) cointegrating long run, after break (1998/99). In (3)-(5) values extracted from Gregory and Hansen (1996), ADF*. Values not reported if no cointegration (ADF). For break dates estimated with the GH refer to Table 5. Dotted horizontal line at value of 1.

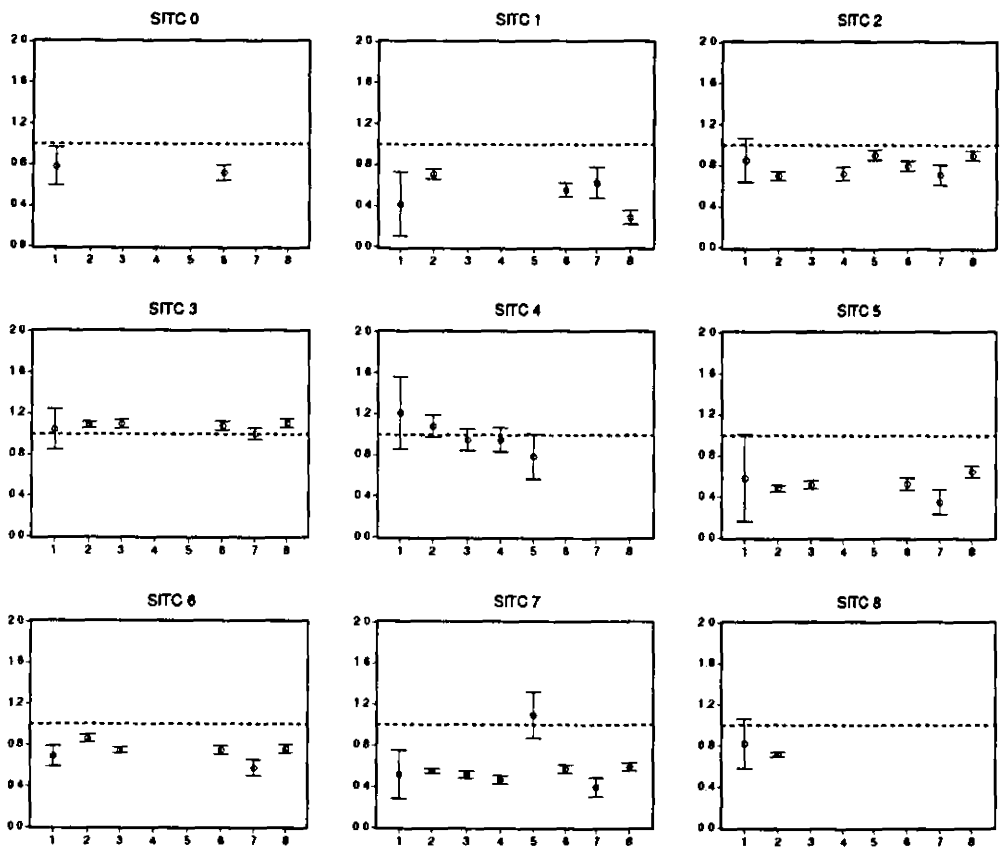

Figure 4: Germany - "long run" exchange rate pass-through estimates with confidence intervals. Individual industries, sample: 1995-2005. Notes - see Figure 3 for explanations.

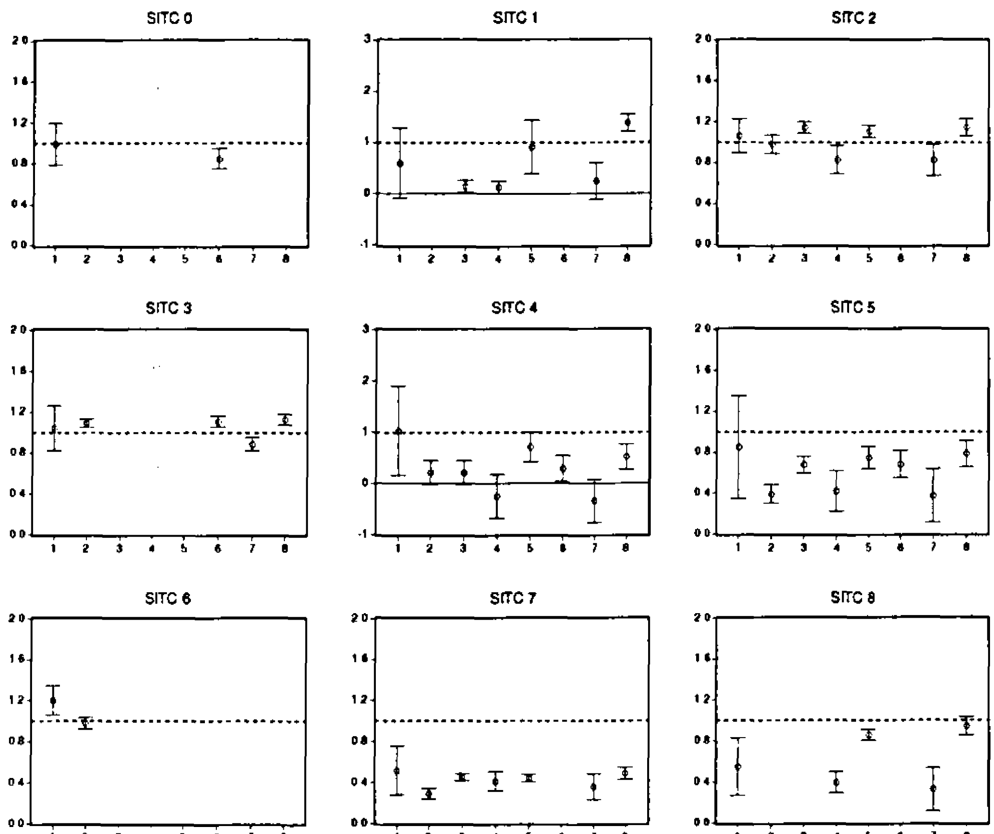

Figure 5: Italy - "long run" exchange rate pass-through estimates with confidence intervals. Individual industries, sample: 1995-2005. Notes - see Figure 3 for explanations.

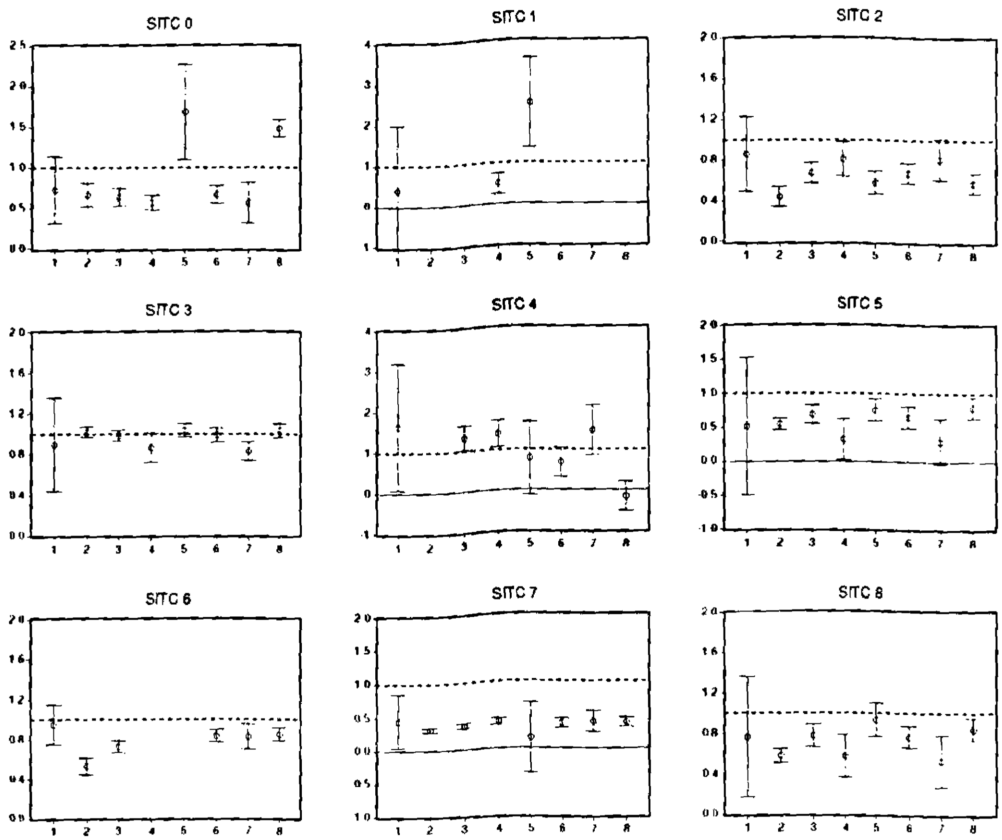

Figure 6: Portugal - "long run" exchange rate pass-through estimates with confidence intervals. Individual industries, sample: 1995-2005. Notes - see Figure 3 for explanations.

constitutes a separate unit. In search of the existence of a cointegrating relationship in the series we try to maximize the dimensions of our panel, and thus will focus on (3). Hence we will apply two types of tests. The so called first generation panel cointegration tests as in Pedroni (1999) test for existence of a cointegrating relationship, assuming no cross-unit interdependence. The modification of the test, based on Gregory and Hansen (1996) is proposed in Banerjee and Carrion-i-Silvestre (2006) and allows for an estimated breakpoint in each individual series. As mentioned however, the tests have the shortcoming of not accounting for possible cross unit dependence. This, as shown by Banerjee, Marcellino and Osbat (2004) in a series of Monte Carlo simulations, can lead to substantial oversize of the tests, and thus increase the possibility of wrongful rejection of the null of no cointegration.

The second generation of tests, as the one proposed in Banerjee and Carrion-i-Silvestre (2006) allows a factor structure for cross-section dependence, but has the limitation of imposing a common (across units) break date.[9] The latter, as we will see, does not seem as much of the problem, as in our main sample of interest, the Pedroni (1999) break date estimates are, as we suspected, relatively close to the date of the introduction of the common currency. Moreover, recomputing the tests for different (imposed) break dates does not lead to any substantial change in our conclusions.

The statistics for the Pedroni (1999) panel cointegration tests with no cross-sectional dependence and no breaks are displayed in the first row of Table 7. They allow for strong rejection of the hypothesis of no cointegration even when the alternative does not allow for a break. This test is restrictive in the sense that we do not allow the cointegration relationship to change within our sample. However as mentioned, we suspect the formation of the euro area constituted a shift in both competition conditions and monetary policy which may

[9]Brief details of these tests are contained in Appendix C.

Model	pseudo-t	pseudo-ρ
No break	-7.73	-35.45
Break in constant	-22.15	-49.38
Break vector	-23.26	-49.50
Under the null hypothesis both statistics have a N(0,1) distribution		

Table 7: Pseudo-t and Pseudo-ρ parametric statistics from Banerjee and Carrion-i-Silvestre (2006). The null hypothesis is no cointegration. Sample: 1996-2004, full panel (N=90), unit specific breaks.

have affected the long run pass-through. We propose running the Pedroni (1999) test which allows for the change in the cointegrating vector. The results allow strong rejection of the null of no cointegration in both the case of a shift in constant and break in the cointegrating relationship between the variables for all the country panels. By construction the test chooses the break date which is consistent with strongest evidence against the null. The test algorithm allows us to extract the break dates for each individual series, as well as the cointegrating coefficients. These are presented in Tables 8 and 9. Within the context of these results derived from the panel tests, it is useful to return briefly to the issue of model choice and to ask whether the more flexible formulation (i.e. equation (13) instead of (12)) is also the more appropriate here. We note from the panel estimates reported in Table 9 that out of 90 series, 34 α_1's are significantly negative, while only for 6 they are significantly positive. We therefore point to the break in slopes and constant specification as being more coherent with the idea that the fixed component of the markup falls (a negative value of α_1), while changes in the pass-through are also observed for a number of sectors and countries.

The estimated break dates for all the individual series are presented in Figure 7. There is some dispersion among the obtained dates, and though there seem to be two modes of the distribution - one relatively close to the introduction of the euro and the other close to the turn around in the euro/dollar exchange rate developments (2000-2001).

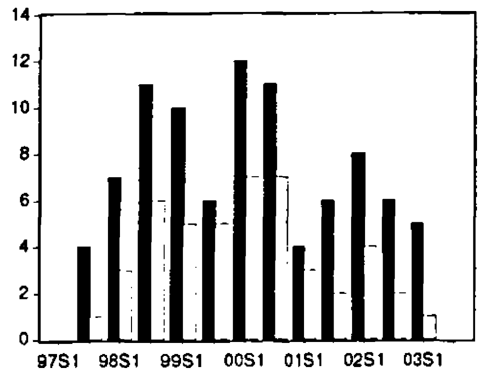

Figure 7: Distribution of estimated break dates by semester (1997s1-2003s2). Breaks in slope taken from Tables 9. Dark color - all breaks, light color - only breaks when long run ERPT changed significantly (10%).

Although the evidence, as presented in Tables 8 and 9, in favor of both cointegration and structural change is unequivocally strong, a few qualifications are worth noting. First, the GH based algorithm here allows for only one, "strongest"[10] break, which is a serious limitation as far as timing the (single) break allowed is concerned. Second, as noted earlier when referring to non-linear methods, the effect of the change in macroeconomic conditions on the ERPT may not have been either instantaneous or linear. Finally, there are other features of this period which are relevant, such as the evolution of the euro/pound rate for Ireland, late euro area membership for Greece etc.

Nevertheless, the sheer fact that despite these limitations (which would in all cases have acted against us) the algorithm identifies a relatively large amount of series where there is cointegration and the change, be it upon the introduction of the euro, or upon the appreciation of the euro, is an interesting finding. Moreover, as we will turn to the

[10]In fact it does not touch upon the notion of the strength of evidence of the break. Generally the break found by this algorithm is a break for which the evidence for a cointegrating relationship is the strongest (i.e. largest - in absolute value - test statistic leading to the rejection of the null of no cointegration).

Industry	"Long run" exchange rate pass-through coefficients											
	France			Netherlands			Germany			Italy		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.88	0.9	12/97	0.63	0.63	8/03	0.89	0.85	3/02	0.98	0.76	9/98
	(0.03)	(0.03)	-***	(0.03)	(0.03)	-***	(0.06)	(0.03)	-***	(0.08)	(0.05)	-***
SITC_1	0.72	0.77	9/03	0.34	0.7	5/98	0.81	0.57	3/98	0.54	0.23	10/00
	(0.06)	(0.07)	-*	(0.06)	(0.08)	-***	(0.03)	(0.04)	+***	(0.07)	(0.08)	+***
SITC_2	0.98	1.06	5/97	0.76	0.81	2/98	0.78	0.85	12/97	0.98	1.1	5/97
	(0.03)	(0.03)	-***	(0.03)	(0.03)	-***	(0.02)	(0.02)	-***	(0.03)	(0.03)	-***
SITC_3	0.96	0.97	3/01	0.85	0.77	7/98	1.1	1.04	9/00	1.08	1.15	6/00
	(0.02)	(0.01)	-***	(0.04)	(0.04)	+***	(0.02)	(0.02)	+***	(0.02)	(0.02)	-***
SITC_4	0.21	0.82	2/00	1.36	1.33	12/98	1.21	1.1	10/01	0.2	0.25	6/97
	(0.14)	(0.11)	+***	(0.05)	(0.05)	-***	(0.06)	(0.06)	-***	(0.13)	(0.13)	-*
SITC_5	0.52	0.68	9/99	0.5	0.57	10/02	0.51	0.55	2/98	0.45	0.66	4/98
	(0.03)	(0.06)	-***	(0.03)	(0.04)	-***	(0.02)	(0.03)	-*	(0.04)	(0.05)	-***
SITC_6	0.86	0.82	7/01	0.85	1.03	1/01	0.85	0.76	3/98	0.99	1.27	10/99
	(0.01)	(0.01)	+***	(0.03)	(0.02)	-***	(0.02)	(0.01)	+***	(0.03)	(0.03)	-***
SITC_7	0.46	0.54	4/98	0.62	0.89	10/98	0.55	0.48	4/00	0.3	0.44	1/98
	(0.02)	(0.02)	-***	(0.03)	(0.04)	-***	(0.01)	(0.03)	+**	(0.02)	(0.02)	-***
SITC_8	0.74	0.7	6/98	0.68	0.86	9/01	0.74	0.78	3/03	0.54	0.83	1/98
	(0.01)	(0.02)	+***	(0.03)	(0.01)	-***	(0.01)	(0.02)	-***	(0.04)	(0.02)	-***

For each country and industry combination columns:
(1) first row: coefficient when no break allowed, second row: standard error;
(2) first row: coefficient if shift in constant allowed, second row: standard error;
(3) first row: estimated shift date for column (2), second row: direction and significance of shift in constant;
*,**,*** *denote shift in constant is significant (t-stat of $\widehat{\alpha}_1$) at 10%, 5% and 1% respectively.*
+ and – denote positive and negative shifts in constant

Table 8: Panel cointegration (Pedroni 1999, modified in Banerjee and Carrion-i-Silvestre, 2006, to account for breaks) results without breaks (equation (9)) and with breaks in constant (equation (12)), reported for group pseudo-t. Cross-section: individual industry in an individual country. No cross-section dependence. Sample: 1996-2004.

Industry	"Long run" exchange rate pass'-through coefficients											
	Ireland			Greece			Portugal			Spain		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.45	0.48	6/03	0.47	0.3	8/01	0.61	0.61	9/03	0.73	0.69	2/00
	(0.12)	(0.08)	+***	(0.06)	(0.04)	+***	(0.08)	(0.05)	+***	(0.05)	(0.04)	+***
SITC_1	-0.57	-0.12	11/00	0.48	0.35	5/02	0.26	0.44	12/02	1.1	0.93	9/03
	(0.09)	(0.11)	-***	(0.03)	(0.03)	+***	(0.12)	(0.14)	-**	(0.08)	(0.08)	+***
SITC_2	0.68	0.59	9/03	0.54	0.36	7/00	0.4	0.74	5/99	0.83	0.97	3/98
	(0.05)	(0.05)	+***	(0.03)	(0.04)	+***	(0.05)	(0.05)	-***	(0.04)	(0.03)	-***
SITC_3	0.72	0.86	5/03	0.99	1.23	7/02	1.04	1.1	3/00	1.14	1.18	6/97
	(0.05)	(0.06)	+***	(0.1)	(0.07)	+***	(0.03)	(0.04)	-**	(0.02)	(0.02)	-***
SITC_4	0.31	0.54	9/01	0.64	0.64	10/01	0.42	1.41	12/02	0.64	0.71	5/99
	(0.11)	(0.11)	+***	(0.14)	(0.11)	+***	(0.24)	(0.19)	+***	(0.09)	(0.09)	+***
SITC_5	0.48	0.51	3/02	0.5	0.54	1/01	0.58	0.72	3/98	0.61	0.76	11/97
	(0.02)	(0.03)	-	(0.01)	(0.02)	-**	(0.05)	(0.07)	-**	(0.03)	(0.04)	-***
SITC_6	0.74	0.61	7/98	0.81	0.69	12/01	0.57	0.73	8/02	0.73	0.85	6/02
	(0.03)	(0.02)	+***	(0.03)	(0.03)	+***	(0.04)	(0.03)	-***	(0.03)	(0.02)	-***
SITC_7	0.56	0.51	10/97	0.51	0.4	11/97	0.35	0.37	8/03	0.31	0.41	10/98
	(0.03)	(0.04)	+*	(0.02)	(0.03)	+***	(0.02)	(0.02)	-**	(0.02)	(0.02)	-***
SITC_8	0.61	0.45	9/99	0.92	0.75	2/02	0.53	0.7	6/98	0.65	0.92	12/99
	(0.04)	(0.08)	+**	(0.03)	(0.04)	+***	(0.04)	(0.05)	-***	(0.03)	(0.06)	-***

Notes: see previous page.

Table 8: continued.

Industry	"Long run" exchange rate pass-through coefficients					
	Finland			Austria		
	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.86	0.92	5/00	0.68	0.69	4/98
	(0.1)	(0.05)	-***	(0.04)	(0.04)	-***
SITC_1	0.72	0.73	3/03	0.17	0.6	3/99
	(0.03)	(0.03)	-***	(0.07)	(0.08)	-***
SITC_2	1.14	0.74	4/99	0.76	0.72	1/98
	(0.07)	(0.06)	+***	(0.02)	(0.02)	+***
SITC_3	1.01	0.87	4/99	0.98	0.87	12/97
	(0.03)	(0.04)	+***	(0.03)	(0.03)	+***
SITC_4	0.25	0.11	9/97	0.04	0.61	11/00
	(0.11)	(0.1)	+***	(0.13)	(0.12)	+***
SITC_5	0.49	0.55	11/02	0.27	0.4	4/98
	(0.03)	(0.03)	-***	(0.03)	(0.05)	-***
SITC_6	0.75	0.7	10/98	0.56	0.44	9/00
	(0.02)	(0.02)	+***	(0.02)	(0.02)	+***
SITC_7	-0.06	0.12	3/02	0.16	0.25	3/02
	(0.03)	(0.03)	-***	(0.02)	(0.02)	-***
SITC_8	0.44	0.36	5/97	0.61	0.57	9/97
	(0.02)	(0.03)	+***	(0.02)	(0.03)	-***

Notes: see previous page.

Table 8: continued.

Industry	"Long run" exchange rate pass-through coefficients											
	France			Netherlands			Germany			Italy		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.84	0.96	6/00	0.62	1.2	11/02***	0.89	0.93	4/00	0.75	1.72	9/00***
	(0.03)	(0.12)	-	(0.03)	(0.18)	-**	(0.03)	(0.11)	+	(0.05)	(0.15)	-***
SITC_1	0.54	1.9	4/02***	0.8	0.59	8/98	0.67	0.42	4/98***	0.21	1.53	3/99***
	(0.07)	(0.36)	-***	(0.08)	(0.12)	-	(0.05)	(0.07)	+**	(0.11)	(0.18)	-***
SITC_2	1.03	0.85	5/02	0.91	0.77	2/98**	0.78	0.88	12/98***	0.89	1.17	5/00***
	(0.03)	(0.13)	+**	(0.06)	(0.03)	-	(0.03)	(0.02)	-	(0.04)	(0.04)	-
SITC_3	0.99	1.13	5/00***	1.13	0.87	6/00**	1.05	1.1	2/99	1.07	1.14	6/00
	(0.02)	(0.02)	-***	(0.07)	(0.07)	+	(0.07)	(0.03)	-	(0.04)	(0.04)	-
SITC_4	1.13	0.45	2/00***	1.08	1.3	4/99**	1.12	0.78	12/00**	-0.58	0.77	6/99***
	(0.16)	(0.13)	+**	(0.08)	(0.05)	-***	(0.08)	(0.12)	+*	(0.22)	(0.13)	-***
SITC_5	0.52	0.79	9/99**	0.58	1.27	2/00***	0.44	0.78	7/00**	0	0.76	12/98***
	(0.08)	(0.08)	-***	(0.08)	(0.19)	-***	(0.04)	(0.14)	-**	(0.12)	(0.06)	-***
SITC_6	0.81	0.81	7/01	1.01	1.12	2/01**	0.66	0.78	1/01***	1.09	1.42	12/99***
	(0.02)	(0.04)	-	(0.02)	(0.04)	-*	(0.02)	(0.04)	-***	(0.04)	(0.04)	-***
SITC_7	0.46	0.58	11/98	1.05	0.94	12/98	0.45	1.12	9/00***	0.35	0.44	1/98
	(0.08)	(0.02)	-	(0.12)	(0.04)	+	(0.02)	(0.12)	-***	(0.07)	(0.02)	+
SITC_8	0.71	0.7	6/98	0.84	1.36	9/01**	0.8	0.41	3/03*	0.58	0.86	1/98***
	(0.05)	(0.02)	+	(0.06)	(0.2)	-**	(0.02)	(0.2)	+**	(0.07)	(0.02)	-***

For each country and industry combination columns:

(1) first row: coefficient on ER before break, second row: standard error; (2) first row: coefficient on ER after break, second row: standard error; (3) first row: estimated shift date with significance of change in coefficient on ER, second row: direction and significance of shift in constant;

*,**,*** *denote shift is significant (t-stat of* $\widehat{\beta}_1$ *or* $\widehat{\alpha}_1$*) at 10%,5% and 1% respectively.*

+ and – denote positive and negative shifts in constant

Table 9: Panel cointegration (Pedroni, 1999, modified in Banerjee and Carrion-i-Silvestre, 2006, to account for breaks) results with breaks in cointegrating vector (equation (13)), reported for group pseudo-t. Cross-section: individual industry in an individual country. No cross-section dependence. Significance tests based on traditional t-stats. Sample: 1996-2004.

Industry	"Long run" exchange rate pass-through coefficients											
	Ireland			Greece			Portugal			Spain		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.42	0.9	5/02*	0.21	0.73	7/98***	0.58	1.91	4/02***	0.62	1.29	10/99***
	(0.06)	(0.27)	-***	(0.06)	(0.09)	-***	(0.05)	(0.28)	-***	(0.03)	(0.11)	-***
SITC_1	-0.33	-0.42	2/03	0.31	0.12	5/02	0.36	1.41	1/03	0.79	0.87	7/02
	(0.1)	(0.81)	-	(0.03)	(0.23)	+	(0.14)	(1.08)	-	(0.1)	(0.53)	-
SITC_2	0.57	0.66	7/02	0.39	0.27	7/00	0.78	0.58	1/99*	0.99	0.96	3/98
	(0.05)	(0.23)	-	(0.05)	(0.09)	+	(0.09)	(0.06)	+	(0.06)	(0.03)	+
SITC_3	0.88	0.93	5/03	1.12	-0.2	5/00***	0.96	1.03	11/98	0.97	1.06	3/00*
	(0.06)	(0.48)	-	(0.11)	(0.14)	+***	(0.11)	(0.04)	+	(0.03)	(0.03)	+
SITC_4	0.82	0.61	1/01	0.67	0.9	7/01	1.54	0.78	12/02	0.35	1.27	6/00***
	(0.14)	(0.17)	+	(0.12)	(0.25)	-	(0.21)	(0.47)	+	(0.13)	(0.15)	-***
SITC_5	0.27	0.47	12/97	0.53	0.41	8/99**	0.28	0.76	8/98**	0.41	0.81	7/00*
	(0.14)	(0.04)	-	(0.04)	(0.04)	+**	(0.19)	(0.09)	-***	(0.07)	(0.22)	-
SITC_6	0.57	0.9	8/02***	0.64	0.49	12/01	0.92	1.17	11/99***	0.9	1.24	2/01***
	(0.03)	(0.1)	-***	(0.03)	(0.12)	+	(0.06)	(0.05)	-**	(0.02)	(0.04)	-***
SITC_7	0.55	1.05	11/01	0.19	0.4	11/97	0.37	0.38	11/02	0.4	0.38	9/98
	(0.04)	(0.43)	-	(0.13)	(0.03)	-**	(0.02)	(0.31)	+	(0.08)	(0.02)	-
SITC_8	0.52	1.16	9/00**	0.74	0.13	2/02*	0.34	0.81	8/98***	1.05	1.08	12/99
	(0.09)	(0.23)	-***	(0.06)	(0.31)	+*	(0.13)	(0.06)	-***	(0.07)	(0.1)	+

Notes: see previous page.

Table 9: continued.

Industry	"Long run" exchange rate pass-through coefficients					
	Finland			Austria		
	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.86	0.73	10/00	0.69	0.79	2/99
	(0.04)	(0.13)	+*	(0.05)	(0.12)	–
SITC_1	0.71	1.01	3/02	0.64	0.5	3/99
	(0.04)	(0.2)	–	(0.09)	(0.19)	+
SITC_2	0.63	0.82	4/99	0.77	0.72	1/98
	(0.09)	(0.07)	–**	(0.06)	(0.03)	+*
SITC_3	0.72	0.9	4/99	0.75	0.88	12/97
	(0.1)	(0.04)	–*	(0.1)	(0.03)	–
SITC_4	-0.17	0.12	9/97	0.91	0.07	7/00***
	(0.34)	(0.11)	–	(0.17)	(0.19)	+***
SITC_5	0.47	0.81	4/03	0.21	0.46	2/99*
	(0.03)	(0.28)	–**	(0.1)	(0.07)	–**
SITC_6	0.72	0.74	12/00	0.43	0.51	9/00**
	(0.02)	(0.04)	–**	(0.02)	(0.03)	–
SITC_7	0.2	1.19	11/01***	0.27	0.23	5/02
	(0.03)	(0.32)	–***	(0.03)	(0.25)	+
SITC_8	0.59	0.49	4/00	0.58	0.56	12/98
	(0.06)	(0.11)	+	(0.07)	(0.03)	–

Notes: see previous page.

Table 9: continued.

interpretation of developments in individual countries and sectors in the following section we will observe some interesting patterns in estimated break points.

We consider the test results from the panel as sufficient evidence in favor of the existence of a "long run" levels relationship between the variables, as implied by the theoretical underpinning - equation (9). Moreover, despite some variability in the estimated breaks in the individual series we can say that at least for some country/industry combinations there is evidence that the formation of the EMU led to a significant change in the equilibrium pass-through rate, be it directly upon its formation or indirectly by tying the currency to the euro, and thus seeing it appreciate against the dollar since about 2001.

However, as given above, the failure of first generation panel cointegration tests to account for cross section dependence tends to oversize the tests and may lead to flawed inference on the existence of the long run relationship. Our final generalization of the testing framework, having already developed tests for cointegration with structural breaks, is to allow for a factor structure to model this type of dependence (as in Bai and Ng, 2004) and apply the test proposed by Banerjee and Carrion-i-Silvestre (2006) which allows for a single (common) estimated break in the series. In this second generation test, we test the null of no cointegration against an alternative hypothesis of cointegration (with up to r common factors modelling cross section dependence) with one common break date for all the series. The results for 1-, 3- and 6-factor dependence structures are reported in Table 10 and we consider them as reconfirmation of the existence of a long run equilibrium relationship.

The results in Table 10 are computed with the common break date imposed at 1999m1. More general results are plotted in Figure 8 where we report the Banerjee and Carrion-i-Silvestre (2006) test statistic for all possible choices of common break dates (subject to

H_0: unit root (no co-integration)

No. of factors	Pseudo-t		
	No break	Break (a)	Break (b)
(1)	-4.51	-3.72	-3.11
(3)	-3.71	-3.47	-2.60
(6)	-3.76	-2.95	-3.47

Under the null the statistics have the normal $N(0,1)$ distribution.
No break, (a) - break in constant within co-integrating equation.
(b)break in the entire co-integrating relationship.
Date of break imposed 1999m1.

Table 10: Test statistics for Banerjee and Carrion-i-Silvestre (2006) panel co-integration tests with cross-section dependence (common factors). Sample: 1996-2005.

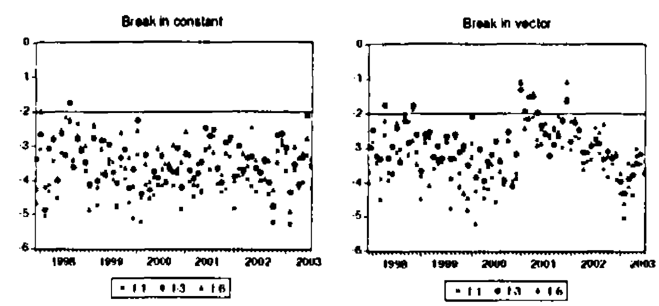

Figure 8: Distribution of the test statistics from Banerjee and Carrion-i-Silvestre (2006) test. The null hypothesis is no cointegration, the alternative cointegration with a common break in constant (left hand side) as in equation (12), and cointegration with common break in cointegrating vector (right hand side) as in equation (13). In both cases cross sectional dependence is modelled with a 1-, 3- and 6-factor structure. Under the null hypothesis, the statistics have a $N(0,1)$ distribution. Each observation is obtained by estimating the model with a given break date.

trimming). Clearly, regardless of the break date imposed, in almost all the cases the null of no cointegration is rejected. The assumption of a common break point may be a limitation, since as reported earlier, the estimated break points may be quite dispersed, and may limit our ability to choose the strongest break point - yet even for the dates for which there would seem to less intuitive reason to impose a break, we are often able to reject the null of no cointegration comfortably.

III.7 Discussion and concluding remarks

The results of our paper show ample evidence for an EG cointegrating relationship between the variables in levels - as in the underlying theoretical equation (9). We have suggested several methods for working with the data that enable the cointegrating relationship to be detected, including better lag-length selection in the tests for cointegration and a consideration of the impact of structural change and conducting inference using a panel (where the N dimension augments the T). By taking care of the adverse effects of cross-section dependence, we have shown that the evidence from the panel tests - with or without allowing for structural breaks - is entirely unambiguous. Thus, even if one were not willing to accept the notion of 'detectable' structural change, as modelled in this paper, or were only willing to attribute the finding of a break to data issues, it should be noted that our main contentions would still hold. We can therefore redefine the long run effect of exchange rate fluctuations on prices to be consistent with the theoretical literature. Instead of a rather arbitrary sum of (mainly insignificant, of opposite signs) coefficients on lags of the exchange rate,[11] which we discussed in Section III.3, we propose using the EG cointegrating equation coefficient. The use of the standard measure of pass-through should be viewed with caution, or re-interpreted substantially.

Our main preoccupation is the fact that despite using data of monthly frequency and aiding ourselves with panel methods, we still deal with a relatively short sample of at most 10 years. This may prove too short for the 'true' long run to reveal itself. This problem is aggravated by the specific developments in our sample - namely the introduction of the euro, the Maastricht criterion de facto restricting the bilateral movements of the exchange

[11] This would, as we have argued, make the estimates problematic to interpret in any case - regardless of whether the notion of a long run, defined as the sum of four or five short run effects, is coherent.

rates of the countries since 1997, and the depreciation and subsequent appreciation of the euro-dollar rate.

However, we think that the sheer fact that we find overwhelming evidence in favor of a cointegrating relationship in the EG sense provides backing to the presence of a theoretically implied relationship as in equation (9). Augmenting this finding with our techniques for (a) dealing with change in the sample; and (b) extracting information from a panel, where pooling or averaging over the 90 units counteracts the noise arising from the T dimension, makes our case more compelling.

When discussing the long run ERPT, anticipating our discussion of the results below, a number of other issues arise concerning the magnitude of the pass-through coefficient as measured by us. First of all, can we reasonably expect ERPT to be lower than one in the long run? Most of the literature, see for example Smets and Wouters (2002), is based on the notion that nominal rigidities cause imperfect ERPT. But as rigidities such as sticky prices tend to be rather a short to medium run phenomena, one may be led to think that producers would be unwilling to accommodate a change in the exchange rate into his mark-up forever, thus leading to expect full ERPT in the long run.

This story is not entirely convincing - the foreign exporter maximizes profit, not mark-up over a set of markets and over time, and thus may be willing to accept to adjust his mark-up in order to maintain market share, adapting to competitive conditions both in the short and long run. The fact that empirically, exchange rates are found to be much more volatile than prices, would also suggest that even in the long run, not all exchange rate fluctuations are passed on to the price level and some of the adjustment may be done through quantity. Consequently, Corsetti, Dedola and Leduc (2006) propose a model where ERPT is lower then one even in the long run, as a result of price discrimination and thus different pricing strategies between markets. Thus overall, a finding of the exchange rate pass-through to be lower than one in the long run is also not unreasonable.

Finally, in this context, can we expect the long run ERPT to be greater than 1? Essentially, in the long run, the answer is 'no'. We do find a handful of the series exhibiting an ERPT estimate significantly greater than one. However, most commonly this occurs when we allow for a break in the slope and the break is estimated as being located rather close to the end of the sample, making inference unreliable.

Before proceeding to a discussion of the estimated break dates, their location, and the direction and nature of the change in the cointegrating relationships, we should emphasize that both our single-equation and panel methods allow for only a single break. This is a limitation imposed by the relatively short span of our sample. If there are, say, two breaks in the data, the algorithm may pick up only one of them, or estimate a break lying somewhere in between the two actual breaks. This may account for some of the heterogeneity reported in the tables.

Generally speaking, both in our single equation framework, as well as in the panel estimates with increased power, we find evidence of a change in the vicinity of the introduction of the common currency (1998-9) or in the vicinity of the exchange rate developments turnaround (2001-2).

First, in the case of the breaks estimated to lie near 1998-9, thus coinciding with the introduction of the euro - there are reasons to expect both 'monetary' and 'real' effects of the common currency. As for the former, a vast literature tends to suggest that we should expect ERPT to fall upon the introduction of the euro (see for instance Devereux, Engel and Tille, 2003, who argue that as the new currency becomes the currency of invoicing, European prices will become more insulated from exchange rate volatility). However, in our estimates we tend to find, especially for Italy and also Portugal and Spain, where the breaks coincide with the euro introduction, there tends to be a significant rise in ERPT - which suggests the story is not as simple. First of all, the above argument concerns primarily short run pass-through, while in this paper we focus on ERPT in the long run. In principle, there is no reason why it would not be possible to observe even opposing movements in the short- and long-run ERPT. Moreover, the acceptance of the euro as an invoice currency may take far longer than we are able to pick up in our short sample. Second, the euro can be expected to have reduced the 'noise' in the exchange rate movements, especially for countries such as Italy, Spain and Portugal. In a noisy, volatile environment producer currency pricing may prove difficult - frequent and often temporary exchange rate changes, confronted with menu costs or costly pricing strategy reviews may lead import goods to be actually more local currency priced. Arguably, especially in the mentioned countries, as the euro was introduced, the amount of noise in exchange rate developments may have fallen, thus actual changes in the exchange rate may have become no longer perceived as noisy, temporary shocks but more of a somewhat permanent and macro-founded nature, which the foreign exporter may become more willing to pass them on to the price. This is in line with the models of Adolfson (2001) and Corsetti, Dedola and Leduc (2006) which generate high volatility of exchange rate associated with low ERPT.

As for the 'real' effects of the common currency following the introduction of the euro, roughly 50% of the imports became by default home currency priced, and thus no longer subject to fluctuations in the exchange rate. This potentially meant a change in competitive conditions for extra-euro imports, for various reasons related to increased price transparency. The latter effect would tend to work in the direction of decreasing ERPT with the formation of the euro, however its strength relies largely on the extent to which extra- and intra-euro imports within a singe 1-digit SITC category actually compete with each other.

We do however show there is some evidence of this effect, namely the estimated reduction of the constant markup (negative estimates α_1 in the most general specification of equation (13)) towards the second part of the sample suggests increased competition between importers.

Second, as for the breaks in the vicinity of 2001, that is coinciding with the period when the euro (and thus the 'local currencies' in our sample), after several years of depreciation against the dollar, started off on a relatively stable appreciation. The reasoning for a possible asymmetrical effect of these exchange rate developments on import prices was briefly provided in the previous sections and is generally based on the notion that as the euro was depreciating, imported goods (which according to our assumptions, and following CM, have a world price in dollars) if priced in dollars in the intra euro market, would be becoming more expensive if the exchange rate change were passed through into the price. Thus in order to stay competitive and maintain market share, the foreign producers could have been expected to accommodate some part of the rise - thus ERPT could be expected to be lower than if a producer currency pricing strategy were adopted. The turn-around in the exchange rate developments meant goods with dollar prices becoming cheaper on the intra euro market, which may have inclined producers to be more willing to shift away from local currency pricing. By passing through more of their dollar price, they would be maintaining their revenue in terms of the dollar, but finding it easier to gain an edge in the market and compete with local products. Notice that as we look at import prices this does not necessarily imply a change (fall) in the price level, nor a gain in market share, as there are many other factors at work (such as changing retailer margins). We treat the fact that in the cases when our estimated break point lies near 2001 the estimated ERPT rises, as strong support for the above story of an asymmetrical ERPT.

Next, there are other developments, arguably harder to date, that may have had an effect on the long run ERPT - among them are: the increase in trade integration, ongoing trade liberalization, specific import compositions of individual countries, such as the Irish large share of pound rather than dollar priced goods and the different evolution of euro/pound and euro/dollar rates especially since 1999.[12]

Last, it is important to mention the fact that the effects of incidents like euro adoption cannot be expected to happen on an exact date - on one hand, they do not come fully unexpected, and thus may be anticipated to some degree, and on the other, the effect may be gradual and thus picked up with a lag.

Having discussed the breaks, we can now turn to analyzing the actual long run pass-through estimates in more detail.

We will focus on the coefficients in Table 9 as the most general setting, which allows for breaks in both the constant and the cointegrating vector. The increase in ERPT in most sectors in countries like Italy, Spain and Portugal usually coincides with the introduction of the euro. This may be a sign of the increase in the credibility of the monetary regime, that occurred when these countries joined the euro area, which led foreign producers to expect more stable conditions, and as argued previously made them more willing to pass on the actual fluctuations. This change is not evident in the case of Greece, which generally has rather low pass-through rates, but joined the euro 2 years later.

As for the sectors, notably sector SITC 5 (Chemical products) faced an increase (signif-

[12]Another suggested issue is the integration of China into the global economy. Although, we expect the effect to work primarily through the 'world price' rather then the degree of pass-through, there may be some room for the latter because of the change in the competition conditions induced by the inflow of Chinese goods. Sector trade shares of imports from China relative to all imports and relative to extra-EU imports grew steadily in the manufacturing sectors throughout our sample and notably in sectors SITC_7 and SITC_8 seemed to accelerate around 2001. This pattern prevailed for most EU countries.

icant in 7 of the 10 countries) in pass-through across almost all the countries in question, rendering it closer to 1. Next, in SITC 0 (Food and live animals) there has been an increase in the pass-through in the Netherlands, Italy, Ireland, Greece, Portugal and Spain from values around 0.5 to values much closer to 1. The estimated dates of this change are close to euro introduction for Italy, Greece and Spain. For the Netherlands, Ireland and Portugal the estimated breaks lie in mid 2002 and even in 2003. In the case of Ireland, an explanation for this may be provided by her intensive trade ties with the United Kingdom, which is by far the most important origin of imports into that country. As opposed to the euro/dollar exchange rate, the euro/sterling rate was relatively stable throughout our sample (see Figure 9). Specifically the British pound did not depreciate against the euro as the dollar did since about 2001. Thus euro movements versus the dollar may have had much lesser influence on the pass-through in this country, and suggests the weakness of the integrated world market assumption for Ireland. Finally in all the specifications it is evident the pass-through in sector SITC 3, i.e. mineral fuels, is practically equal to or very close to 1, and has not changed substantially upon the introduction of the euro. This may be explained by strong foreign market power of producers in this sector, who for instance face practically close to zero domestic competition in products like oil, and thus a common world price is fully passed on when the exchange rate fluctuates.

Generally we are able to reject zero pass-through rates much more often than using the arbitrary long run definition of a sum of 5 lags. Our estimated pass-through coefficients tend to be closer to 1 in magnitude, although they are often significantly different (from 1) due to much narrower confidence intervals. Moreover, we are able to provide an explanation for the increase in the pass-through rate that seemed to occur after 2000 in many countries. As mentioned previously, in the first part of the sample, the euro, and thus most currencies

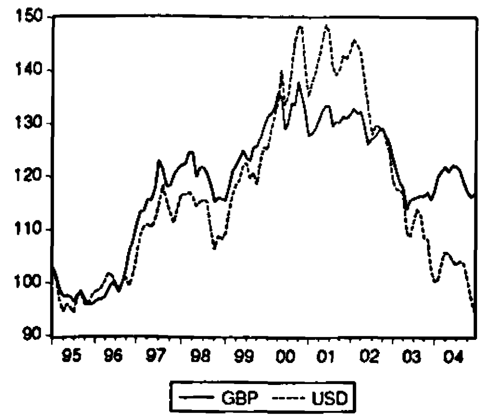

Figure 9: Evolution of EUR(ECU)/USD and EUR(ECU)/GBP exchange rates(1996=100).

related to it (e.g. through the ERM), was depreciating. Foreign producers may have been forced to absorb some of this increase in the relative price of their good in order to maintain the market share. Since 2001 the euro started to appreciate. As foreign producers where able to receive their relatively unchanged income in the foreign currency, they may have been able to pass on a larger part of the change in the exchange rate.

To summarize, in this paper, we propose a new estimate of the long run ERPT. The incorporation of the levels equilibrium relationship that we propose renders the empirical estimation of ERPT more consistent with the theoretical underpinnings which are in fact a levels relationship.

The empirical literature has been somewhat forced to look for alternative, more arbitrary definitions of long run ERPT, because of a failure to find a cointegrating relationship between the variables. We show that proper choice of lag lengths in unit root tests, allowing for breaks in the series and using panel methods facilitates the discovery of such an equilibrium relationship in the data, and thus improved estimation of both long and short run ERPT.

Overall, ERPT in the long run is found to be equal to one or close to one in the commodity sectors, throughout the entire sample, while it tends to be rather lower than one in the manufacturing, food, beverages and tobacco and chemical sectors. As there are a number of reasons, such as the euro introduction and exchange rate developments that lead us to suspect a potential change in the long run relationship, we use up-to-date panel methods, to estimate possible break dates and changes in ERPT and account for possible cross-section dependence.

We tend to favor the most flexible specification, i.e. the one allowing for a break in the entire cointegrating relationship, as in this case the estimated shift in the constant is in line with the expected increase in competition expected with trade integration and the introduction of the euro.

Allowing for a structural break in the relationship we find that ERPT has generally increased in the vicinity of the euro introduction and the change is especially evident in Southern European countries. This may be the effect of perceived stabilization in the monetary regime, which led to less noise in exchange rate developments. Moreover the increase in ERPT in the second part of our sample may be due to specific exchange rate developments (euro/dollar depreciation till 2000, and subsequent appreciation) which may suggest asymmetrical response of the import prices. When we allow for the change in the long run relationship, we find that, towards the second part of our sample, that is after the estimated break date, apart from Greece and perhaps a number of manufacturing sectors in Austria, long run ERPT was not generally substantially (in most cases not significantly) lower than 1.

Obviously in order to be able to speak more confidently of the EG long run ERPT, we would require a longer series, ranging both further back and beyond the date of the introduction of the euro. While this is the subject of on-going research, we hope we have been able in this paper to question the basis of the empirical literature surrounding estimation of ERPT and to propose a set of alternative ideas for discussion.

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Appendix A - Data

Sources: Eurostat, COMEXT.
import prices - monthly indexes of import unit values (calculated to be based on local currency) for imports originating outside the euro area.
foreign prices - monthly indexes of import unit values (calculated to be based on US dollars) from imports originating outside the euro area into the euro zone.
exchange rates - index of monthly average exchange rate of local currency against the US dollar.
All variables are in logs.
SITC code - Industry
0 - Food and live animals chiefly for food
1 - Beverages and Tobacco
2 - Crude materials, inedible, except fuels
3 - Mineral fuels, lubricants and related materials
4 - Animal and vegetable oils, fats and waxes
5 - Chemicals and related products, n.e.s.
6 - Manufactured goods classified chiefly by materials
7 - Machines, transport equipment
8 - Manufactured goods n.e.c.

- CM data set 1989-2001 - series for 1989m1-2001m3: Belgium+Luxembourg, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal and Spain. Series for 1996m1-2001m3: Austria and Finland.

- "new" data set 1995-2005 - 1995m1-2005m3 for 10 out of 11 countries of the CM data set (Belgium+Luxembourg excluded, Austria and Finland start 1996m1, Portugal and Austria stop 2004m1)

- full panel - reduced version of 1995-2005 data set, trimmed in order to obtain a balanced panel. Covers 1996m1-2004m12 for all 10 countries.

- full panel for CM 1989-2001 sample - 9 countries: Austria and Finland excluded, due to short series. Series Ireland SITC 4 and Portugal SITC 4 also excluded due to missing values.

Appendix B - Tables for the CM 1989-2001 sample

H_0: Unit root (no cointegration)

Country	SITC0	SITC1	SITC2	SITC3	SITC4	SITC5	SITC6	SITC7	SITC8
France	-3.58***	-1.95**	-2.15**	-4.51***	-3.46***	-4.77***	-2.38**	-4.56***	-5.36***
Bel+Lux.	-4.35***	-3.82***	-4.26***	-3.97***	-3.83***	-3.52***	-2.05**	-2.53**	-3.96***
Netherlands	-4***	-3.96***	-2.24**	-5.12***	-2.63***	-4.09***	-2.63***	-3.23***	-3.72***
Germany	-3.04***	-2.86***	-2.87***	-2.86***	-3.52***	-3.67***	-1.86*	-4.28***	-4.29***
Italy	-2.81***	-2.15**	-5.77***	-7.44***	-3.86***	-3.68***	-2.39**	-4.32***	-2.19**
Ireland	-3.22***	-2.51**	-7.1***	-2.72***	-3.76***	-2.48**	-3.42***	-2.41**	-3.38***
Greece	-4.24***	-4.33***	-3.15***	-4.01***	-3.08***	-2.98***	-3.16***	-1.96**	-2.65***
Portugal	-5.28***	-1.45	-3.06***	-4.83***	-3.28***	-4.57***	-2.65***	-5.6***	-7.8***
Spain	-3.38***	-3.1***	-2.43**	-5.8***	-4.47***	-2.02**	-3.64***	-2.59**	-2.27**
Finland	-1.65*	-3.83***	-2.17**	-2.87***	-4.41***	-1.1	-1.99**	-4.36***	-4.86***
Austria	-5.02***	-3.72***	-2.78***	-3.63***	-3.15***	-5.27***	-4.13***	-2.92***	-2.78***

For each sector first line reports the ADF t-statistic, and the second p-value.
Specification: no constant, no trend. Maximum lags number = 12. Selection: Akaike (AIC)
****,**,* denote the null hypothesis of no cointegration can be rejected at 1%, 5% and 10%.*

Table 11: ADF tests on the errors from the OLS regression of the "long run" equation (9). Equivalent for Table 3 on sample 1989-2001.

Industry	France (1)	France (2)	Belg.+Lux. (1)	Belg.+Lux. (2)	Netherlands (1)	Netherlands (2)	Germany (1)	Germany (2)	Italy (1)	Italy (2)	Ireland (1)	Ireland (2)
	"Long run" exchange rate pass-through coefficients											
SITC_0	0.87	*98/6*	(.)		1.03	93/4	(.)		(.)		(.)	
	(0.02)	-***	(.)		(0.02)	-***	(.)		(.)		(.)	
SITC_1	0.79	95/5	(.)		(.)		(.)		0.28	95/9	0.48	98/10
	(0.05)	+***	(.)		(.)		(.)		(0.08)	+***	(0.17)	+***
SITC_2	(.)		0.91	94/10	0.72	97/7	(.)		1.11	93/8	0.72	92/7
	(.)		(0.03)	-***	(0.07)	+***	(.)		(0.02)	-***	(0.05)	-***
SITC_3	(.)		(.)		(.)		(.)		1.06	94/1	(.)	
	(.)		(.)		(.)		(.)		(0.02)	+***	(.)	
SITC_4	(.)		1.2	98/11	(.)		(.)		(.)		n.a.	
	(.)		(0.05)	-***	(.)		(.)		(.)		n.a.	
SITC_5	0.94	91/1	1.17	95/9	(.)		1.07	98/3	0.87	97/11	1.29	96/7
	(0.03)	+***	(0.05)	+***	(.)		(0.03)	+***	(0.03)	-***	(0.11)	+***
SITC_6	(.)		1.1	96/9	1.36	92/12	0.76	98/3	1.11	92/6	0.55	95/1
	(.)		(0.02)	-***	(0.02)	+***	(0.01)	+***	(0.02)	-***	(0.05)	+***
SITC_7	1.14	91/8	1.17	97/6	1.04	96/3	0.99	97/3	0.88	90/12	1.8	96/2
	(0.02)	-***	(0.03)	-***	(0.04)	-***	(0.02)	-**	(0.01)	+***	(0.12)	+***
SITC_8	1.07	93/1	(.)		(.)		1.05	95/10	0.75	*95/3*	(.)	
	(0.02)	+***	(.)		(.)		(0.02)	-***	(0.03)	+***	(.)	

Value not reported (.) if the hypothesis of unit root (no cointegration) cannot be rejected at 10%
(Italics) if the hypothesis of unit root (no cointegration) can be rejected at 10%, but cannot be rejected at 5%;
For each country/industry combination columns: (1) first row estimated coefficient on ER, second row standard error, (2) first row estimated break date, second row direction of change in constant
+ or − indicate change in constant is positive or negative,
,,*** indicate this change is significant (t-stat of $\widehat{\alpha}_1$) at 10%, 5% and 1%*

Table 12: Estimated ERPT coefficients and break dates extracted from GH. Specification: break in constant - equation (12). Statistic: ADF*. Sample: 1989-2001 equivalent of Table 5.

Industry	Greece		Portugal		Spain		Finland		Austria	
	"Long run" exchange rate pass-through coefficients									
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
SITC_0	0.36	92/10	1.04	91/2	1.02	99/5	(.)		(.)	
	(0.04)	-***	(0.05)	+***	(0.03)	+***	(.)		(.)	
SITC_1	0.91	*93/10*	(.)		1.65	94/10	(.)		0.11	98/9
	(0.09)	-***	(.)		(0.06)	-***	(.)		(0.21)	-***
SITC_2	0.47	94/10	(.)		(.)		1.28	97/10	0.66	00/3
	(0.06)	-***	(.)		(.)		(0.07)	-***	(0.04)	-***
SITC_3	0.88	92/1	0.65	*91/1*	1.13	99/2	1.01	00/2	0.78	00/2
	(0.08)	-***	(0.03)	-***	(0.02)	-***	(0.08)	-***	(0.06)	-***
SITC_4	(.)		n.a.		1.28	95/2	0.54	99/3	(.)	
	(.)		n.a.		(0.03)	+***	(0.36)	+***	(.)	
SITC_5	0.28	95/6	0.98	*93/9*	0.94	96/9	1.63	97/6	0.71	97/6
	(0.05)	-***	(0.09)	-***	(0.03)	-***	(0.17)	-***	(0.1)	-**
SITC_6	(.)		(.)		1.02	94/9	0.99	99/7	(.)	
	(.)		(.)		(0.03)	-***	(0.06)	+***	(.)	
SITC_7	(.)		0.77	97/12	(.)		0.6	99/11	(.)	
	(.)		(0.03)	-***	(.)		(0.08)	+***	(.)	
SITC_8	(.)		1.01	98/2	(.)		0.67	98/4	(.)	
	(.)		(0.06)	-***	(.)		(0.08)	-***	(.)	

Notes: See previous page for explanations.

Table 12: continued.

Industry	France (1)	France (2)	France (3)	Belg.+Lux. (1)	Belg.+Lux. (2)	Belg.+Lux. (3)	Netherlands (1)	Netherlands (2)	Netherlands (3)	Germany (1)	Germany (2)	Germany (3)
SITC_0	(.)	(.)		(.)	(.)		1.12	0.53	*97/8****	(.)	(.)	
							(0.01)	(0.09)	+***			
SITC_1	0.39	0.99	95/1***	(.)	(.)		0.32	1.6	97/8***	0.2	0.71	*91/11****
	(0.09)	(0.07)	-***				(0.06)	(0.18)	-***	(0.12)	(0.03)	-*
SITC_2	(.)	(.)		1.15	0.76	97/5***	0.95	0.36	97/11***	1.02	1.14	96/12***
				(0.05)	(0.06)	+***	(0.07)	(0.09)	-	(0.03)	(0.03)	-
SITC_3	0.83	1.02	92/9***	(.)	(.)		(.)	(.)		(.)	(.)	
	(0.05)	(0.02)	-***									
SITC_4	(.)	(.)		1.2	1.29	98/11	(.)	(.)		(.)	(.)	
				(0.05)	(0.25)	-						
SITC_5	0.85	0.9	93/6	0.97	1.29	95/9**	(.)	(.)		0.71	1.15	92/2***
	(0.15)	(0.03)	+	(0.12)	(0.08)	-**				(0.15)	(0.02)	-***
SITC_6	0.86	1	93/11***	1.15	0.97	94/6***	(.)	(.)		0.73	0.81	98/3***
	(0.02)	(0.02)	-***	(0.03)	(0.03)	+**				(0.02)	(0.02)	-
SITC_7	0.91	1.11	92/9**	(.)	(.)		0.84	1.15	96/3***	0.86	0.95	*91/1*
	(0.08)	(0.02)	-***				(0.08)	(0.07)	-**	(0.23)	(0.01)	-
SITC_8	0.87	1.06	93/8***	1.16	1.11	92/3	(.)	(.)		(.)	(.)	
	(0.06)	(0.02)	-	(0.11)	(0.03)	-						

"Long run" exchange rate pass-through coefficients

Value not reported (.) indicates H_0: unit root (no cointegration) cannot be rejected at 10%
(Italics) indicate H_0 cannot be rejected at 5%, but can be rejected at 10%;
*For each country and industry combination columns: (1) first row coefficient on ER before break date, second row standard error, (2) first row coefficient on ER after break date, second row standard error, (3)first row: estimated break date, second row direction of change in constant + positive, - negative, *,**,*** indicate the change of the long run pass-through (first row) or constant (second row) on the date of the break is significant (t-stat of $\widehat{\beta}_1$ or $\widehat{\alpha}_1$ respectively) at 10%, 5% and 1%.*

Table 13: Estimated coefficients and break dates extracted from GH. Specification: break in cointegrating vector. Statistic: ADF*. Sample: 1989-2001 equivalent of Table 6

Industry	"Long run" exchange rate pass-through coefficients											
	Italy			Ireland			Greece			Portugal		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	(.)	(.)		(.)	(.)		0.27	0.95	94/11***	0.52	1.08	90/12
							(0.04)	(0.07)	-***	(0.38)	(0.05)	-
SITC_1	0.74	-0.08	96/2***	-0.47	1.66	*97/9****	(.)	(.)		(.)	(.)	
	(0.11)	(0.21)	+	(0.21)	(0.68)	-						
SITC_2	1.02	1.31	97/4***	0.4	0.79	92/7***	(.)	(.)		(.)	(.)	
	(0.02)	(0.03)	-***	(0.12)	(0.06)	-*						
SITC_3	1.06	1.02	94/6	(.)	(.)		0.98	0.84	92/3	0.84	0.75	99/3
	(0.03)	(0.03)	-				(0.27)	(0.07)	-***	(0.05)	(0.13)	+***
SITC_4	0.54	0.67	94/2	n.a.	n.a.		(.)	(.)		n.a.	n.a.	
	(0.06)	(0.09)	-***									
SITC_5	0.86	0.92	97/11	1.37	1.17	96/7	0.23	0.76	95/3***	(.)	(.)	
	(0.03)	(0.09)	-	(0.18)	(0.2)	+	(0.07)	(0.14)	-***			
SITC_6	(.)	(.)		0.32	0.8	93/11***	-0.11	1.13	*96/11****	(.)	(.)	
				(0.07)	(0.05)	-***	(0.03)	(0.09)	-***			
SITC_7	1.54	0.89	91/12***	(.)	(.)		(.)	(.)		0.87	0.59	95/2***
	(0.14)	(0.02)	+***							(0.05)	(0.04)	+
SITC_8	0.92	0.55	96/1***	(.)	(.)		(.)	(.)		1.04	0.96	98/2
	(0.02)	(0.08)	+							(0.07)	(0.11)	+

Notes: See previous page.

Table 13: continued.

Industry	"Long run" exchange rate pass-through coefficients								
	Spain			Finland			Austria		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
SITC_0	0.96	1.71	97/9***	(.)	(.)		(.)	(.)	
	(0.03)	(0.13)	-***						
SITC_1	1.83	1.47	94/11***	(.)	(.)		(.)	(.)	
	(0.1)	(0.09)	+**						
SITC_2	(.)	(.)		0.97	1.36	*97/11***	0.68	0.44	00/3
				(0.15)	(0.07)	–	(0.04)	(0.15)	+
SITC_3	1.12	1.26	99/2	0.91	-0.24	*99/6****	0.91	0.55	00/1*
	(0.02)	(0.13)	–	(0.12)	(0.23)	+***	(0.05)	(0.2)	+***
SITC_4	1.2	1.05	98/6	0.43	0.66	99/3	(.)	(.)	
	(0.03)	(0.24)	+*	(0.5)	(0.87)	–			
SITC_5	0.9	1.17	97/3***	0.25	1.89	97/11***	0.23	0.8	98/6***
	(0.04)	(0.09)	-*	(0.32)	(0.14)	-**	(0.18)	(0.1)	-**
SITC_6	(.)	(.)		1.03	0.72	99/11**	0.42	1.5	00/4***
				(0.09)	(0.12)	+*	(0.05)	(0.19)	-***
SITC_7	(.)	(.)		0.76	0.1	00/2**	(.)	(.)	
				(0.08)	(0.32)	+***			
SITC_8	(.)	(.)		1.02	0.52	99/8**	(.)	(.)	
				(0.15)	(0.13)	+*			

Notes: see previous page

Table 13: continued.

Model	pseudo-t	pseudo-ρ
No break	-8.27	-35.53
Break in constant	-15.61	-40.18
Break vector	-19.51	-42.22
Under the null hypothesis both statistics have a N(0,1) distribution		

Table 14: Pseudo-t and Pseudo-ρ parametric statistics from Banerjee and Carrion-i-Silvestre (2006) version of Pedroni(1999). The null hypothesis is no cointegration. Sample: 1989-2001, full panel (N=79) excl. Austria and Finland, unit specific breaks. Equivalent of Table 7.

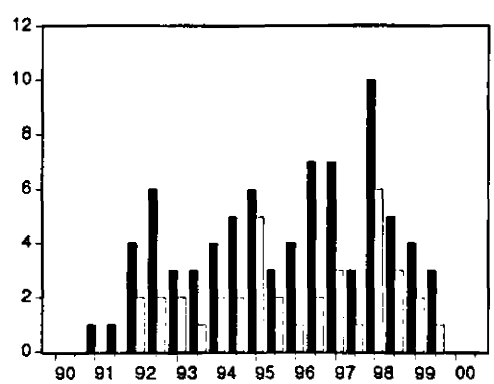

Figure 10: Semi-annual distribution of estimated break dates (1990s1-2000s2). Breaks in cointegrating vector. Dark color - all breaks, light color - only breaks when long run ERPT changed significantly (10%). Sample: 1989-2001, excl. Finland and Austria. Equivalent of Figure 7.

H_0: unit root (no co-integration)

No. of factors	Pseudo-t						
	No break	Break (a)			Break (b)		
		ave.	min.	max	ave.	min.	max
(1)	-2.64	-2.51	-3.59	-1.55	-2.32	-3.77	-0.29
(2)	-2.89	-2.49	-3.95	-1.54	-2.07	-3.62	-0.75
(3)	-1.88	-1.76	-2.79	-0.66	-1.86	-3.67	-0.52

Under the null the statistics have the normal $N(0,1)$ distribution.
No break, (a) - break in constant within co-integrating equation,
(b)break in the entire co-integrating relationship.

Table 15: Test statistics for Banerjee and Carrion-i-Silvestre (2006) panel co-integration tests with cross-section dependence (common factors). Sample: 1989-2001, Finland and Austria excluded. Columns (ave), (min) and (max) report average, minimum and maximum values of the statistic when estimation is repeated for each common break between 1990m4 and 1998m10. Extended version of Table 10.

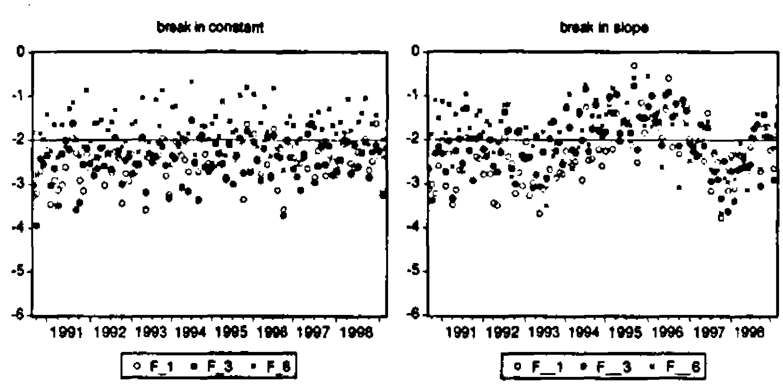

Figure 11: Distribution of the test statistics from Banerjee and Carrion-i-Silvestre (2006) test. The null hypothesis is no cointegration, the alternative cointegration with a common break in constant (left hand side) as in equation (12), and cointegration with common break in cointegrating vector (right hand side) as in equation (13). In both cases cross sectional dependence is modelled with a 1-, 3- and 6-factor structure. Under the null hypothesis, the statistics have a $N(0,1)$ distribution. Each observation is obtained by estimating the model with a given break date. Sample: 1989-2001, Finland and Austria excluded.

Appendix C - Descriptions of tests

Single equations with breaks - Gregory and Hansen (1996), panel cointegration without cross-sectional dependence: Pedroni (1999) - without breaks and with breaks as in Banerjee and Carrion-i-Silvestre (2006)[13]

For the purpose of describing the formal setup of the tests, let $\{Y_{i,t}\}$ be a $(m \times 1)$-vector of non-stationary stochastic process with the following representation

$$\Delta x_{i,t} = v_{i,t} \tag{14}$$

$$y_{i,t} = f_i(t) + x'_{i,t}\delta_{i,t} + e_{i,t}; \quad e_{i,t} = \rho_i e_{i,t} + \varepsilon_{i,t}, \tag{15}$$

where $Y_{i,t} = \left(y_{i,t}, x'_{i,t}\right)'$ is conveniently partitioned into a scalar $y_{i,t}$ and the $((m-1) \times 1)$-vector $x_{i,t}$, $i = 1, \ldots, N$, $t = 1, \ldots, T$. Let $\xi_{i,t} = \left(\varepsilon_{i,t}, v'_{i,t}\right)'$ be a random sequence assumed to be strictly stationary and ergodic, with mean zero and finite variance. In addition, the partial sum process constructed from $\{\xi_{i,t}\}$ satisfy the multivariate invariance principle defined in Phillips and Durlauf (1986). At this stage and in order to set the analysis in a simplified framework, let us assume that $\{v_{i,t}\}$ and $\{\varepsilon_{i,t}\}$ are independent.

The general functional form for the deterministic term $f(t)$ is given by:

$$f_i(t) = \mu_i + \beta_i t + \theta_i DU_{i,t} + \gamma_i DT^*_{i,t}, \tag{16}$$

where

$$DU_{i,t} = \begin{cases} 0 & t \leq T_{bi} \\ 1 & t > T_{bi} \end{cases}; DT^*_{i,t} = \begin{cases} 0 & t \leq T_{bi} \\ (t - T_{bi}) & t > T_{bi} \end{cases}, \tag{17}$$

with $T_{bi} = \lambda_i T$, $\lambda_i \in \Lambda$, denoting the time of the break for the i-th unit, $i = 1, \ldots, N$. Note also that the cointegrating vector is specified as a function of time so that

$$\delta_{i,t} = \begin{cases} \delta_{i,1} & t \leq T_{bi} \\ \delta_{i,2} & t > T_{bi} \end{cases}. \tag{18}$$

Using these elements, Banerjee and Carrion-i-Silvestre (2006) propose up to six different model specifications, of which for the purpose of this paper we will review two:

- Model 1. Constant term with a change in level but stable cointegrating vector:

$$y_{i,t} = \mu_i + \theta_i DU_{i,t} + x'_{i,t}\delta_i + e_{i,t} \tag{19}$$

- Model 4. Constant term with change in both level and cointegrating vector:

$$y_{i,t} = \mu_i + \theta_i DU_{i,t} + x'_{i,t}\delta_{i,t} + e_{i,t} \tag{20}$$

Using any one of these specifications the authors propose testing the null hypothesis of no cointegration against the alternative hypothesis of cointegration (with break) using the ADF test statistic applied to the residuals of the cointegration regression as in Engle and Granger (1987) and Gregory and Hansen (1996) but in the panel data framework

[13]This Appendix is a extract from Banerjee and Carrion-i-Silvestre (2006). For more details, including setup, derivations, asymptotic properties and finite sample simulations we refer the reader to the original papers.

developed in Pedroni (1999, 2004). In fact, Gregory and Hansen (1996) propose both of the specifications given by models 1 and 4 above.

The Banerjee and Carrion-i-Silvestre (2006) proposal starts by following Gregory and Hansen (1996), to the OLS estimation of one of the models given above (in our case (19) and (20)) and run the following ADF type-regression equation on the estimated residuals $(\hat{e}_{i,t}(\lambda_i))$:

$$\Delta\hat{e}_{i,t}(\lambda_i) = \rho_i\hat{e}_{i,t-1}(\lambda_i) + \sum_{j=1}^{k}\phi_{i,j}\Delta\hat{e}_{i,t-j}(\lambda_i) + \varepsilon_{i,t}. \tag{21}$$

The notation used refers to the break fraction (λ_i) parameter, which (if it exists) is in most cases unknown. In order to get rid of the dependence of the statistics on the break fraction parameter, Gregory and Hansen (1996) suggest estimating the models given above for all possible break dates, subject to trimming, obtaining the estimated OLS residuals and computing the corresponding ADF statistic.
With the sequence of ADF statistics in hand, one can also estimate the break point for each unit as the date that minimizes the sequence of individual ADF test statistics – either the t-ratio, $t_{\hat{\rho}_i}(\lambda_i)$, or the normalized bias, computed as $T\hat{\rho}_i(\lambda_i) = T\hat{\rho}_i\left(1 - \hat{\phi}_{i,1} - \cdots - \hat{\phi}_{i,k}\right)^{-1}$ – see Hamilton (1994), pp. 523. Note that the estimation of the break point $\hat{T}_{bi}$ is conducted as

$$\hat{T}_{bi} = \frac{\arg\min}{t}{}_{\hat{\rho}\lambda_i\in\Lambda_i}(\lambda_i); \quad \hat{T}_{bi} = \frac{\arg\min}{\lambda_i \in \Lambda} T\hat{\rho}_i(\lambda_i), \tag{22}$$

$\forall i = 1, \ldots, N$. At this point Gregory and Hansen (1996) test the null hypothesis for each unit. Gregory and Hansen (1996) derive the limiting distribution of $t_{\hat{\rho}_i}\left(\hat{\lambda}_i\right) = \inf_{\lambda_i\in\Lambda} t_{\rho_i}(\lambda_i)$ and $T\hat{\rho}_i\left(\hat{\lambda}_i\right) = \inf_{\lambda_i\in\Lambda} T\hat{\rho}_i(\lambda_i)$, which are shown not to depend on the break fraction parameter. Specifically, Gregory and Hansen (1996) show that $T\hat{\rho}_i\left(\hat{\lambda}_i\right) \Rightarrow \inf_{\lambda_i\in\Lambda} \int_0^1 Q(\lambda_i, s)\, dQ(\lambda_i, s) \Big/ \int_0^1 Q(\lambda_i, s)^2 ds$, and $t_{\hat{\rho}_i}\left(\hat{\lambda}_i\right) \Rightarrow \inf_{\lambda_i\in\Lambda} \int_0^1 Q(\lambda_i, s)\, dQ(\lambda_i, s) \Big/ \left[\int_0^1 Q(\lambda_i, s)^2 dr \left(1 + \varrho(\lambda_i)' D(\lambda_i)\, \varrho(\lambda_i)\right)\right]^{1/2}$, where $\Rightarrow$ denotes weak convergence, $Q(\lambda_i, s)$ and $\varrho(\lambda_i)$ are functions of Brownian motions and the deterministic component, and $D(\lambda_i)$ depends on the model – see the Theorem in Gregory and Hansen (1996) for further details.

Banerjee and Carrion-i-Silvestre (2006) propose combining the unit-specific information in a panel data statistic.

The panel statistics on which they focus in order to test the null hypothesis are given by the $Z_{\hat{\rho}_{NT}}$ and $Z_{\hat{t}_{NT}}$ tests in Pedroni (1999, 2004), which can be thought as analogous to the residual-based tests in Engle and Granger (1987). These test statistics are defined by pooling the individual ADF tests, so that they belong to the class of between-dimension test statistics. Specifically, they are computed as:

$$N^{-1/2} Z_{\hat{\rho}_{NT}}\left(\hat{\lambda}\right) = N^{-1/2}\sum_{i=1}^{N} T\hat{\rho}_i\left(\hat{\lambda}_i\right) \tag{23}$$

$$N^{-1/2} Z_{\hat{t}_{NT}}\left(\hat{\lambda}\right) = N^{-1/2}\sum_{i=1}^{N} t_{\hat{\rho}_i}\left(\hat{\lambda}_i\right). \tag{24}$$

where $\hat{\rho}_i\left(\hat{\lambda}_i\right)$ and $t_{\hat{\rho}_i}\left(\hat{\lambda}_i\right)$ are the estimated coefficient and associated t-ratio from (21) and

$$\hat{\lambda} = \left(\hat{\lambda}_1, \hat{\lambda}_2, \ldots, \hat{\lambda}_i, \ldots, \hat{\lambda}_N\right)' \tag{25}$$

is the vector of estimated break fractions.

Note that this framework allows for a high degree of heterogeneity since the cointegrating vector, the short run dynamics and the break point estimate might differ among units. The use of the panel data cointegration test aims to increase the power of the statistical inference when testing the null hypothesis of no cointegration, but some heterogeneity is preserved when conducting the estimation of the parameters individually.

Following Pedroni (1999), the panel test statistics are shown to converge to standard Normal distributions once they have been properly standardized.

Panel cointegration with cross-sectional dependence(Bai and Ng 2004[13]): Banerjee and Carrion-i-Silvestre (2006)[13]

The setup above extended static-regression based tests for cointegration to allow for structural breaks in the components of the regression. The underlying assumption was that panel units are cross-sectionally independent, which is quite rarely the case in economic applications. The extended approach in Banerjee and Carrion-i-Silvestre (2006) models cross-sectional dependence using common factors such as in Bai and Ng(2004). The tests can accommodate a (common) structural break.

Break point known

The underlying model is given in the following structural form:

$$\begin{aligned} y_{i,t} &= f_i(t) + x'_{i,t}\delta_{i,t} + u_{i,t} && (26)\\ u_{i,t} &= F'_t\pi_i + e_{i,t} && (27)\\ (I-L)F_t &= C(L)w_t && (28)\\ (1-\rho_i L)e_{i,t} &= H_i(L)\varepsilon_{i,t} && (29)\\ (I-L)x_{i,t} &= G_i(L)v_{i,t}, && (30) \end{aligned}$$

$t = 1,\ldots,T$, $i = 1,\ldots,N$, where $C(L) = \sum_{j=0}^{\infty} C_j L^j$, and $f_i(t)$ denotes the deterministic component (which may be broken as in 16 above), F_t denotes a $(r \times 1)$-vector containing the common factors, with π_i the vector of loadings. Despite the operator $(1-L)$ in equation (28), F_t does not have to be I(1). In fact, F_t can be I(0), I(1), or a combination of both, depending on the rank of $C(1)$. If $C(1) = 0$, then F_t is I(0). If $C(1)$ is of full rank, then each component of F_t is I(1). If $C(1) \neq 0$, but not full rank, then some components of F_t are I(1) and some are I(0). Our analysis is based on the same set of assumptions in Bai and Ng (2004), and Bai and Carrion-i-Silvestre (2005). With a number of assumptions on the loadings and error terms from the above equations we one can continue the estimation of common factors as is done in Bai and Ng (2004). We need to compute the first differences:

$$\Delta y_{i,t} = \Delta f_i(t) + \Delta x'_{i,t}\delta_{i,t} + \Delta F_t\pi_i + \Delta e_{i,t}, \tag{31}$$

and take the orthogonal projections:

$$\begin{aligned} M_i\Delta y_i &= M_i\Delta F\pi_i + M_i\Delta e_i && (32)\\ &= f\pi_i + z_i, && (33) \end{aligned}$$

with $M_i = I - \Delta x_i^d \left(\Delta x_i^{d\prime} \Delta x_i^d\right)^{-1} \Delta x_i^{d\prime}$ being the idempotent matrix, and $f = M_i \Delta F$ and $z_i = M_i \Delta e_i$. The superscript d in Δx_i^d indicates that there are deterministic elements. The estimation of the common factors and factor loadings can be done as in Bai and Ng (2004) using principal components. Specifically, the estimated principal component of $f = (f_2, f_3, \ldots, f_T)$, denoted as $\tilde{f}$, is $\sqrt{T-1}$ times the r eigenvectors corresponding to the first r largest eigenvalues of the $(T-1) \times (T-1)$ matrix $y^* y^{*\prime}$, where $y_i^* = M_i \Delta y_i$. Under the normalization $\tilde{f}\tilde{f}'/(T-1) = I_r$, the estimated loading matrix is $\tilde{\Pi} = \tilde{f}' y^*/(T-1)$. Therefore, the estimated residuals are defined as

$$\tilde{z}_{i,t} = y_{i,t}^* - \tilde{f}_t \tilde{\pi}_i. \tag{34}$$

One can recover the idiosyncratic disturbance terms through cumulation, i.e. $\tilde{e}_{i,t} = \sum_{j=2}^{t} \tilde{z}_{i,j}$, and test the unit root hypothesis ($\alpha_{i,0} = 0$) using the ADF regression equation

$$\Delta \tilde{e}_{i,t}\left(\hat{\lambda}_i\right) = \alpha_{i,0} \tilde{e}_{i,t-1}\left(\hat{\lambda}_i\right) + \sum_{j=1}^{k} \alpha_{i,j} \Delta \tilde{e}_{i,t-j}\left(\hat{\lambda}_i\right) + \varepsilon_{i,t}. \tag{35}$$

We denote by $ADF_{\tilde{e}}^{c}(i)$, $ADF_{\tilde{e}}^{\tau}(i)$ and $ADF_{\tilde{e}}^{\gamma}(i)$ the pseudo t-ratio ADF statistics for testing $\alpha_{i,0} = 0$ in (35), for the model that includes a constant, a linear time trend, and a time trend with a change in trend, respectively. When $r = 1$ we can use an ADF-type equation to analyze the order of integration of F_t as well. However, in this case we need to proceed in two steps. In the first step we regress $\tilde{F}_t$ on the deterministic specification and the stochastic regressors. In the second step we estimate the ADF regression equation using the detrended common factor $\left(\tilde{F}_t^d\right)$, i.e. the residuals of the first step:

$$\Delta \tilde{F}_t^d = \delta_0 \tilde{F}_{t-1}^d + \sum_{j=1}^{k} \delta_j \Delta \tilde{F}_{t-j}^d + u_t, \tag{36}$$

and test if $\delta_0 = 0$ – $ADF_{\tilde{F}}^d(\lambda)$ denotes the pseudo t-ratio ADF statistic for testing $\delta_0 = 0$ in (36).

Finally, if $r > 1$ one should use one of the two statistics proposed in Bai and Ng (2004) to fix the number of common stochastic trends (q). As before, let $\tilde{F}_t^d$ denote the detrended common factors. Start with $q = r$ and proceed in three stages – which are reproduced for completeness:

1. Let $\tilde{\beta}_{\perp}$ be the q eigenvectors associated with the q largest eigenvalues of $T^{-2} \sum_{t=2}^{T} \tilde{F}_t^d \tilde{F}_t^{d\prime}$.
2. Let $\tilde{Y}_t^d = \tilde{\beta}_{\perp} \tilde{F}_t^d$, from which we can define two statistics:

   (a) Let $K(j) = 1 - j/(J+1)$, $j = 0, 1, 2, \ldots, J$:

   i. Let $\tilde{\xi}_t^d$ be the residuals from estimating a first-order VAR in $\tilde{Y}_t^d$, and let

$$\tilde{\Sigma}_1^d = \sum_{j=1}^{J} K(j) \left(T^{-1} \sum_{t=2}^{T} \tilde{\xi}_t^d \tilde{\xi}_t^{d\prime} \right). \tag{37}$$

   ii. Let $\tilde{v}_c^d(q) = \frac{1}{2}\left[\sum_{t=2}^{T}\left(\tilde{Y}_t^d \tilde{Y}_{t-1}^{d\prime} + \tilde{Y}_{t-1}^d \tilde{Y}_t^{d\prime}\right) - T\left(\tilde{\Sigma}_1^d + \tilde{\Sigma}_1^{d\prime}\right)\right]\left(T^{-1}\sum_{t=2}^{T} \tilde{Y}_{t-1}^d \tilde{Y}_{t-1}^{d\prime}\right)^{-1}$.

iii. Define $MQ_c^d(q) = T\left[\tilde{v}_c^d(q) - 1\right]$ for the case of no change in the trend and $MQ_c^d(q,\lambda) = T\left[\tilde{v}_c^d(q,\lambda) - 1\right]$ for the case of a change in the trend.

(b) For p fixed that does not depend on N and T:

i. Estimate a VAR of order p in $\Delta\hat{Y}_t^d$ to obtain $\tilde{\Pi}(L) = I_q - \tilde{\Pi}_1 L - \ldots - \tilde{\Pi}_p L^p$. Filter $\hat{Y}_t^d$ by $\tilde{\Pi}(L)$ to get $\tilde{y}_t^d = \tilde{\Pi}(L)\hat{Y}_t^d$.

ii. Let $\tilde{v}_f^d(q)$ be the smallest eigenvalue of

$$\Phi_f^d = \frac{1}{2}\left[\sum_{t=2}^{T}\left(\tilde{Y}_t^d\tilde{Y}_{t-1}^{d\prime} + \tilde{Y}_{t-1}^d\tilde{Y}_t^{d\prime}\right)\right]\left(T^{-1}\sum_{t=2}^{T}\tilde{Y}_{t-1}^d\tilde{Y}_{t-1}^{d\prime}\right)^{-1}. \tag{38}$$

iii. Define the statistic $MQ_f^d(q) = T\left[\tilde{v}_f^d(q) - 1\right]$ for the case of no change in the trend and $MQ_f^d(q,\lambda) = T\left[\tilde{v}_f^d(q,\lambda) - 1\right]$ for the case of a change in the trend.

3. If $H_0 : r_1 = q$ is rejected, set $q = q-1$ and return to the first step. Otherwise, $\tilde{r}_1 = q$ and stop.

Now, note that the definition of the common factors framework implies that the matrix of projections M_i that is used above cannot depend on i, which means that all elements that are defined in Δx_i^d should be the same across i. There are two different kind of elements in Δx_i^d: (i) the deterministic regressors and (ii) the stochastic regressors. Regarding the latter, BS have shown in their Appendix that the limiting distribution of the statistics do not depend on the presence of stochastic regressors, so that we can ignore the effect of these elements when defining M_i. Unfortunately, this is not true for the deterministic regressors. Thus, to warrant that M_i does not (asymptotically) depend on i one has to assume common break dates, i.e. one has to assume that the break points are the same for all units. This restriction can be seen as a limitation of the analysis, but in fact it is due to the definition of the common factors framework. Thus, (33) specifies a common factor structure for all units, so that f_t cannot depend on i. Looking at the definition of $f_t = M_i\Delta F_t$ one can see that the specification of heterogeneous structural breaks implies that the idempotent matrix M_i depends on i. The only way to overcome this situation is to impose $M_i = M$ $\forall i$ so that the structural breaks are the same for all units.

Next, the limiting distribution of the ADF statistic for the idiosyncratic disturbance term does not depend on the presence of stochastic regressors. Moreover, the presence of changes in level does not affect the limiting distribution of the ADF statistic that is computed using the idiosyncratic disturbance term.

The individual ADF statistics for the idiosyncratic disturbance terms can be pooled to define a panel data cointegration test. Thus, one can define

$$N^{-1/2}Z_{\hat{t}_{NT}}^e(\lambda) - \Theta_2^e(\lambda)\sqrt{N} \Rightarrow N\left(0, \Psi_2^e(\lambda)\right), \tag{39}$$

where the superscript e denotes the idiosyncratic disturbance. The moments $\Theta_2^e(\lambda)$ and $\Psi_2^e(\lambda)$ depend on the deterministic specification used and, except for the case of changes in trend, are the same as the ones for the statistics in Bai and Ng (2004) (where these do not depend on the break fraction λ).

Break point unknown When the break point is unknown we can proceed to estimate it using the infimum functional as described above. However in contrast with case where factors were not present, we have to constrain the (unknown) break point to be common to all units in the panel data set and to estimate both the subspace spanned by the common factors and the idiosyncratic disturbance terms for all possible break points. We then compute the $Z^{e}_{t_{NT}}(\lambda) = N^{-1}\sum_{i=1}^{N} t_{\hat{\rho}_i}(\lambda)$ statistic for each break point using the idiosyncratic disturbance terms and estimate the break point as the argument that minimizes the sequence of standardized $Z^{e}_{t_{NT}}(\lambda)$ statistics.

Chapter IV
Exchange Rates and Exchange Rate Pass-Through - a Glance at the EU New Member States

Abstract

We analyze the exchange rate pass-through from the nominal effective exchange rates to import prices in 8 Central Eastern European countries, taking a separate look at pass-through in commodity and manufacturing sectors. We find that the degree of pass-through was not primarily driven by the choice of the exchange rate regime, and that it tended to be full and instantaneous in a number of countries and sectors while incomplete in the Czech Republic and Slovenia. As there is some evidence exchange rate changes are further passed on to CPI, we speculate on how high and instantaneous ERPT may contribute to the vulnerabilities arising from the ERM II requirements once the countries decide to adopt the euro.

JEL Classification Numbers: F14, F31, F36, F42

Keywords: *CEECs, exchange rates, pass-through, import prices.*

IV.1 Introduction

Exchange rate pass-through into prices (ERPT) has been recognized as an important input for the design of monetary policy. It reflects the interaction of a number of microeconomic factors, such as the market power of producers and importers, market segmentation and competition conditions. Moreover it also reflects a number of macroeconomic issues, such as the credibility of monetary policy and the expectations of future macroeconomic developments.

In this paper, we use data for sectoral import unit values (IUVs) for 8 New Member States from Central and Eastern Europe to determine what part of nominal effective exchange rate developments is passed on to import prices. We adopt different assumptions about the long-run ERPT in order to be able to draw conclusions on the instantaneous pass-through. We find evidence of full or close to full short run ERPT in Hungary, both in commodity and manufacturing sectors; in Poland and Lithuania especially in the commodity sectors and in Estonia in the manufacturing sectors. Pass-through for all sectors together tends to be full in all three of these countries, and with lesser confidence in Estonia. On the other hand we find evidence of incomplete pass-through in the Czech Republic, Slovenia and Latvia, especially in the manufacturing sectors. Most of the ERPT estimates appear stable within the examined period. The consequences of high pass-through in the short term may include potential vulnerabilities in future attempts in joining the common European currency.

Our findings, though seriously limited by the quality of the data, have a number of important implications for the countries - once in the euro waiting room - the ERM II. As Central Eastern European economies (CEECs) catch up in terms of development, productivity growth, through the Balassa-Samuelson effect,[1] can be expected to cause the real exchange rates in developing, strongly growing economies to appreciate in the longer term. This can take place through two channels - the nominal exchange rates and (relative) price inflation. A country with a fixed exchange rate policy, in principle leaves one of the channels closed. Thus recent higher (than in inflation targeting CEECs) inflation in the Baltic States should not come as a surprise. However, as for these countries inflation has been the main obstacle to joining the European common currency in 2007, any additional rise in prices caused by temporary nominal exchange rate shocks (e.g. fluctuations in the euro/dollar rate), could bring with them the risk of failing to enter the Economic and Monetary Union (EMU) on a certain date. On the other hand, although this risk seems mitigated in inflation targeting regimes, the fact of a floating currency brings about potentially large fluctuations in exchange rates, which can in turn feed in to prices, thus must be taken into account by the monetary authorities. The importance is emphasized by the fact that the Maastricht inflation criterion has been shown to be applied with great scrutiny when Lithuania attempted to join the EMU starting 2007.

As we will show, Hungary, Poland and Lithuania tend to have high short-term values of ERPT into import prices. This means a change in the nominal (effective) exchange rate tends to be passed on fully to the local import price within one month, though admittedly in some cases this tendency is stronger in the commodity sectors than in manufactured goods. Thus an adverse (effective) exchange rate development, both in case of a currency board such as Lithuania and in case of flexible regimes such as in Poland and Hungary may potentially lead to a failed attempt to join the euro. This reasoning involves the assumption

[1]The effect of productivity growth differentials between tradeables and non-tradeables in different countries resulting in bilateral changes in the real exchange rate, was first noticed in the works of David Ricardo, and among others Roy Harrod and later independently by Balassa (1964) and Samuelson (1964). The most common nomenclature refers to it as the "Balassa-Samuelson effect".

that at least part this change is further passed on into the consumer price index (CPI), which as will show, there are grounds to believe is the case. We find indications that in these three countries, ERPT into the overall price level is non-zero in the short run and full or close to full in the long run. One of the purposes of the exercise is to highlight some problems of the vulnerabilities of both regimes under ERM II. On the other hand, the finding of rather low values of ERPT in Czech Republic and Slovenia, makes them less sensitive to the issue, though admittedly for the latter, on which the decision to enter to the Eurozone has already been made, the issue is of lesser relevance.[2]

The paper is structured as follows. In Section IV.2 we describe the basic setup of our ERPT estimation. Next, in Section IV.3 we discuss other empirical work in this area. Turning to Section IV.4 we describe the data including the construction of the variables, trade patterns of the CEECs and devoting special attention to exchange rate developments. Section IV.5 presents the estimation method, and the following sections summarize and discuss the results.

IV.2 Exchange Rate Pass-Through into Import Prices

By definition[3], import prices for any type of good j, MP_t^j are a transformation of export prices of a country's trading partners XP_t^j using the bilateral exchange rate ER_t:

$$MP_t^j = ER_t \cdot XP_t^j \tag{1}$$

[2] The decision to approve Slovenia's accession to the EMU on the 1st of January 2007 was made by the Council of the EU on July 11th, 2006.

[3] This section is based on Campa, Goldberg and Gonzalez-Minguez (2005), CGM hereafter. It is similar to the relevant section in De Bandt, Banerjee and Kozluk (2006) and thus in Chapter III of this thesis.

In logarithms (depicted in lower case):

$$mp_t^j = er_t + xp_t^j \tag{2}$$

Where the export price consists of the exporters marginal cost and a markup:

$$XP_t^j = FMC_t^j \cdot FMKUP_t^j \tag{3}$$

So that in logarithms we have:

$$xp_t^j = fmc_t^j + fmkup_t^j \tag{4}$$

Substituting for xp_t^j into equation (2) yields:

$$mp_t^j = er_t + fmkup_t^j + fmc_t^j \tag{5}$$

for each sector in each of the countries.

The literature on industrial organization yields insight into why exporters of a given product may decide to absorb some of the exchange rate variations instead of passing them through to the price in the importing country currency. This responsiveness of the mark-up is determined primarily by the competition conditions, such as market share, competition structure, the ability to discriminate between markets etc. that the producer or exporter faces in a given market. If the pass-through is complete (producer currency pricing), their mark-ups will not respond to fluctuations of the exchange rates, thus leading to a pure currency translation. At the other extreme, they can decide not to vary the prices in the destination country currency (local currency pricing or pricing to market) and absorb the fluctuations within the mark-up.

Thus, mark-ups of the exporting industry are assumed to consist of a component specific to the type of good, independent of the exchange rate and a reaction to exchange rate movements:

$$fmkup_t = \alpha + \Phi er_t \tag{6}$$

Also important to consider are the effects working through the marginal cost. These are a function of demand conditions in the importing country; marginal costs of production (labor wages) in the exporting country and the commodity prices denominated in foreign currency:

$$fmc_t = c_0 \cdot y_t + c_1 \cdot fw_t + c_2 \cdot er_t + c_3 \cdot fcp_t \tag{7}$$

Substituting (6) and (7) into (4), we have:

$$mp_t = \alpha + \underbrace{(1 + \Phi + c_2)}_{\beta} er_t + c_0 \cdot y_t + c_1 \cdot fw_t + c_3 \cdot fcp_t + \varepsilon_t \tag{8}$$

where the coefficient β on the exchange rate er_t is the pass-through elasticity.[4] Therefore estimated pass-through elasticities are a sum of three effects:

- effects of the unity translation effects of the exchange rate movement,
- the response of the markup in order to offset this translation effect,
- the effect on the marginal cost that are attributable to the exchange rate movements, such as the sensitivity of input prices to exchange rates.

And thus does not necessarily equal 1. In the 'integrated world market' approach adopted

[4]Obviously, this is a simple approach, with a highly reduced form representation, where one can have no hope of identifying Φ from c_2.

by the CGM specification, $c_0 \cdot y_t + c_1 \cdot fw_t + c_3 \cdot fcp_t$, is independent of the exchange rate and can be thought of as the opportunity cost of allocating those same goods to other customers, as reflected in the world price of the product fp_t in the world currency.[5]

IV.2.1 Long run ERPT

From the previous section we can re-write the final equation (8) as follows:

$$mp_t = \alpha + \beta \cdot er_t + \gamma \cdot fp_t + \varepsilon_t \tag{9}$$

which gives the long run relation between the import price, exchange rate and a measure of foreign price.

Exchange rate pass-through is usually specified as the effect of exchange rate fluctuations on import prices (the so-called 'first stage' pass-through, which we will mostly focus on in this paper) or on domestic consumer prices ('second stage' pass-through). Since ERPT (both into import prices and into domestic prices) is a channel linking exchange rates to prices, it is often named as one of the key determinants of monetary policy design. The level of ERPT is an important input for inflation forecasting and monetary policy. There is a vast literature on optimal monetary policy, which takes into account ERPT into

[5] The integrated market hypothesis in CGM is based on the assumption that there exists a single world market for each good. Therefore, regardless of the origin of the product, on the world market, it has one world price. This price constitutes the opportunity cost of selling to a local market. Thus, in the CGM setup for the integrated market it proxies for the foreign price. The currency denomination does not in fact matter, as long as the exchange rate for the local currency is taken vis-a-vis this 'world' currency. In the CGM case the extra-euro area imports into the euro area denominated in US dollars are taken as a proxy for the world price. CGM propose an alternative specification - the 'segmented world market' where both the exchange rate and foreign price are, respectively, the trade-weighted average of exchange rates with the 5 biggest trading partners and the trade-weighted price index of imports from the same trading partners. CGM test and reject the hypothesis that the segmented market specification for most of their EMU sample combinations adds any information relative to the integrated market hypothesis. As will become clear later, in this paper we take an intermediate approach.

import prices and final consumer prices.[6] Most generally, earlier literature (see Devereux and Engel, 2002; Adolfson, 2001) tends to suggest that as incomplete ERPT lowers the relative importance of the exchange rate as the shock transmission channel, and allows the attainment of low price volatility together with large exchange rate fluctuations. However, Corsetti and Pesenti (2005) and Smets and Wouters (2002) show that internal stabilization policies are not successful in the presence of imperfect pass-through if they do not take into account the cost of excessive exchange rate volatility. High fluctuations in the exchange rate, in part engineered by central banks exploiting the above finding, together with the firms' inability to fully adapt prices in response to these fluctuations (due to sticky prices - the main underlying source of low ERPT in most models), may cause producers to charge inefficiently high prices in order to assure a certain revenue level.

Moreover a large part of the literature focuses on short run ERPT assuming full pass-through in the long run. Notably Corsetti, Dedola and Leduc (2006) distinguish two sources of imperfect pass-through: price stickiness which is a short run phenomenon and thus does not affect the long run, and market segmentation which allows for an imperfect pass-through even in the long run.

Finally, there is a large degree of endogeneity in the observed ERPT and monetary policy. That is, pricing strategies of firms depend not solely on competition conditions in the market, but also on monetary policy, or rather the expected monetary policy and the policy makers' credibility. On the other hand monetary policy can be expected to depend on a consideration of ERPT and the effect of exchange rate variability on prices. Therefore it will be interesting to compare results for countries with very different exchange rate regimes. On the downside, there are a number of problems with the data for CEECs, which

[6] Section 2 of Chapter III reviews the literature in more detail.

we describe in detail in Section IV.4.

IV.2.2 ERPT in the short run

Both economic theory and relevant tests lead us to think of each of the series (import price, exchange rate and world price) as being characterized by a unit root. Moreover, the underlying specification of the model being an equilibrium relationship in 'levels' suggests that these series can be expected to be cointegrated. If we can find evidence of the latter, we will proceed by estimating an error correction model (ECM), as in Engle and Granger (1987), of the the following form:

$$\Delta mp_t = a+\sum_{k=0}^{K_1} b_k\cdot\Delta er_{t-k}+\sum_{k=0}^{K_2} c_k\cdot\Delta fp_{t-k}+\lambda(\underbrace{mp_{t-1}-\alpha-\beta\cdot er_{t-1}-\gamma\cdot fp_{t-1}}_{ECM})+u_t \quad (10)$$

where the estimated coefficient b_0 will yield the instant reaction of the price level to the exchange rate movement, and thus the short run or the instantaneous ERPT. The total number of lags is selected automatically, starting from the maximum of 4 and dropping lags until the last is significant, or the simultaneous model is obtained, which in our estimates leaves only simultaneous effects in the vast majority of cases.

IV.3 Empirical literature

Empirical investigations of ERPT have been given much attention in the recent literature (see for example Campa and Gonzalez-Mingues, 2006; Campa, Goldberg and Gonzalez-Minguez, 2005; De Bandt, Banerjee and Kozluk, 2006; Marazzi et al., 2005). But while ERPT has been subject to a relatively large amount of empirical analysis for the industrialized economies, and even some developing countries (see for example Frankel, Parsley and

Wei, 2005), rather little work has been done on the Central Eastern European transition economies. To our knowledge, there are two papers concerning the issue of ERPT into the aggregate price indices, of which both focus on the 'secondary' ERPT, that is pass-through into domestic prices. Davras (2001) attempts to examine ERPT into domestic fundamental prices (i.e. domestic price aggregates which exclude food, energy and administered prices) for Czech Republic, Hungary, Poland and Slovenia in the second half of the 1990s. The author attempts to estimate impulse-response functions from vector auto-regressions (VARs) and a version of the error-correction (ECM) specification. The results obtained tend to suggest short run ERPT in the Czech Republic was close to zero, while in the other countries above zero, significantly lower than one (point estimates of 0.1 to 0.3, though not significantly different from each other). The robustness of the analysis is weakened by a scarcity of observations (the short time series consist of twelve quarterly observations for the Czech Republic and 29 for the other countries), and by problems with estimated parameter instability, possibly due to the fact that the sample period includes the turbulent period of relatively early transition.

Coricelli, Jazbec and Masten (2006) take a different approach. They use a cointegrated VAR framework to estimate ERPT into consumer price inflation in the same four CEECs (Czech Republic, Hungary, Poland and Slovenia) in the period 1993m1:2002m5. They supplement the set of explanatory variables by industrial production and the interest rate differential (with respect to Euribor) in order to account for cyclical fluctuations.[7] They find estimates of long run pass-through into the CPI in Slovenia, Hungary and Poland

[7]In a working paper version of the paper, Coricelli, Jazbec and Masten (2004), the authors attempt to support their conclusions with I(2) analysis. However, ADF tests performed on the variables over our sample (1999m1:2005m12) do not confirm the hypothesis of the variables being I(2). This may be because the data starts rather late compared to the Coricelli, Jazbec and Masten (2004) paper and thus does not cover the turbulent early period of transition.

insignificantly different from one, while significantly lower than one in the Czech Republic. Short-term, or instantaneous pass-through rates from exchange rate changes to inflation differentials tend to be highest in Slovenia, lower, though admittedly not significantly lower for Poland and Hungary, and the lowest in the Czech Republic. Generally, the authors tentatively conclude that accommodative monetary policy causes much of the inflation pressures in the CEECs.

Both investigations in the field of ERPT in Central and Eastern Europe are rather rough and only very indicative results are obtained. The major determinant is certainly the lack of long time series over a stable period. Consequently, in this paper, we want to introduce a number of new issues. First, we attempt to distinguish between ERPT in different types of goods, mainly between primary goods and processed or manufactured goods, which empirical estimation for the industrialized countries has shown to have different importer pricing behavior.[8] We use a data set which, starting in 1999, is characterized by much more stable exchange rate, price and monetary policy developments. Admittedly many of the problems encountered in previous work still linger and we have preferred in consequence to adopt a simple estimation approach.

IV.4 Data

For the purpose of this paper we use data for Import Unit Values (IUVs) and exchange rates of local currencies against the euro and US dollar. The data is taken from Eurostat

[8]One of the reasons why one can expect ERPT into import prices to be higher (at least in the short run) in primary and commodity sectors is the fact that trade in these sectors is to a large extent done through future contracts, upon a 'world' price, set in one of the main currencies (usually USD). Hence there is little room for adjustment, and the IUV, which is reported in local currency is affected directly by the change in exchange rate, consisting to a large extent of a product of the incidental(upon the declaration of import) exchange rate and the pre-agreed contract price. Anecdotal evidence and empirical estimates in papers quoted later in this work seem, somewhat unsurprisingly, to point to more flexibility in the manufactured sectors.

(COMEXT) and described in more detail in Appendix A. The observations concern total trade and trade in nine individual 1-digit SITC sectors for 8 of the Central and Eastern European EU members (Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia and Slovenia).[9] They range from 1999m1 to 2005m12.

The issues related to using IUVs in ERPT estimation are mostly related to the underlying properties of the indexes. These have been well overviewed in Campa and Gonzalez-Minguez (2006) and also in De Bandt, Banerjee and Kozluk (2006) and concern the issue of aggregation. Generally IUVs have the downside of measuring the value of a basket which is varying at each moment of observation. That is the IUV may change from month to month not solely because the price of a good has changed, but because the composition of goods imported in one month may change in another, partly perhaps, because of the exchange rate movements. Moreover, using import unit values entails essentially looking at the values of kilograms of each type of goods, be it commodities, automobiles or computers.

The problems described above are standard and well identified in the analysis of ERPT. However, in the case of the CEECs there are a number of additional issues that make the analysis more complicated and troublesome. First of all, for this group of countries the IUVs are not directly provided by the Eurostat, and thus we have had to compute them by dividing the value of trade by the quantity. This renders the series to appear less smooth than the series for "old" EU members, as they are not aggregated up from lower classification levels. Moreover, the CEECs have been undergoing a period of transition to market economies since the early 1990s. Prices were gradually liberalized, economies opened up to international trade and integrated both with the global and European economies. All of the countries

[9]The series SITC 1 (Beverages and Tobacco) IUVs appear to be corrupt (large break in mean and variance) for six of the countries, and have been excluded from the sample, especially as the series are relatively short.

had to set up new monetary regimes, and many have switched between various arrangements for their exchange rates. Moreover, on the 1st of May 2004 they joined the EU. Each of these changes may be suspected to constitute a break in the sample. The issues of exchange rates in CEECs and trade integration with the EU are given more attention in the following sections, but generally, the transition, exchange rate regime changes and EU integration justify using the sample provided in COMEXT, that is starting 1999m1. Going back earlier, would force us to deal with the problem of the introduction of the common currency in the biggest CEECs' trading partner - the EMU, leading us to a suspicion of a possible break.

In our exercise we distinguish imports from two sources - the (current) EU and from outside of it. This means the import price (in local currency) is a weighted average of the import price of imports from inside the EU pm^{EU} and from other sources pm^{nonEU} (both in local price):

$$pm_{it} = pm_{it}^{EU} * \frac{imp\_val_{it}^{EU}}{imp\_val_{it}^{TOT}} + pm_{it}^{nonEU} * \frac{1 - imp\_val_{it}^{EU}}{imp\_val_{it}^{TOT}} \tag{11}$$

where imp_val^{X} is the value of imports from X - in our case the EU-25 (EU) and the total outside world (TOT). Consequently, the exchange rate is an effective exchange rate, i.e. the weighted average between the rates against the euro and the dollar:

$$er_{it} = er_{it}^{EUR} * \frac{imp\_val_{it}^{EU}}{imp\_val_{it}^{TOT}} + er_{it}^{USD} * \frac{1 - imp\_val_{it}^{EU}}{imp\_val_{it}^{TOT}} \tag{12}$$

The foreign price is proxied by the 'world price' which consists of the weighted average of the the EU price of imports originating from within and from outside the EU:

$$fp_{it} = fp_{it}^{EU} * \frac{imp\_val_{it}^{EU}}{imp\_val_{it}^{TOT}} + fp_{it}^{nonEU} * \frac{1 - imp\_val_{it}^{EU}}{imp\_val_{it}^{TOT}} \tag{13}$$

This setup is equivalent to assuming that trade done with the EU partners is denominated in euros, and with outside partners in dollars. The assumption may seem questionable, but is actually broader than the "integrated market" specification of Campa and Gonzalez-Minguez (2006)[10] and in a sense a natural extension. Moreover due to specific exchange rate regimes in some of the CEECs, including fixing to the euro or dollar and switching between regimes, the use of a type of effective exchange rate is necessary. Finally, one may wonder how reasonable it is to include for instance the UK together with the euro area, but having tried varying this approach by using the EMU or the EU-15 instead of the EU-25 did not affect the results.[11] One of the reasons could be the small amount of trade done with the non EMU members of the "old" EU relative to the dominating trade partners - Germany, Italy and France.

IV.4.1 Exchange rate regimes and monetary policy

The choice of exchange rate regime is determined by the choice of monetary policy. In the simple Mundell-Flemming framework, with capital flows, in a small (relative to global capital) economy, in order to pursue a fully independent monetary policy, focusing on domestic shocks and pursuing a domestic target, like for instance pure inflation targeting, the authority must cease attempts to control the exchange rate and allow it to float. At the other extreme fixing its currency to another will require the adoption of the monetary policy of the target country (in order for the peg to be sustainable), and the adjustments to go through the price channel. In a sense in the real exchange rate link of the nominal exchange rate and the nominal price level, there is only one degree of freedom, thus for instance an

[10]Our approach is actually a variation of the Campa and Gonzalez-Minguez (2006) 'segmented market' specification, as we use the effective exchange rate and effective import prices.

[11]Some of the results in the two specifications are compared in Appendix B.

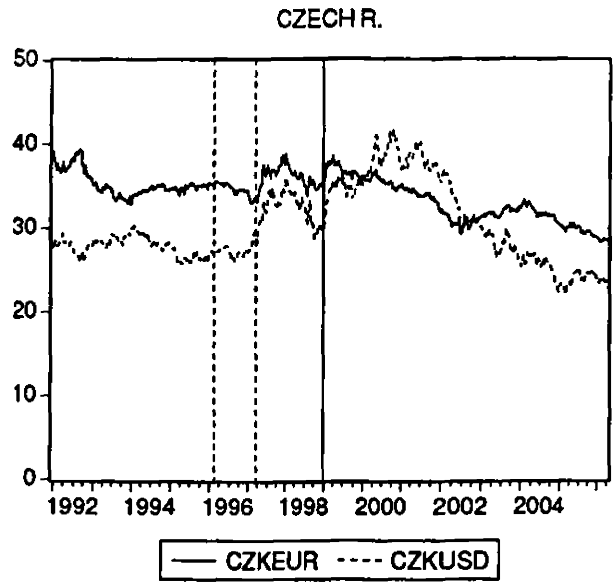

Figure 1: Weekly CZK exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 1 Jan 1991 - fixed peg against basket 65% DEM, 35% USD. C.B. intervention band ±0.5%; 28 Feb 1996(*) band widened to ±7.5%; 26 Mar 1997(*) (managed) float; Vertical solid line indicates 1999m1d1. Note that the CSK (Czecho-Slovak koruna) was replaced by the CZK on the 8^{th} of February 1993 at par. Source: Thomson DataStream, transitioneconomies.blogspot.com

inflation target serving as a nominal anchor for expectations, determines the absorption of shocks through a flexible exchange rate and vice versa, a fixed nominal exchange rate, will require adjustments through the level of prices.

In practice, many intermediate arrangements exist, but any type of policy that should claim to overcome the above explained trade-off is unsustainable without strict capital controls. As we will see in the next section, CEECs exhibit a wide range of regimes, spanning from full float inflation targeting through various types of intermediate regimes to a fully fixed exchange rate.

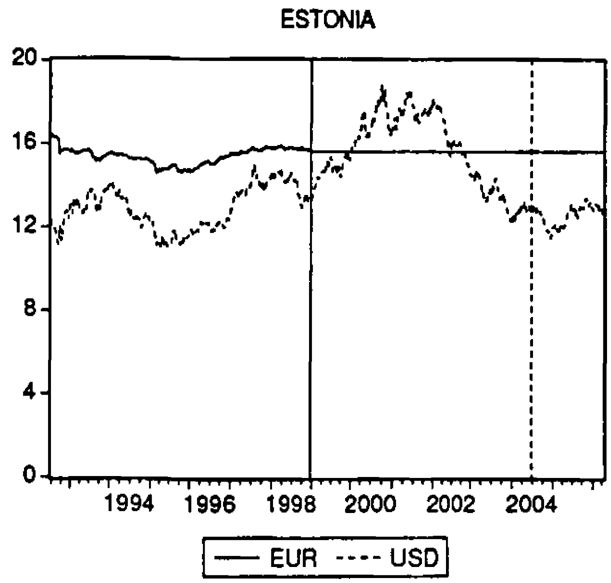

Figure 2: Weekly EEK exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 20 Jun 1992 currency board with DEM (subsequently EUR); 27 Jun 2004(*) ERM II; Vertical solid line indicates 1999m1d1. Source: see above

IV.4.2 Exchange rate regimes in CEECs

The time-lines of exchange rate regimes and the historical exchange rates against the USD and the euro in the CEECs are described in Figures 1 to 8. In the 1990s' early years of transition the Central and Eastern European economies have experienced high inflation, and partly as a result of this, frequent monetary regime changes. The quest for credibility in monetary policy caused the policy makers to take very different, and in some cases varying routes. There were notable attempts to 'borrow' credibility by fixing domestic currencies to well-recognized external ones, such as the USD or DEM, or a basket of currencies, however the way to proceed was uncertain, and the exchange rate developments often turbulent. The strongest commitments to fixing the currency were visible in the adoption of currency boards in Estonia in 1992 and in Lithuania two years later. Latvia tightly pegged its lat to the SDR basket, and maintained the arrangement till 2005. Poland, having started off

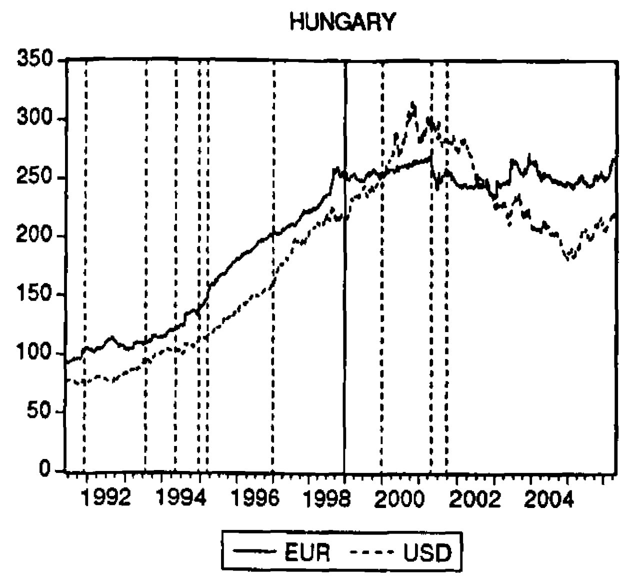

Figure 3: Weekly HUF exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 9 Dec 1991(*) - peg to basket 50%ECU, 50% USD, C.B. intervention band ±0.3%; 2 Aug 1993(*) basket changed to 50%DEM, 50%USD; 16 May 1994(*) - basket changed to 70%ECU, 30%USD; 22 Dec 1994(*) - intervention band widened to ±2.25%; 13 Mar 1995(*) - crawling peg/band to basket. Bands at ±2.25%. Crawl rate decreasing; 1 Jan 1997(*) basket changed to 70%DEM, 30% USD; 1 Jan 1999(*) basket changed to 70%EUR, 30% USD; 1 Jan 2000(*) basket changed to 100% EUR; 3 May 2001(*) band widened to ±15%; 1 Oct 2001(*) fixed horizontal band ±15%. Notes: Devaluations in early 1990s. Vertical solid line indicates 1999m1d1. Source: see above

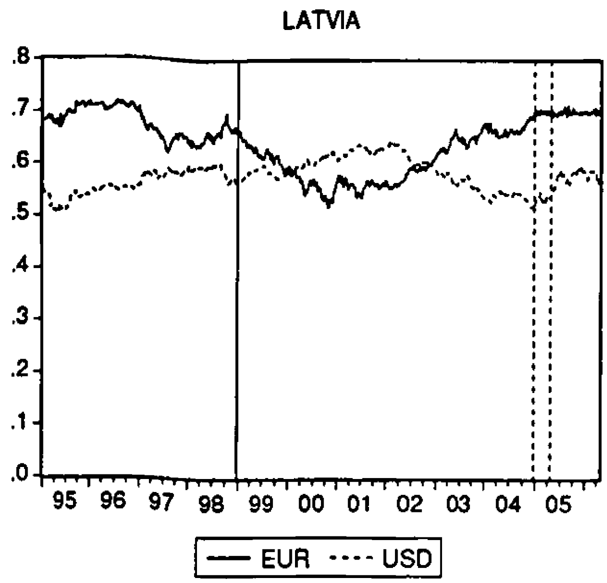

Figure 4: Weekly LAT exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 20 Jul 1992 managed float; 12 Feb 1994 fixed peg to SDR basket. C.B. margin ±1%; 1 Jan 2005(*) fixed peg to EUR. C.B. margin ±1%; 2 May 2005 ERM II; Vertical solid line indicates 1999m1d1. Source: see above

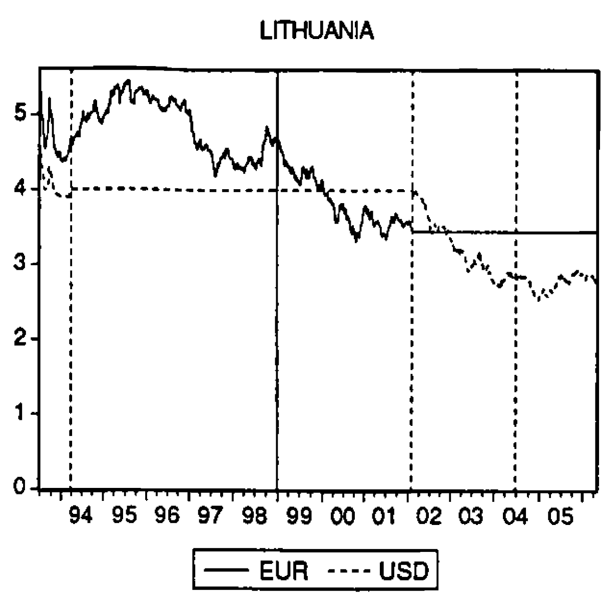

Figure 5: Weekly LTL exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 1 Jul 1992 managed float; 1 Apr 1994(*) currency board with USD; 2 Feb 2002(*) currency board with EUR; 27 Jun 2004(*) ERM II; Vertical solid line indicates 1999m1d1. Source: see above

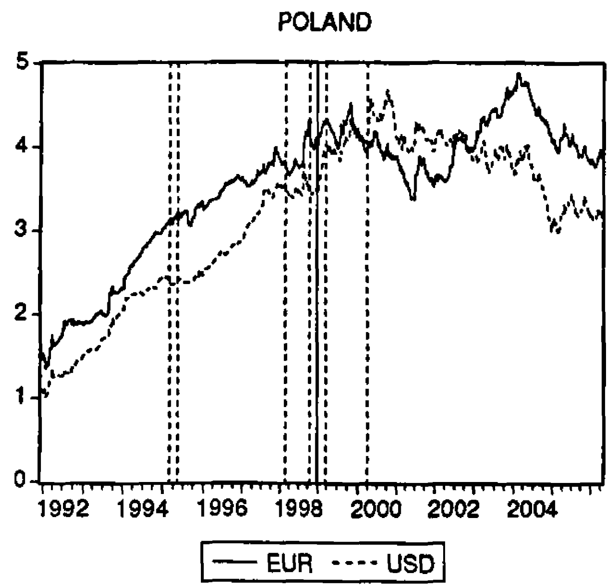

Figure 6: Weekly PLN exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 1 Jan 1990 fixed to USD; 16 May 1991 fixed to basket (45% USD, 55% DEM+GBP+FF+CHF); 14 Oct 1991 crawling peg to (same) basket. Crawl rate decreasing, C.B. intervention margin ±0.5%; 6 Mar 1995(*) C.B. intervention margin widened to ±2.0%; 16 May 1995(*) crawling band (to basket) ±7%. Crawl rate decreasing; 26 Feb 1998(*) band widened to ±10%; 28 Oct 1998 band widened to ±12.5%; 1 Jan 1999(*) basket changed to 55% EUR, 45% USD; 25 Mar 1999(*) band widened to ±15%; 12 Apr 2000(*) free float (but C.B. reserves extraordinary right to intervene) Notes: Devaluations in early 1990s. Vertical solid line indicates 1999m1d1. Source: see above

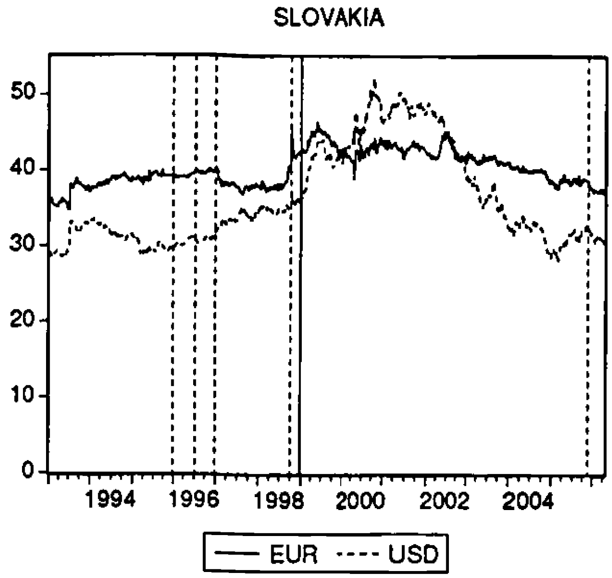

Figure 7: Weekly SKK exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 1 Jan 1991 fixed peg against basket 60% DEM, 40% USD, C.B. intervention band ±1.5%; 1 Jan 1996(*) band widened to ±3%; 16 Jul 1996(*) band widened to ±5%; 1 Jan 1997(*) band widened to ±7%; 1 Oct 1998(*) managed float; 25 Nov 2005(*) - ERM II; Vertical solid line indicates 1999m1d1. Note that the CSK (Czecho-Slovak koruna) was replaced by the SKK on the 8^{th} of February 1993 at par. Source: see above

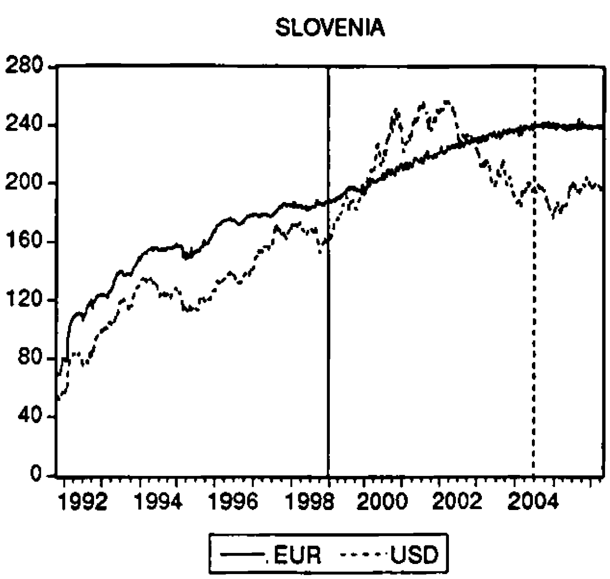

Figure 8: Weekly SIT exchange rates against ECU/EUR (solid line) and USD (dotted). Vertical lines indicate dates described below with (*): 8 Oct 1991 managed float; 27 Jun 2004(*) ERM II; Vertical solid line indicates 1999m1d1. Source: see above

	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
CZ	-	9.9	9.5	8.8	8.5	10.7	2.1	3.9	4.7	1.8	0.1	2.8	1.9
ES	89.8	47.7	28.8	23.1	10.6	8.2	3.3	4	5.7	3.6	1.3	3.1	4.1
HU	22.5	18.9	28.3	23.6	18.3	14.2	10	9.8	9.2	5.3	4.6	6.8	3.6
LV	108.8	35.9	25	17.6	8.4	4.7	2.4	2.7	2.5	1.9	2.9	6.2	6.8
LI	410.2	72.2	39.7	24.6	8.9	5.1	0.8	1	1.3	0.3	-1.2	1.2	2.7
PL	36.9	33.3	28.1	19.8	15.1	11.7	7.3	10.1	5.5	1.9	0.8	3.6	2.1
SK	-	13.4	9.9	5.8	6.1	6.7	10.6	12	7.3	3.3	8.6	7.6	2.7
SN	31.7	21	13.4	9.9	8.4	7.9	6.2	8.9	8.4	7.5	5.6	3.6	2.5

Table 1: Annual CPI inflation. Source: IMF (International Financial Statistics)

by attempting to fix to the USD, and forced to frequently devalue its currency in the early 1990s followed the path of gradual increase in its exchange rate flexibility in order to reach a practically full float by the end of the decade. The Czech Republic, Slovakia and Slovenia closed the decade with managed float regimes of very different degrees of tightness, and while Slovenia introduced this arrangement since the establishment of its currency, allowing for a gradual depreciation of the tolar, the other two went through years of pegging to currency baskets. Finally Hungary also went through crawling pegs, in order to finish with a wide horizontal band.

Despite the different regimes adopted, by the late 1990s all of the 8 EU New Member States managed to stabilize inflation quite successfully (see Table 1). However, we still can observe very different monetary policy regimes from inflation targeting with a fully floating exchange rate to currency boards with a full fix.

Going into more detail about the period relevant for our data set, i.e. starting 1999m1, the exchange rate regimes were still subject to some changes, albeit to a lesser extent than in the previous periods. This makes the analysis and comparison of ERPT estimates in the CEECs more complicated and questionable but also, in a sense, more interesting.

The Czech Republic, Slovakia and Slovenia essentially maintained a managed floating exchange rate in the analyzed period, the second and third joining the ERM II in 2005 and 2004 respectively. Poland entered 1999 with a crawling (at a decreasing rate) peg against a basket of currencies (55% EUR, 45% USD). The intervention band was ±12.5% and was subsequently widened in March 1999 to ±15%. One year later the central bank switched to a fully floating exchange rate, albeit reserving itself the right of intervention in extraordinary situations.

In Hungary the regime was a crawling (at a decreasing rate) band to a euro-dollar basket with ±2.25% margins. One year after (2000m1) the basket was replaced with the euro as the reference currency, a year later the band was widened to ±15%, and in October 2001 the crawl of the peg ceased, thus turning to a fixed horizontal band.

As for the Baltic States, all three maintained relatively fixed exchange rate arrangements. Estonia, being the first to introduce a currency board with the DEM, and subsequently the euro, entered the ERM II in 2004. Lithuania, which previously chose to fix its currency to the USD, switched the fix to the euro in early 2002 with the objective of future EMU membership. It also joined the ERM II in 2004. Finally Latvia maintained a tightly fixed peg (with ±1% bands) to the virtual SDR, or in other words a basket of international currencies until the end of 2004, switching then to the euro as the reference currency and joining the ERM II four months later. The last, to date, to join the waiting room for the euro, was Slovakia, which entered the ERM II in November 2005.

The issue of changes in exchange rate regime may pose a problem for our estimation. The sample is too short be divided into sub-samples, but we will use recursive estimation as a guide to the trustworthiness of the results.

IV.4.3 CEECs - trade patterns

In practically all the SITC sectors, aside perhaps from sector 3 (Mineral fuels), the 8 EU New Member States import the majority of goods from other EU states. Their integration with the EU, already high in 1999, increased further during the run up to EU membership. As visible in Figures 9 and 10 the share of imports originating from outside the EU, in individual sectors, resembles that of the entire EU. Nevertheless, the CEECs trade more intensively with EU members than the 'old' EU, except for commodity sectors SITC 2 (Crude materials) and again SITC 3 (Mineral fuels).

IV.4.4 CEECs and the Euro

Part of the accession agreement which the CEECs had to sign in order to join the EU, entails the adoption of the common currency 'as soon as they are ready'. No opt-out clause has been put in place for these countries, and in the mid to late 1990s all countries seemed rather determined to join. In fact many analysts predicted that due to large potential benefits there may be a rush to the euro. The situation however, changed rather drastically. At the date of entry to the EU, all but Poland, of the 8 countries were pursuing an official target date for the adoption of the common currency. However it seems now, that in 2007 only Slovenia will join, while most other countries have either withdrawn from or postponed their target dates. While Lithuania made an attempt to join on the same date, its participation was ruled out by the European Commission on the grounds of missing the Maastricht inflation criterion by a marginal amount.[12] Estonia, experiencing higher inflation, decided beforehand not to pursue the earlier target date i.e. 2007, partly in fear

[12]The criterion of inflation lower than the average in the three EU countries with lowest inflation +1.5 percentage point. Notably, when Lithuania failed to fulfil the criterion by a narrow margin, i.e. in mid 2006, Poland was among the three countries upon which the cut-off value was calculated.

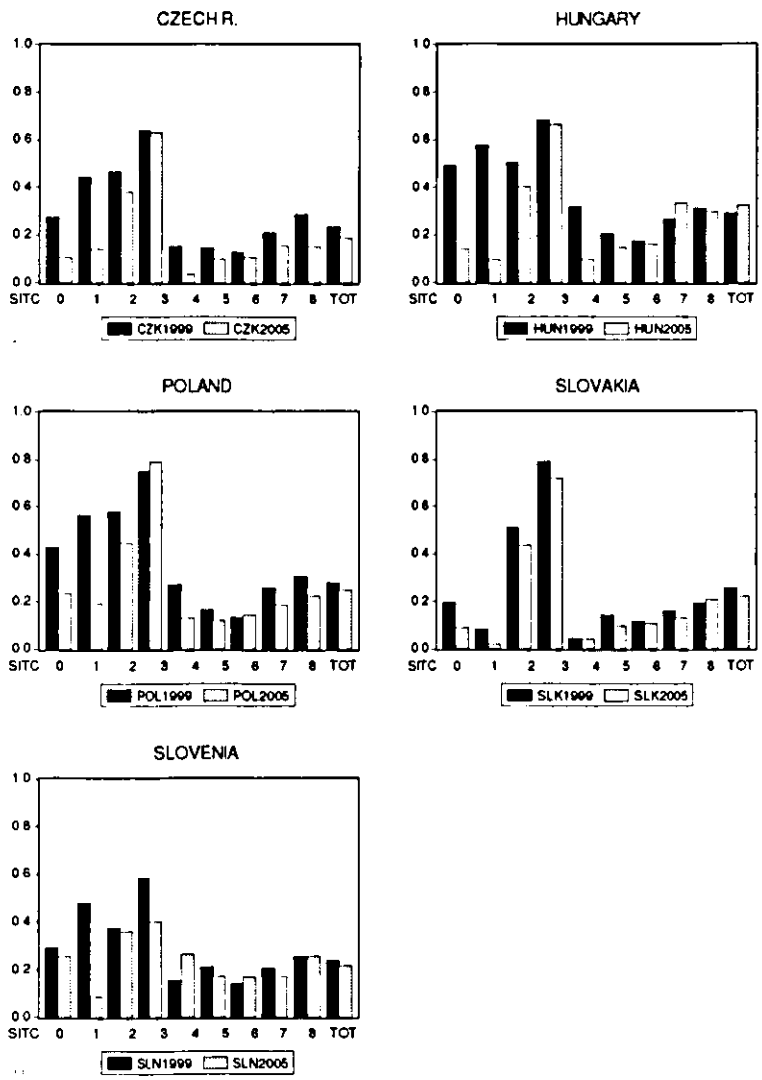

Figure 9: Share of Extra-EU25 imports in the value of total imports - SITC sectors 0 to 8 and TOTAL trade. Averages for 1999 and 2005.

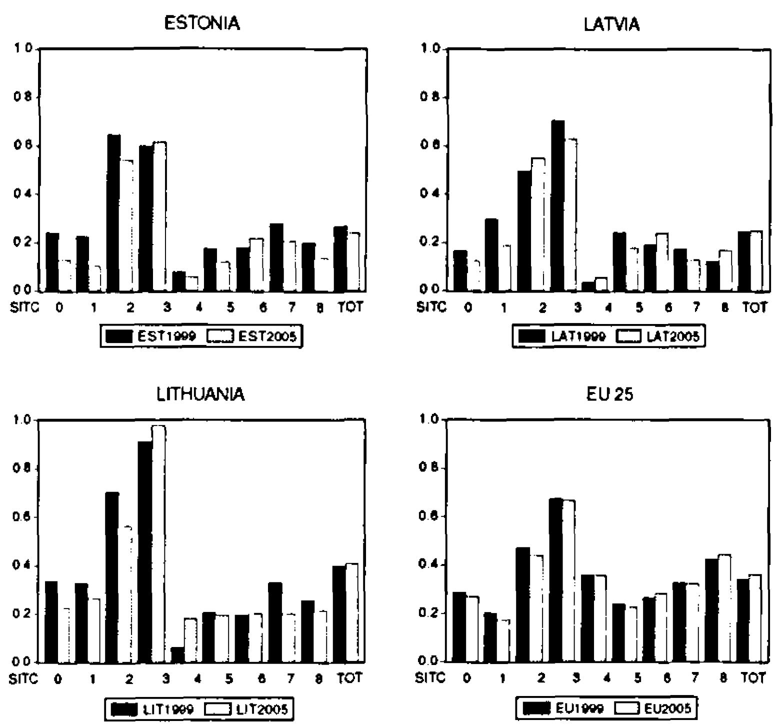

Figure 10: Share of Extra-EU25 imports in the value of total imports - SITC sectors 0 to 8 and TOTAL trade. Averages for 1999 and 2005.

of the adverse signal a rejected attempt would make. Together with Latvia, all three remain in the ERM II and now officially declare target dates in the area of 2008-2009, though the persistent higher inflation makes these dates questionable. Slovakia, the last to join the ERM II, officially aims at 2009, while the Czech Republic and Hungary withdrew from their previous declarations, and now follow Poland's policy of not setting an official target date. Contrary to Slovenia, the Baltic States, and Slovakia these countries do not seem determined to rush into the euro, partly because of lack of political support, but more probably due to the reluctance to reduce fiscal deficits. Nevertheless declarations of officials from these countries rather exclude any attempt to adopt the euro before 2011. Finally, there does not seem to be pressure from the EU itself for the CEECs to adopt the euro, which together with the previously described developments means that for most CEECs the issue of exchange rate management and monetary policy will remain vital for the coming years. ERPT may prove important both in the years they follow an independent monetary policy, as well as once they are in the euro waiting-room - the ERM II.

IV.5 Estimation Methods

As described in the previous section, there are several problems concerning the data on IUVs for the New Member States. The shortness of the available sample puts in question whether one can claim to be able to capture the "long run", that is the levels cointegration relationship as in equation (9). The 6 years we are dealing with may just be too short a sample for the actual long run to reveal itself. In order to somewhat circumvent this we propose using two alternative strategies. Firstly, we nevertheless attempt estimating the long run relationship with the available series (we will refer to this as the 'estimated long-run' specification), and second we in a sense impose a full pass-through in the long run ('imposed long-run'). In both cases, we test for cointegration between the variables by looking at the obtained error correction terms, and subsequently use them for estimating the instantaneous, short run ERPT as represented by b_0 in equation (10).

The assumption of a full pass-through in the long run, though not necessarily correct,[13] can be backed with several arguments. First of all, as mentioned in Section IV.2 a large amount of macroeconomic literature assumes a full (unit) ERPT in the long run. Second, it does not seem totally unreasonable that in small economies the effects of exchange rate movements are, in the long term, passed on to the price level. Third, as we will see, having estimated long run coefficients on the exchange rate in the levels equation (9) in most cases we are unable to reject the hypothesis that they are equal to 1 (see Table 2), although admittedly the confidence intervals are relatively wide, and thus in quite some cases we can neither reject them being equal to zero.

Despite the possible suspicion of breaks within our sample, we decide not to use break detection methods, as the sample length is relatively short. Therefore we proceed with: 1) estimating the "long run" ERPT coefficients, 2) testing for cointegration in the estimated and imposed long run relationships, both using single equation ADF tests and Pedroni (1999) panel test, 3) estimating the short run, first difference relationship.

Finally, in Appendix B we provide some insight on ERPT in aggregate CPI, aggregate IUVs and experiment with using the longer 1995-2005 sample.

Long run ERPT estimates

Sector	Czech R.	Estonia	Hungary	Latvia	Lithuania	Poland	Slovakia	Slovenia
SITC_0	1.03	1.72	2.05	-0.16	1.91	0.74	-1.04	0.62
	(.23)	(0.6)	(.32)	(.32)	(.32)	(.28)	(.67)	(.36)
SITC_1	-	3.19	1.88	-	-	-	-	-
		(1.8)	(.61)					
SITC_2	0.28	1.64	2.13	-1.95	0.85	0.34	0.65	1
	(.26)	(0.5)	(.67)	(.72)	(.32)	(.26)	(.22)	(0.2)
SITC_3	0.75	0.77	0.79	1.22	0.8	1.25	0.6	0.37
	(.31)	(.32)	(.58)	(0.6)	(.15)	(.17)	(.11)	(.18)
SITC_4	0.03	0.58	2.06	1.14	-0.19	2.02	-0.14	0.56
	(.24)	(.93)	(.25)	(.13)	(.27)	(.14)	(0.5)	(.19)
SITC_5	-0.07	0.47	1.38	-0.04	1.32	0.92	2.02	0.6
	(.48)	(1.2)	(.24)	(0.4)	(.39)	(.13)	(.44)	(0.2)
SITC_6	0.43	0.63	1.84	0.18	1.04	1.28	0.84	-0.11
	(0.4)	(.49)	(0.2)	(.34)	(.32)	(0.1)	(0.5)	(.23)
SITC_7	0.16	3.97	0.96	2.2	0.43	0.79	-0.38	0.06
	(.51)	(1.3)	(.25)	(.41)	(.61)	(.41)	(.91)	(.22)
SITC_8	-0.68	1.02	1.63	1.58	0.78	0.66	1.02	0.95
	(.71)	(.72)	(.26)	(.23)	(.41)	(.15)	(.48)	(.36)

For each country/industry combination the long run ERPT estimate from equation (9) is reported. Standard errors in brackets.

Table 2: Long run ERPT estimates for individual countries and sectors.

H_0: Unit root (no cointegration)

Sector	Czech R.	Estonia	Hungary	Latvia	Lithuania	Poland	Slovakia	Slovenia
SITC_0	-6.35***	-3.62***	-5.31***	-4.38***	-4.53***	-5.11***	-3.97***	-5.3***
SITC_1		-4.4***	-3.96***					
SITC_2	-4.29***	-5.66***	-4.95***	-3.59***	-3.43***	-5.37***	-3.9***	-4.34***
SITC_3	-3.27***	-4.23***	-3.17***	-4.24***	-6.1***	-4.95***	-5.03***	-5.77***
SITC_4	-2.06**	-5.28***	-2.71**	-4.39***	-2.37**	-4.07***	-4.75***	-2.4**
SITC_5	-4.66***	-5.25***	-4.74***	-6.98***	-5.08***	-2.6**	-6.03***	-5.01***
SITC_6	-4.79***	-4.32***	-5.58***	-5.3***	-3.14***	-4.16***	-4.88***	-7.54***
SITC_7	-4.18***	-2.66***	-2.54**	-3.88***	-3.92***	-2.19**	-5.63***	-5.17***
SITC_8	-4.23***	-2.13**	-3.17***	-2.25**	-5.64***	-6.02***	-4.58***	-5.85***

*For each sector the ADF t-statistic reported. *,**,*** indicate H_0 can be rejected at 10%, 5% and 1%.*
Specification: no constant, no trend. Maximum lags number = 6. Downward t selection.

Table 3: ADF tests on the errors from the OLS regression of the "long run" single equations (9). Sample: 1999m1-2005m12.

H_0: Unit root (no cointegration)

Sector	Czech R.	Estonia	Hungary	Latvia	Lithuania	Poland	Slovakia	Slovenia
SITC_0	-1.54	0.54	-2.23	-4.54***	-4.11***	-1.21	-2.7*	-4.57***
SITC_1	-	-1.95	-3.84***	-	-	-	-	-
SITC_2	-1.49	-6.01***	-4***	-3.26**	-3.2**	-5.17***	-2.23	-4.43***
SITC_3	-2.81*	-4.05***	-2.81*	-1.65	-3.04**	-3.94***	-1.57	-3.6***
SITC_4	-2.35	-2.11	-1.9	-4.77***	-1.74	-2.04	-4.6***	-3.32***
SITC_5	-3.88***	-2.99**	-3.24**	-1.74	-0.9	-1.14	-4.4***	-2.56*
SITC_6	-1.39	-3.27**	-5.09***	-4.57***	-0.16	-3.48***	-4.79***	-5.03***
SITC_7	-4.23***	-2.34	-2.48	-2.09	-3.41***	-1.36	-5.12***	-1.21
SITC_8	-2.33	-2.38	-3.21**	-2.23	-4.73***	-3.43***	-5.45***	-7.37***

*For each sector the ADF t-statistic reported. *,**,*** indicate H_0 can be rejected at 10%, 5% and 1%.*
Specification: constant, no trend. Maximum lags number = 6. Downward t selection.

Table 4: ADF tests on the errors from the imposed "full pass-through long run" single equations (9). Sample: 1999m1-2005m12.

Model	pseudo-t
Estimated long run	-15.82
Imposed long run	-5.79
Under the null hypothesis the statistics have a N(0,1) distribution	

Table 5: Pseudo-t statistics from Pedroni (1999). The null hypothesis is no cointegration. Sample: 1999m1-2005m1, full panel (N=66), no cross-unit dependence.

IV.6 Results

Generally, we find rather strong evidence of the existence of a cointegrating relationship between our variables (see Tables 3, 4 and 5). With the data span that we deal with, it is hard to claim to have a good overview of the actual long-run developments, moreover the sheer long-run estimates do not tend to be very precise. Both of these issues however, support our proposal to use both the 'estimated long-run' and 'imposed long-run' specifications, in to account for the possibility of both a full and an imperfect pass-through in the long run.

In general the sectoraly disaggregated short run estimates are not very precise. In many cases of the single equation short-run estimates, though point-estimates fall within the expected values, more than 50% cannot reject the short run being equal to 0 nor it being equal to 1, which may be due to the exacerbation of the (dis)aggregation problems of dealing with IUVs, especially in short series. In order to increase the explanatory power, we decided to propose grouping sectors within a country in order to perform a panel estimation. Obviously, this is a rough way to go about the problem, but it seems the only one available. Therefore we proceed by estimating three types of common coefficient, fixed effect panels for each country: (i) TOTAL - one which groups all industries in a country in the cross-section dimension; (ii) SITC_0-4, also referred to as primary sector, which groups

[13]Several empirical papers (see De Bandt, Banerjee and Kozluk, 2006; Campa and Minguez, 2006), using different definitions of long-run ERPT into import prices find incomplete pass-through even in the long run, in a number of sectors/countries.

Short run estimates with "estimated long run"

	Czech R.	Estonia	Hungary	Latvia	Lithuania	Poland	Slovakia	Slovenia
				Estimated long run				
TOTAL	0.28^{***}	0.9^{+}	1.08^{+++}	0.18^{***}	0.74^{+++}	0.8^{+++}	0.55	0.53^{+++**}
	(.21)	(.56)	(.29)	(.30)	(.26)	(.14)	(.52)	(.18)
SITC_0-4	0.79^{+++}	0.4	1.14^{+++}	0.12^{**}	0.81^{+++}	1.08^{+++}	0.66	0.61^{+++*}
	(.21)	(.59)	(.43)	(.45)	(.30)	(.17)	(.38)	(.22)
SITC_5-8	-0.53^{***}	2.52^{++}	1.05^{+++}	0.27^{*}	0.55	0.54^{++**}	-0.16	0.45^{+**}
	(.43)	(1.2)	(.32)	(.39)	(.44)	(.22)	(1.5)	(.27)
				Imposed long run				
TOTAL	0.35^{***}	0.99^{+}	0.78^{+++}	0.25^{**}	0.88^{+++}	0.81^{+++}	0.23^{*}	0.64^{+++**}
	(.23)	(.57)	(.29)	(.32)	(.26)	(.15)	(.45)	(.18)
SITC 0-4	0.66^{+++}	0.55	0.72^{+}	0.45	1.03^{+++}	0.99^{+++}	0.47	0.66^{+++}
	(.23)	(.60)	(.43)	(.47)	(.31)	(.18)	(.41)	(.22)
SITC 5-8	-0.16^{***}	2.46^{++}	0.92^{+++}	0.02^{**}	0.48	0.63^{+++}	-0.5	0.71^{++}
	(.44)	(1.24)	(.33)	(.42)	(.41)	(.23)	(1.47)	(.30)

For each country/industry combination rows show: (1)short run ERPT estimate from equation (10), (2) standard deviation.
+++,++,+ indicate estimate is different from 0 at 99%, 95%, 90% respectively.
****,**,* indicate estimate is different from 1 at 99%, 95%, 90% respectively.*

Table 6: Short run ERPT estimates from equation (10). SITC_0-4: 0 - Food and live animals chiefly for food, 1 - Beverages and Tobacco, 2 - Crude materials, inedible, except fuels, 3 - Mineral fuels, lubricants and related materials, 4 - Animal and vegetable oils, fats and waxes; SITC_5-8: 5 - Chemicals and related products, n.e.s., 6 - Manufactured goods classified chiefly by materials, 7 - Machines, transport equipment, 8 - Manufactured goods n.e.c.; , TOTAL: SITC_0-4 and SITC_5-8.

sectors of commodity, agricultural goods and foodstuffs; (iii) SITC_5-8 also referred to as the processed or manufactured goods sector, which encompasses sectors of manufactured and chemical goods. In order to justify our grouping approach we must notice that retaining separate estimates for each country is in line the country specific characteristics such as monetary policy regimes and their changes, which have been found to matter for ERPT. More importantly, much of the work in the area of ERPT tended to show specific patterns that distinguish the behavior of prices of commodity/primary sectors from manufactured sectors (see for example Campa and Gonzalez-Minguez, 2006; De Bandt, Banerjee and Kozluk, 2006). Finally, the shares of CEECs trade done with the EU tend to be higher in the manufactured goods than in commodity sectors, while they remain similar between countries. This is yet another reason to justify not only the grouping strategy adopted, but also the fact that we allow different ERPT estimates in primary and manufactured sectors. Thus we obtain common coefficient estimates which will be some sort of averages between sectors and between sector groups in a country. The results are advertised in Table 6.

First, the results do not differ significantly regardless whether the estimated or imposed long run are used.[14] Next, the only country where the results are problematic is Slovakia, where all the estimated rates are both not significantly different from 0 nor from 1. In case of 4 countries (Estonia, Hungary, Lithuania and Poland) the point estimates of the total ERPT into imports price is close to 1 and in fact insignificantly different from it. Of course, can we not be certain this is evidence of full pass-through, but the estimates are reasonably close. Next, in case of Slovenia, the Czech Republic and Latvia the full pass-through is rejected, and only in the first case, we cannot reject the one-period pass-through being

[14]Admittedly, the fact that we cannot reject the hypothesis of long-run ERPT smaller than one in Czech Republic and Slovenia (see Table 2 could lead us to favor the 'estimated long-run' specification in the case of the two countries. However, overall conclusions do not differ vastly, regardless of which of the two specifications we adopt.

zero. Switching to the SITC subgroups (ii) and (iii), we find that in case of the "commodity and primary sectors" in Poland, Lithuania, Hungary and the Czech Republic, we cannot reject full ERPT in the short term, while easily rejecting the zero pass-through. In all of these countries, aside the Czech Republic, the point estimates of ERPT in the commodity sectors are rather close to one, indicating full or close to full ERPT. On the other hand, in Slovenia the commodity ERPT is somewhat smaller, while still much higher than 0. The estimates for Estonia and Latvia, do not provide much insight. As for the production sectors, only in Hungary and Estonia the pass-through can be interpreted as full or close to full, while in Poland and Slovenia it is clearly non-zero, but rather not full. Finally, the short-run ERPT in production sectors is rather low and close to zero in the Czech Republic and Latvia, while not much can be said about Lithuania.

Turning to the insight on the robustness of the results, Figures 11 to 13 show that reducing the sample by dropping up to 21 either last or first observations, does not, in principle, change the estimates or their significance. This confirms that during the analyzed period (1999-2005) ERPT developments in the CEECs where rather stable, and thus our results seem to be robust.

Finally Appendix B presents the results for ERPT into CPI and aggregate import prices. Though these results must be treated with even more caution than the above, we can see that pass-through into aggregate IUVs in Hungary, Lithuania and Poland seems generally significant and rather high in the short run, and close to full or full in the long run. As for the second degree pass-through - i.e. into the aggregate price indices, there is some evidence of non-zero short-term ERPT and still rather high, albeit not full aside Lithuania, ERPT in the long run.

IV.7 Discussion

Overall, we can say that exchange rate regimes are not the primary driving force behind the pass-through of the nominal effective exchange rate to import prices, as evidence pointing towards full or close to full immediate pass-through into IUVs in the total and commodity sectors can be found both for Poland and Hungary, with rather flexible regimes as well as for Lithuania which has had a currency board regime. Similarly, there is some weak indication of full ERPT in Estonia.[15] On the other hand, Slovenia with a managed float has rather intermediate values of ERPT in the aggregate estimates and in both sector groups, while the Czech Republic has low aggregate, basically zero ERPT in production sectors, and rather high commodity sector short run pass-through rates. Generally, only in the case of Slovenia and Latvia can we reject full ERPT in the commodity sectors, and only in Poland, Czech Republic, and Latvia can we reject full ERPT in the production sectors.

Altogether, although our results are rough, they point to the issue of potential vulnerabilities in future strategies to adopt the euro. By the EU aquis, all the CEECs are required to join the common currency, and the ERM II waiting room imposes bounds on both exchange rate movements (for two years) and price changes (practically for 1 year). In countries like Poland and Hungary, where the full and immediate pass-through is coupled with a rather flexible exchange rate regime, the central bank will have to combine both the focus on price stability and on the effect of instantaneous movements in the exchange rate versus the euro and the euro/dollar rate on domestic prices. This may prove both tricky and hard to justify by the statutes of the central banks. At this point in time, this seems more of a problem for Hungary, which both historically and currently has higher inflation,

[15]The lack of precision of the results for Estonia may be driven by the fact of the kroon being tied to the currency of its largest trading partner (the EU) throughout the entire sample and thus experiencing very little nominal effective exchange rate variability.

and if this were to continue, would need a smaller price swing (and thus, as a consequence of larger ERPT into the CPI, a smaller change in the exchange rate) to breach the benchmark. Admittedly by the time the two countries decide to join the euro, the situation may turn around, but the fact that not only there seems to be a strong immediate effect on import prices, but also an immediate effect on CPI backed by a strong or even full pass-through effect in the long run indicates this issue should not be overlooked. With lower and slower pass-through the Czech Republic seems much less prone to the same problems.

On the other hand, Lithuania's currency board arrangement insulates it from fluctuations of the lita against the euro, and makes the fulfillment of the exchange rate criterion seem rather simpler. However, with a full and immediate pass-through from the effective nominal exchange rate its prices are exposed to euro/dollar exchange rate changes, which are beyond any control of the central bank. Its attempt to enter the Eurozone in 2007 was set back because of inflation marginally exceeding the entry requirement, which proved the strict approach of the EU to applying the criteria. The fact that Lithuania exceeded the benchmark by less than 0.1 percentage points means that the difference between qualifying and not managing to is subtle and therefore an unfavorable euro/dollar development with full and instantaneous pass-through into import prices could, if it were to feed in to the CPI, jeopardize its euro prospects. Backing this argument we find a tendency for exchange rate fluctuations to be passed on to the CPI, to some extent in the short run and close to fully in the long run. Therefore, the currency board, will not provide full insulation against exchange rate changes. As for Estonia, this may be the case, but the evidence is rather weak, while in Latvia, more room for exchange rate fluctuations is accompanied by a rather low pass-through. Finally Slovenia with its imperfect pass-through is already on a steady road to the common currency.

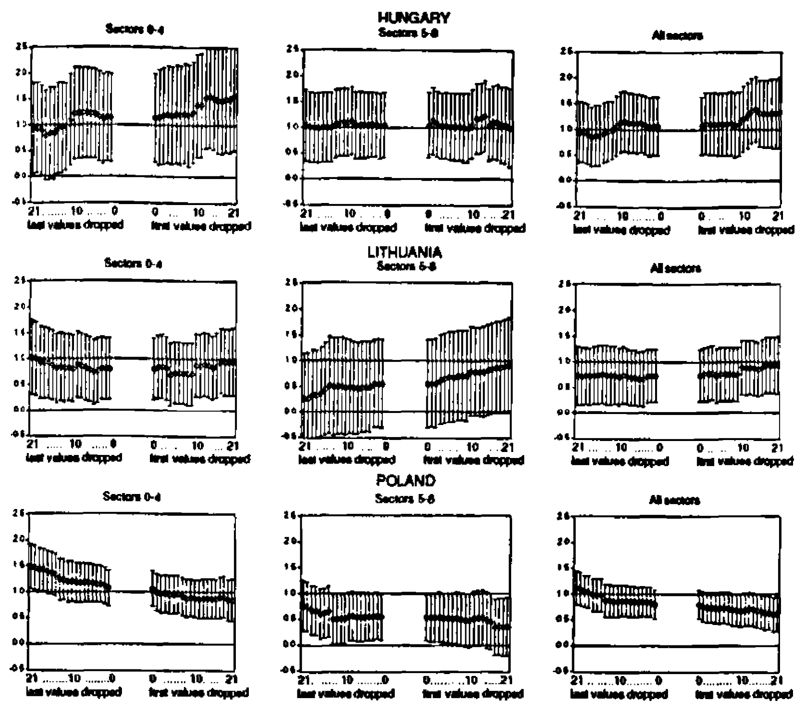

Figure 11: Recursive estimations of short-run ERPT into IUVs. Point estimates and 95% confidence intervals. In each panel, the observations are obtained as follows: (left part of the panel) we start by dropping the 21 last observations and estimate equation (10) and plot the short-run ERPT, next we add one observation and reestimate. So on until reaching the full sample. Then (right part of each panel) we take the full sample and drop the first observation, and plot the resulting ERPT, and so on until having dropped the first 21 observations. The specification is 'estimated long-run'. Solid horizontal lines are at values 0 and 1.

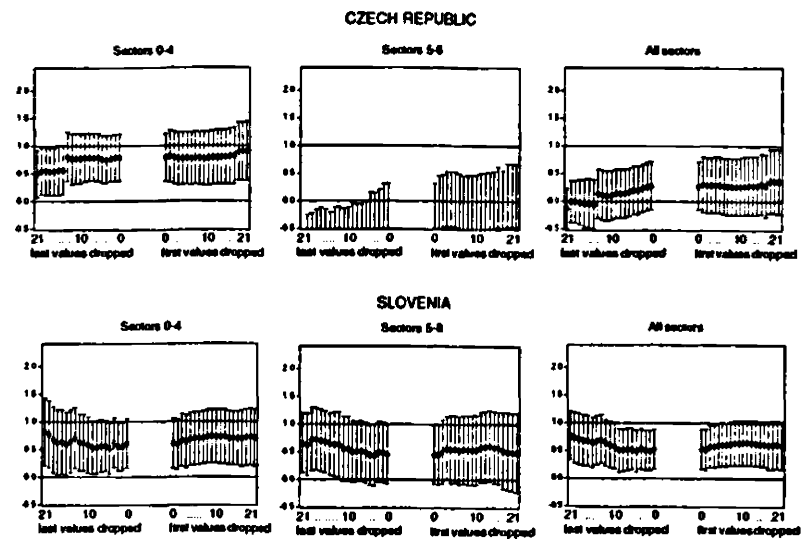

Figure 12: Recursive estimations of short-run ERPT into IUVs. Point estimates and 95% confidence intervals. See Fig. 11 for details.

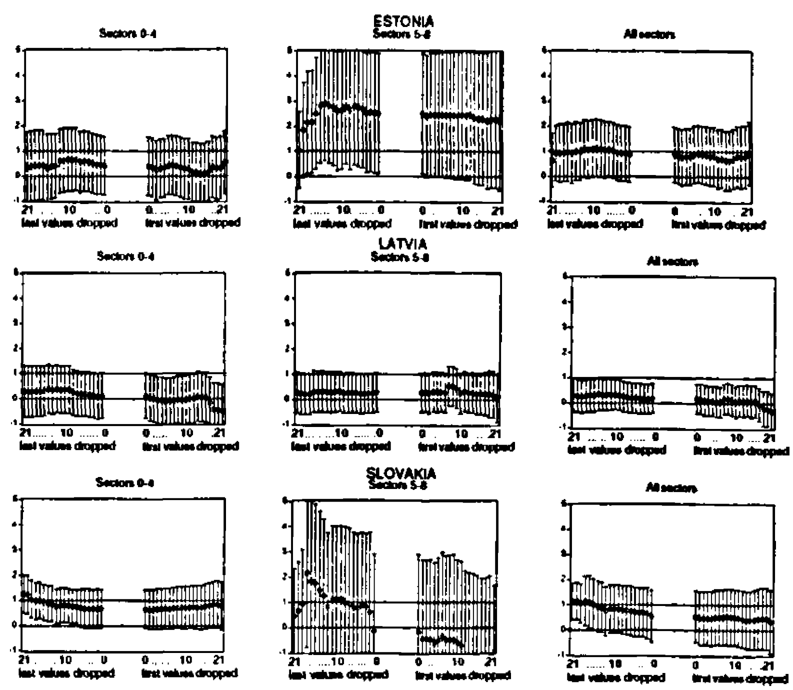

Figure 13: Recursive estimations of short-run ERPT into IUVs. Point estimates and 95% confidence intervals. See Fig. 11 for details.

IV.8 Conclusions

Exchange rate pass-through has been been given significant attention, due to its insight on foreign and domestic firm market power, market competition structure and the design of monetary policy. In this paper we attempt to estimate the ERPT into sectoral import prices in 8 Central European countries. The main problems regard the data - the short transition history is further troubled with structural change issues of major shifts in exchange rate regimes, ongoing integration with the EU, economic transition itself and a change in regime in the regions largest trading partners - the introduction of the euro. Despite this, we manage to obtain several interesting results.

Admittedly, the relatively short time span and consequently imprecise estimates do not allow us to say too much about the long run ERPT in the new Member States. However, we can say a number of things about the estimates of the short run pass-through. First of all they are rather robust to the different long run specifications proposed - i.e. they are unaffected by whether we use the estimated values of the long run or impose an assumption about full pass-through in the long run. Moreover, despite rather disappointing power of the sectoral estimates, estimating the values for the groups of commodity sectors and production sectors yields relatively strong results. Short run pass-through of the effective exchange rate fluctuations in the two most advanced (in terms of GDP per capita) Central European countries, Czech Republic and Slovenia, which suggests foreign exporters are less willing to pass-through short term fluctuations of the exchange rate thus more pricing to market. Notably the two countries differ to a large extent on the degree of exchange rate flexibility - Slovenia has a rather tightly managed float, while the Czech koruna is rather flexible. On the other hand in Poland, Hungary and Lithuania ERPT is full or close to full in the short term, meaning that any effective exchange rate fluctuations are passed on immediately to the import prices. Again the exchange rate regime does not seem a crucial determinant here - Poland with practically a free float is behaving similarly as Hungary with a very wide fluctuation band and Lithuania with a currency board. Finally, except for Slovenia, and less confidently Estonia ERPT in commodities is full or close to full and is generally higher than in the production sectors, where it tends to take intermediate values. As most of trade in commodities is done with extra-euro partners, this indicates that even a currency board with the euro may fail to provide insulation of commodity prices against the fluctuations of exchange rates.

Although tentative, the finding that 'first-degree' ERPT tends to be full or close to full, tends to be an important indicator for the future strategies regarding the waiting room for the euro. Full and immediate pass-through into import prices, if largely fed through further to CPI, which as we show seems to be the case, could potentially jeopardize euro entry prospects by causing problems with fulfilling the inflation criterion. Up-to-date experience has shown, that the assessment tends to be strict, and together with the fact that even a currency board does not insulate against global exchange rate fluctuations, a tendency to pass on these changes into the price may prove sufficient for the benchmark to be exceeded. Of course this will surely not be the only obstacle, and quite possibly not a major one, however it is a vulnerability that should be kept in mind.

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Appendix A - Data Sources: Eurostat, COMEXT.

import prices - monthly indexes of effective import unit values (calculated to be based on local currency, combines imports from inside and from outside EU-25). Logs taken. Rebased for 2000=100.

foreign prices - monthly indexes of import unit values into the EU-25 (calculated to be based on the effective euro/dollar aggregate currency, combines imports from inside and from outside EU-25). Logs taken. Rebased for 2000=100.

exchange rates - index of monthly average exchange rate of local currency against the effective euro/dollar aggregate. Logs taken. Rebased for 2000=100.

SITC code - Industry

0 - Food and live animals chiefly for food
1 - Beverages and Tobacco
2 - Crude materials, inedible, except fuels
3 - Mineral fuels, lubricants and related materials
4 - Animal and vegetable oils, fats and waxes
5 - Chemicals and related products, n.e.s.
6 - Manufactured goods classified chiefly by materials
7 - Machines, transport equipment
8 - Manufactured goods n.e.c.

Appendix B - Total ERPT estimates

Long run ERPT estimates

Country	CPI 1995 (1)	IUV 1999 (2)	CPI 1999 (3)	CPI 1995 (4)	IUV 1999 (5)	CPI 1999 (6)
CZK	0.54	-0.2	0.1	0.52	0.39	0.2
	(.03)	(.35)	(.05)	(.03)	(0.4)	(.05)
EST	1.1	2.58	0.53	0.8	1.93	0.38
	(.12)	(.76)	(.15)	(.09)	(.59)	(.12)
HUN	0.58	1.79	0.33	0.58	1.07	0.38
	(.02)	(.18)	(.04)	(.02)	(.16)	(.04)
LAT	-0.08	-0.69	0.04	0.51	-1.22	0.51
	(0.1)	(.33)	(.11)	(.18)	(.61)	(0.2)
LIT	0.81	0.41	0.44	1.23	0.67	0.56
	(.28)	(.36)	(.24)	(.22)	(.39)	(.23)
POL	0.65	0.75	-0.18	0.71	1.02	-0.16
	(.06)	(0.1)	(.04)	(.05)	(0.1)	(.05)
SLK	0.19	1.64	0.16	0.15	0.66	0.16
	(.03)	(.41)	(.09)	(.02)	(.22)	(.05)
SLN	0.42	0.79	0.53	0.41	0.43	0.48
	(.02)	(.26)	(.04)	(.02)	(.19)	(.03)

For each country the long run ERPT estimate from equation (9) is reported. Standard errors in brackets.

Table 7: Long run ERPT estimates for individual countries in the following specifications: columns (1)-(3) euro denominated imports defined as from EU-25; (4)-(6) euro denominated imports defined as EU15;

H_0: Unit root (no cointegration)

Country	CPI (1)	CPI (2)	IUV (3)	IUV (4)	CPI (5)	CPI (6)
CZK	-3.79***	-1.76	-4.44***	-2.54	-3.07***	-1.43
EST	-5.49***	-3.88***	-3.96***	-3.47***	-2.38**	-3.28**
HUN	-3.01***	-1.43	-4.52***	-2.92**	-2.37**	-1.45
LAT	-4.57***	-2.11	-6.11***	-2.3	-2.83***	-0.79
LIT	-4.92***	-3.08**	-5.14***	-3.88***	-5.32***	-2.5
POL	-2.73***	-2.3	-5.11***	-2.38	-2.49**	-1.82
SLK	-2.36**	-1.24	-5.81***	-2.99**	-2.86***	-1.33
SLN	-4.64***	-1.38	-7.28***	-7***	-2.33**	-0.72

For each country the ADF t-statistic reported.

,,*** indicate H_0 can be rejected at 10%, 5% and 1%.*

Specification: constant: yes (2),(4),(6), no (1),(3),(5), no trend.

Maximum lags number = 6. Downward t selection.

Table 8: ADF tests on the errors from the "long run" single equations (9). Columns: (1) - ERPT into total CPI, estimated long run, 1995m1-2005m12; (2) - ERPT into total CPI, imposed long run, 1995m1-2005m12; (3) - ERPT into total IUV, estimated long run, 1999m1-2005m12; (4) - ERPT into total IUV, imposed long run, 1999m1-2005m12; (5) - ERPT into total CPI, estimated long run, 1999m1-2005m12; (6) - ERPT into total CPI, imposed long run, 1999m1-2005m12. Euro imports defined as EU25.

H_0: Unit root (no cointegration)

Country	CPI (1)	CPI (2)	IUV (3)	IUV (4)	CPI (5)	CPI (6)
CZK	-3.63***	-1.84	-3.87***	-2.77*	-1.64*	-1.04
EST	-5.77***	-3.39***	-5.26***	-3.95***	-8.52***	-2.52
HUN	-2.97***	-0.9	-5.12***	-2.89**	-2.51**	-0.97
LAT	-3.68***	-3.05**	-6.33***	-3.51***	-6.77***	-3.12**
LIT	-4.98***	-2.8*	-5.80***	-0.71	-6.28***	-1.89
POL	-3.01***	-2.6*	-6.13***	-2.86**	-3.3***	-2.13
SLK	-3.14***	-0.36	-5.68***	-0.45	-3.1***	-0.52
SLN	-4.54***	-1.95	-6.81***	-6.12***	-3.46***	-1.82

For each country the ADF t-statistic reported.
,,*** indicate H_0 can be rejected at 10%, 5% and 1%.*
Specification: constant: yes (2),(4),(6), no (1),(3),(5), no trend.
Maximum lags number = 6. Downward t selection.

Table 9: ADF tests on the errors from the "long run" single equations (9). Columns: (1) - ERPT into total CPI, estimated long run, 1995m1-2005m12; (2) - ERPT into total CPI, imposed long run, 1995m1-2005m12; (3) - ERPT into total IUV, estimated long run, 1999m1-2005m12; (4) - ERPT into total IUV, imposed long run, 1999m1-2005m12; (5) - ERPT into total CPI, estimated long run, 1999m1-2005m12; (6) - ERPT into total CPI, imposed long run, 1999m1-2005m12. Euro imports defined as EU-15.

Short run estimates of total ERPT

	EU-25						EU-15					
	CPI 1995		IUV 1999		CPI 1999		CPI 1995		IUV 1999		CPI 1999	
Country	(1)	(2)	(3)	(4)	(5)	(6)	(1)	(2)	(3)	(4)	(5)	(6)
CZK	0	0	0.58	0.35	-0.01	-0.05	0	0.01	0.21	0.13	-0.01	-0.01
	(.03)	(.03)	(.5)	(.56)	(.01)	(.01)	(.03)	(.03)	(.52)	(.55)	(.01)	(.03)
EST	0.09	0.1	-1.01	-0.81	0.15	0.2	0.09	0.08	-0.11	-0.58	0.11	0.18
	(.12)	(.12)	(1.3)	(1.3)	(.23)	(.21)	(.12)	(.12)	(1.3)	(1.3)	(.22)	(.23)
HUN	0.2^{+++}	0.16^{+++}	1.36^{+++}	0.85^{+}	0.09^{++}	0.06^{+}	0.2^{+++}	0.17^{+++}	0.88^{++}	0.67	0.08^{++}	0.07^{+}
	(.01)	(.01)	(.41)	(.47)	(.01)	(.01)	(.01)	(.01)	(.38)	(.43)	(.01)	(.01)
LAT	0	-0.01	1.01	0.81	-0.25	-0.29	0.11	0.14	2.16^{++}	1.99^{++}	-0.03	-0.02
	(.13)	(.14)	(.71)	(.75)	(0.2)	(0.2)	(.15)	(.16)	(.89)	(.93)	(.21)	(.25)
LIT	0.26	0.33	0.85	0.65	0.21	0.21	0.62^{++}	0.77^{++}	1.37^{+}	1.72^{++}	0.5	0.67
	(.29)	(.31)	(.69)	(.74)	(.39)	(.41)	(.31)	(.32)	(.76)	(.87)	(.37)	(.43)
POL	0.06^{+}	0.06^{++}	0.63^{++}	0.73^{++}	0.01	0.01	0.07^{+}	0.05^{+}	0.56^{++}	0.6^{++}	0.05	0.01
	(.01)	(.03)	(.27)	(0.3)	(.05)	(.05)	(.01)	(.03)	(.25)	(.28)	(.01)	(.01)
SLK	0	-0.01	0.19	-0.58	0	-0.02	0.01	0.01	0.72	-0.26	0.03	0.02
	(.05)	(.05)	(.95)	(.09)	(.08)	(.08)	(.05)	(.05)	(.81)	(.91)	(.07)	(.08)
SLN	0.05	0.12^{+++}	0.8	1.18	0.07	0.18^{++}	0.05	0.12^{+++}	0.72	0.6	0.07	0.16^{+}
	(.01)	(.01)	(.88)	(.91)	(.09)	(.09)	(.01)	(.01)	(.81)	(.84)	(.08)	(.09)

For each country short run ERPT estimate from equation (10) reported, standard deviation in brackets.
+++, ++, + indicate estimate is different from 0 at 99%, 95%, 90% respectively.

Table 10: Total short run ERPT estimates (into CPI and IUV, different samples: 1995m1-2005m12 and 1999m1-2005m12) from equation (10). In the ERPT into CPI, coefficient b_1 is reported.