Towards the estimation of equilibrium exchange rates for CEE acceding countries: methodological issues and a panel cointegration perspective[*]

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Abstract: *This paper provides a discussion of methodological issues relating to the estimation of the long-run relationship between exchange rates and fundamentals for Central and Eastern European acceding countries, focusing on the so-called behavioural equilibrium exchange rate (BEER) approach. Given the limited availability and reliability of data as well as the rapid structural change acceding countries have been undergoing in the transition phase, this paper identifies several pitfalls in following the most straightforward and standard econometric procedures. As an alternative, it looks at the merits of a two-step strategy that consists of estimating the relationship between exchange rates and economic fundamentals in a panel cointegration setting – using a sample which excludes acceding countries – and then "extrapolating" the estimated relationships to the latter. While focusing on the first step of such a strategy, the paper also delves into discussing technical aspects underlying the "extrapolation" stage. As a result, the paper endows the reader with the methodological and empirical ingredients for computing equilibrium exchange rates for acceding countries, providing estimates for the long-run coefficients between real exchange rates and economic fundamentals and a discussion of how to apply these results to acceding countries data.*

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1. Introduction

With the completion of the negotiations on EU accession, ten countries will join the European Union in May 2004 as members with a derogation. These countries will eventually adopt the euro, upon reaching a high level of sustainable convergence as spelled out in the Maastricht treaty. Directly after EU accession, each member state will treat exchange rate policy as a "matter of common interest" and, in principle, will have the option to participate in the ERM II. For doing so, it is necessary to reach a multilateral agreement on a sustainable central parity for the exchange rate of new members' currencies vis-à-vis the euro. Assessing exchange rate developments in relation to their "fair level" is conceptually difficult, as one needs to rely on a wide set of indicators and a broad range of models. In this context, various concepts of "equilibrium exchange rates" may serve as a helpful tool to address the issue of the appropriate level for a currency. Estimating such equilibrium exchange rates is a challenging task for major currency pairs and is even more complicated for the exchange rates of acceding countries, on account of substantial problems of data availability and measurement as well as difficulties in choosing an appropriate econometric methodology. Moreover, as these countries were in transition within the short time span for which data are available, their equilibrium exchange rate may have undergone rapid change in itself, further complicating the assessment.

Indeed, the sizeable and in some cases massive real appreciation of acceding countries' currencies may reflect to a large extent this transition to more market-oriented economies. Since 1993, the Baltic countries faced the strongest real appreciation. The external value of the Estonian and Latvian currencies more than doubled in real terms (CPI-based) and in the extreme case of Lithuania the rise in the real exchange rate amounted to more than 600%.[1] In these countries, the pace of the real appreciation was very high up to 1997 and moderated thereafter. In comparison to the Baltic States, the increase in the real external value of the currencies of the Czech Republic, Poland and Slovakia since 1993 was smaller but still substantial. In Hungary, the real exchange rate of the forint appreciated steadily between 1995 and 2002, apart from a period of depreciation during the Russian crisis. By contrast, the real appreciation of the Slovenian tolar was more moderate.

This paper focuses on a so-called behavioural equilibrium exchange rate (BEER) approach and presents an empirical analysis of the long-run relationship between exchange rates and fundamentals based on panel cointegration techniques. On account of limitations of standard econometric models when applied to acceding countries' currencies, this paper looks at the merits of a two-step approach. The first step involves the estimation of long-run relationships between the real exchange rate and economic fundamentals, using data from non-acceding OECD countries ("estimation stage"). The second step involves the derivation of an "equilibrium real exchange rate" as a function of these economic fundamentals using data from acceding countries ("extrapolation" stage).[2] This paper focuses on the first step and

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[1]   Overall, the change in the real exchange rate on the basis of producer prices was lower but in most countries still significant.

[2]   Unlike other contexts, in the context of this paper, the term "extrapolation" should not be intended as an out-of-sample exercise along the temporal dimension, but as an application of the estimated coefficients out of sample across the cross-sectional dimension.

discusses the extrapolation stage only at a methodological level. Accordingly, the paper should be understood as a kind of a guide, providing the ingredients and the procedural steps to compute "equilibrium" exchange rates for acceding countries. In this context, we provide estimates for the long-run equilibrium coefficients between economic fundamentals and real exchange rates and a discussion of how to apply these results to acceding countries data.

The paper is structured as follows: The next section briefly discusses some fundamental methodological issues with a special emphasis on the BEER approach and provides a selected overview of the academic literature on this topic. Section 3 presents a simple cross-section analysis, which focuses on the relationship between productivity advances and exchange rate developments. Section 4 extends this analysis by applying panel cointegration techniques. Moreover, it broadens the perspective by introducing additional macroeconomic fundamentals, which could be relevant for the determination of equilibrium exchange rates. Section 5 discusses technical aspects underlying the second stage (so-called "out-of-sample" approach). Section 6 concludes.

2. Methodological issues and review of the literature

The literature on acceding countries' exchange rates has employed different concepts of equilibrium exchange rates, from the fundamental equilibrium exchange rate (FEER) approach (see, e.g. Šmídková *et al.* (2002) as well as Coudert and Couharde (2002)) to the the monetary model (see Crespo-Cuaresma *et al.* (2003)) to the BEER approach. The focus of this paper is on the BEER, which constitutes a more empirical approach to modelling the relationship between real exchange rates and various economic fundamentals.[3] Overall, the available empirical BEER studies on acceding countries' equilibrium exchange rates can be classified into three categories (as illustrated in Chart 1): a) studies which estimate equilibrium exchange rates on a country-by-country basis,[4] b) cross-section analyses and c) studies based on panel data, which exploit simultaneously the cross-section and the time series information contained in the data. The merits and limitations of these approaches are briefly addressed below.
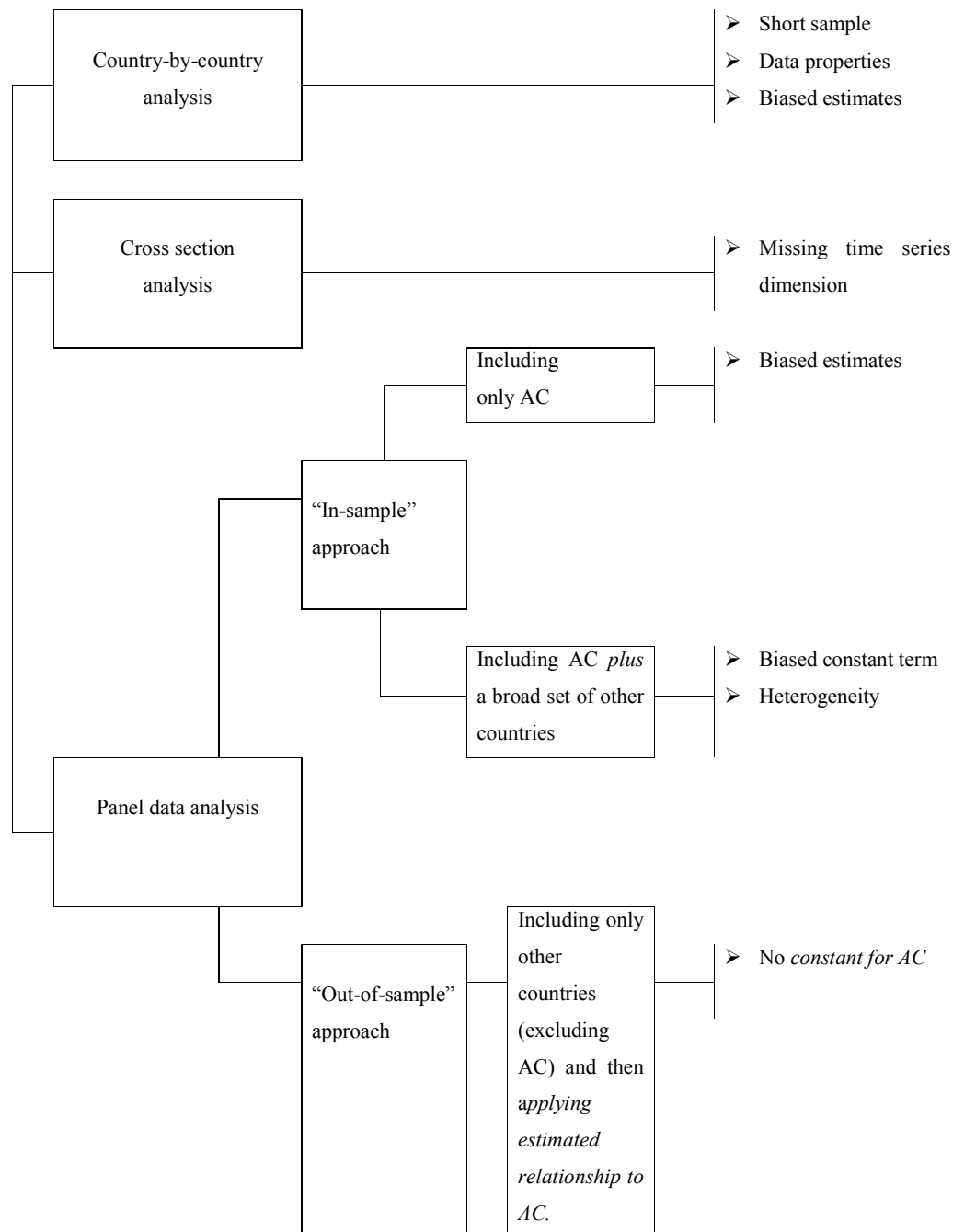
2.1 The country-by-country approach

Carrying out *country-by-country analyses* is intuitively the most straightforward strategy, as the equilibrium exchange rate is estimated for each country, taking into account the peculiarities of each individual economy. However, this approach could be subject to drawbacks in terms of interpretation of the results, due to data problems and the particularity of the transition period.

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[3] Different concepts for assessing exchange rates are discussed in ECB (2002) and in MacDonald (2000). Obviously, the BEER is only one of the methods that could be employed. Using the FEER or a NATREX model could be considered as an alternative, but a thorough assessment on the basis of alternative methodologies is beyond the scope of this paper.

[4] For instance, Jakab and Kovacs (2000) estimate the relationship between economic fundamentals and the Hungarian forint in a SVAR framework. Egert (2002) estimates the relationship between economic fundamentals and the real exchange rate of individual CEE countries, but also shows results of a panel test. Similarly Rahn (2003) estimates both country-by-country and panel regressions.

Chart 1: Methodological issues in choosing an appropriate econometric strategy in a BEER framework

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┌──────────────────────┐                                    ➢  Short sample
│  Country-by-country   │                                    ➢  Data properties
│      analysis         │                                    ➢  Biased estimates
└──────────────────────┘

┌──────────────────────┐                                    ➢  Missing time series
│    Cross section      │                                       dimension
│      analysis         │
└──────────────────────┘
                          ┌──────────────┐                  ➢  Biased estimates
                          │  Including    │
                          │  only AC      │
                          └──────────────┘
                     ┌──────────────┐
                     │ "In-sample"  │
                     │  approach    │
                     └──────────────┘
                          ┌──────────────────┐              ➢  Biased constant term
                          │ Including AC plus │              ➢  Heterogeneity
                          │ a broad set of    │
                          │ other countries   │
                          └──────────────────┘
┌──────────────────────┐
│  Panel data analysis  │
└──────────────────────┘
                          ┌──────────────────┐              ➢  No constant for AC
                          │ Including only    │
                     ┌────────────┐ other     │
                     │"Out-of-    │ countries  │
                     │ sample"    │ (excluding │
                     │ approach   │ AC) and then│
                     └────────────┘ applying   │
                          │ estimated         │
                          │ relationship to   │
                          │ AC.               │
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First of all, the length of the sample in these studies is rather short, commonly spanning a period of around 10 years as it would be futile to include data reaching back to the 1980s when these countries were still operating in a planned-economy environment. Since the real exchange rate and the underlying macroeconomic fundamentals are commonly found to be non-stationary variables, cointegration analysis is often employed. However, cointegration

test statistics and estimates are strongly biased in such small samples. The fluctuations of real exchange rates can exhibit rather long-lasting swings, so that it could well be that even a sample period of ten years fails to reflect broad movements around some equilibrium schedule. Importantly, some studies have suggested that acceding countries' exchange rates have converged gradually from an initial substantial undervaluation towards their equilibrium level in the 1990s.[5] Accordingly, the actual exchange rates might have gradually approached the equilibrium level after their strong devaluation at the beginning of the reform process and it cannot even be ruled out *per se* that they may have not reached this equilibrium value yet. If this possibility is not properly accounted for, some of the ensuing coefficient estimates (in particular, the estimate of the intercept) could be biased.

Chart 2 provides a very stylised illustration of this point. It shows a typical pattern for acceding countries' real exchange rates over the last twelve years (bold solid line).[6] At the early stage of transition, these currencies appreciated strongly, to increasingly level off in recent years. If the correct path of the equilibrium exchange rate is given by the thin black solid line, the exchange rate was initially undervalued but converged over time to more "reasonable" levels and has been close to its equilibrium or even slightly overvalued in recent years.[7] The equilibrium exchange rate is assumed to gradually appreciate, for example owing to higher productivity gains in acceding countries compared with the euro area. If one simply estimates a cointegration model linking the real exchange rate to productivity on a country-by-country basis, without taking the initial period of undervaluation into account, one would expect the constant term to be biased (dotted line). This would give rise to misleading results for the evaluation of the exchange rate level.

In order to account for a gradual adjustment of the initially undervalued currency towards equilibrium, Krajnyak and Zettelmeier (1998), e.g., explicitly address the issue of adjustment to equilibrium from an undervalued position by including in their panel data analysis a time-varying "transition dummy". However, if the coefficient on the transition dummy is the same for all countries, it reflects only the average extent of "undervaluation" of these currencies, although the adjustment path may have been rather diverse. As an alternative, one could consider including a "transition variable", which could be shaped for example like a logarithmic trend, on a country-by-country basis. Experimenting with such an approach, however, we found that over-fitting the equilibrium schedule results, which amounts to finding by construction that the currencies are close to their equilibrium at the end of the sample, independently of how much they have appreciated in the past.

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[5]   See Halpern and Wyplosz (1997) as well as Krajnyak and Zettelmeier (1998). As this analysis is built on real dollar wages, it is difficult to draw strong conclusions for the euro exchange rates of CEE currencies. It also needs to be emphasised that between 1995-2002, the real appreciation of the currencies of CEE countries was stronger against the euro than against the US dollar. However, these studies raise the awareness that such an eventuality needs to be taken into account in choosing an appropriate econometric strategy. Darvas (2001), by contrast, suggests that Hungary did not start the transition period with a substantially undervalued currency.

[6]   The series is constructed as an unweighted average of accession countries' real exchange rates between 1993 and 2004. Forecasts provided by the AMECO database were used for the end of the sample.

[7]   The series was constructed using the trend in per capita income of accession countries relative to the euro area, multiplied by an elasticity of 0.5. The level is set *ad hoc* to illustrate a situation of an exchange rate which is broadly in line with fundamentals in recent years.

Chart 2: Illustration: Estimating an appreciating equilibrium exchange rate

Index; number of years on the horizontal axis
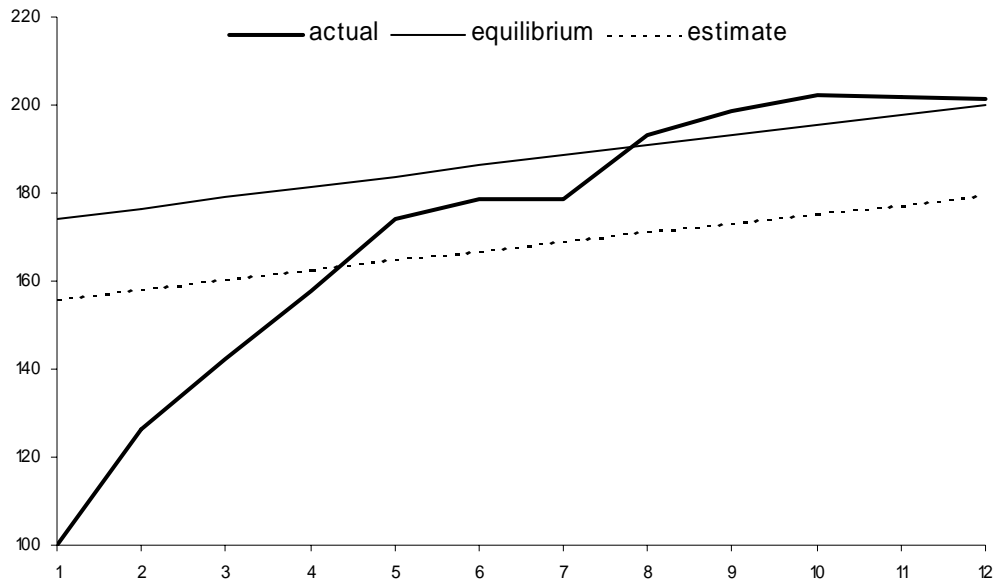
2.2 The cross-section approach

Instead of employing a time series approach, one could employ a ***cross-section regression***, to avoid the problems outlined above regarding the data properties. This strategy is often followed when analysing equilibrium exchange rates of acceding countries in a more illustrative setting, mainly based on graphical analysis. Such an approach follows, in principle, Kravis et al. (1982) and use US-dollar based purchasing power parity conversion factors provided by the International Comparison Programme of the University of Pennsylvania or data provided by Eurostat.[8] These papers commonly explain, in a cross-section context, the gap between this PPP exchange rate and the actual exchange rate (the so-called "exchange rate gap") on the basis of per capita income (in dollar or euro PPP terms). An analogous approach for the euro is employed in section 3 of this paper.

2.3 The panel data approach

More sophisticated studies have employed panel data models in order to overcome the problem of short time series.[9] These papers can be separated into studies following either an "in-sample" or an "out-of-sample" approach (see Chart 1). The ***"in-sample"*** approach focuses

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[8] This approach has been employed by DeBroek and Sloek (2001) in the first part of their paper, Randveer and Rell (2002) in an analysis for Estonia, and Brook and Hargraeves (2001) in an analysis for New Zealand. Cihak and Holub (2003) discuss this approach very comprehensively. They also carry out a similar exercise from a Czech perspective controlling – albeit rather *ad hoc* – for other variables.

[9] See, for instance, DeBroek and Sloek (2001) and Fischer (2002). Rahn (2003) employs a "in-sample" panel cointegration framework for effective exchange rates of acceding countries and derives equilibrium exchange rates on the basis of a algebraic transformation along the lines of Alberola et al. (1999). He finds that acceding countries' currencies may be substantially overvalued.

on the panel including only acceding countries (or a sub-set of these countries) and has the advantage that the cross-section sample is fairly homogeneous. However, such a procedure may entail similar drawbacks as the country-by-country approach, leading to biased estimates of the deviations from equilibrium. Adding more countries which were not subject to the transition process could mitigate the problem of a bias in the estimates, but the constant terms may still be distorted towards finding an overvalued exchange rate.

As an alternative, an *"out-of-sample"* approach is based on a two-step procedure for estimating equilibrium exchange rates for acceding countries. In the first step, the equilibrium exchange rate is estimated for non-acceding countries. In the second step, equilibrium exchange rates for acceding countries are then "extrapolated" on the basis of the estimated structural relationships. In this context, the choice of the cross-section coverage requires a compromise between maximising the degrees of freedom in the estimation by including as many countries as possible and maintaining a reasonable degree of sample homogeneity. Halpern and Wyplosz (1997) and Krajnyak and Zettelmeier (1998) for instance, employed this approach and opted for a very broad country coverage, using a sample of 85 industrialised, developing, planned and transition economies. While a broad panel may be useful in identifying more general patterns, which were less dented by the transition process itself, it also amplifies the problems related to heterogeneity, as countries as diverse as the United States, the Netherlands, Japan, Myanmar, Papua New Guinea and Zimbabwe are combined. Both papers employ static regressions including a broad set of economic fundamentals. However, this does not take the time series properties of the data properly into account, as the data are in all likelihood non-stationary. Using fewer data points in the time dimension (five-year intervals), as done in these studies, does not overcome this limitation.[10]

In this context, both the sample selected and the econometric approach followed by Kim and Korhonen (2002) appear to be more compelling. They employed data from 29 middle and high income countries, arguing that the acceding countries share common characteristics with both groups: while their per capita GDP is more similar to middle-income countries, their industrial and trade structure is more similar to high-income countries. They derive equilibrium exchange rate indices, both against the US dollar and in effective terms, for five acceding countries.

In the "out of sample" approach, the estimation stage would be followed by an extrapolation stage. In this second step, the fitted value for the exchange rate, calculated on the basis of the acceding country's data using the point estimates of the long-run parameters estimated in the previous step, is interpreted as the equilibrium exchange rate for each acceding country. The derivation of an appropriate "equilibrium exchange rate" is, however, far from trivial. Apart from controlling for changes in economic fundamentals affecting the "equilibrium exchange rate" over time, in order to provide a guide about the appropriateness of the current level of

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[10] Halpern and Wyplosz (1997) use for sample covering 1970 to 1990 observations every five years. They find a specification including, as exogenous variables, GDP per worker, school enrolment, the share of agriculture relative to industry, the share of the government sector, and the inflation rate, which, in their original paper, are statistically significant according to standard critical values. As this study estimates a model based on a panel data set excluding acceding countries for the period 1975 to 1990 and applies estimates to acceding countries for the period after 1990, this can be interpreted as a "double" out-of-sample exercise. In an update of their paper (Begg et al. (1999)), after including one more data point for 1995 for each country, some of these fundamentals become insignificant while others, such as demographic factors, trade openness and external indebtedness, become important. On the whole, this may point to some stability problems in this approach.

the exchange rate, one needs to choose an appropriate constant term too. The latter is estimated with a bias in "in-sample" regressions and cannot be estimated directly in "out-of-sample" exercises. Section 5 discusses some methodological issues involved in choosing a constant term in such exercises.

3. A cross-section analysis

This section presents an assessment of the relationship between real exchange rates and fundamentals, based on a cross-section approach using data on PPP exchange rates against the euro as provided by the European Commission's Ameco database (See Annex 1 for the data sources). The PPP exchange rate is defined as the number of units of a country's currency required to purchase the same basket of goods in that country as one unit of the numeraire currency would buy in the numeraire country. Accordingly, PPP exchange rates are commonly used as conversion factors, which allow cross-country comparisons of economic aggregates in real terms. For instance, for the Czech koruna against the euro the PPP conversion factor is defined as the number of koruna needed in the Czech Republic to buy the same basket of goods and services as one euro would buy in the euro area. In 2002, the PPP exchange rate of the Czech koruna was CZK/EUR 15.18, while the actual euro-koruna exchange rate was CZK/EUR 30.8. This implies that the real value of the Czech currency is almost 50% lower in the euro area than in the Czech Republic. This mainly results from the fact that prices of non-traded goods and services are much higher in the euro area than in the Czech Republic (while traded goods should have more similar prices in both countries).

On the basis of this concept one can compute so-called "exchange rate gaps", defined as the ratio of the PPP exchange rate and the actual exchange rate. In a time series context, they have the same pattern as real exchange rate indices, but PPP exchange rates have the advantage of being meaningful in levels. Chart 3 shows these gaps for 38 industrialised countries, emerging markets and the accession countries in 2002. For the CEE acceding countries, the gaps are strongly negative, ranging from –0.33 for Slovenia to more than –0.60 in the case of Slovakia. In this sample, only Bulgaria and Romania have larger exchange rate gaps. For Malta and Cyprus, exchange rate gaps are smaller than for some euro area countries, such as Portugal or Greece, and comparable with Korea and Spain. At the other end of the spectrum are countries such as Japan and Switzerland, where the cost of living is much higher than in the euro area. Overall, this indicates that the purchasing power of CEE acceding countries' currencies is much lower in the euro area than in their home countries.

The presence of higher non-traded goods and services prices in one country relative to another country – given traded good prices – is often related to the Balassa-Samuelson theory, according to which relative prices of non-traded and traded goods in each country are inversely related to the relative productivity in the two sectors.[11] In a nutshell, since traded goods sector productivity can be assumed to be higher in industrialised countries, non-traded

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[11] In practice, the differentiation between traded and non-traded goods is a artificial simplification. While the general pattern that prices of non-traded goods are lower than prices of traded goods seems to prevail, in Central and Eastern European countries, for instance, the price of semi-durable goods and food is also sizeably lower than in the EU. Accordingly, the real appreciation experienced by acceding countries' currencies in recent years may also reflect price increases in the tradable sector. We are grateful to the anonymous referee for pointing this out.

goods prices are also higher there. In addition, the relative price between non-traded and traded goods is influenced by demand-side factors as well as price regulation and tax policies. Demand effects may arise, for instance, from non-homothetic preferences when the demand for non-traded services, which are more likely to have luxury good characteristics, increases with prosperity (see Bergstrand (1991)). Regarding the possible effect of government inference, Cihak and Holub (2003) cannot find a statistically significant relationship between the exchange rate gap and several fiscal variables in a similar cross-section analysis. By contrast, they find the size of agricultural employment as being significant. This is interpreted as a proxy for the political temptation to government inference in this sector, which may in turn have an impact on food prices.

Chart 3: Exchange rate gap in 2002

Source: ECB and AMECO
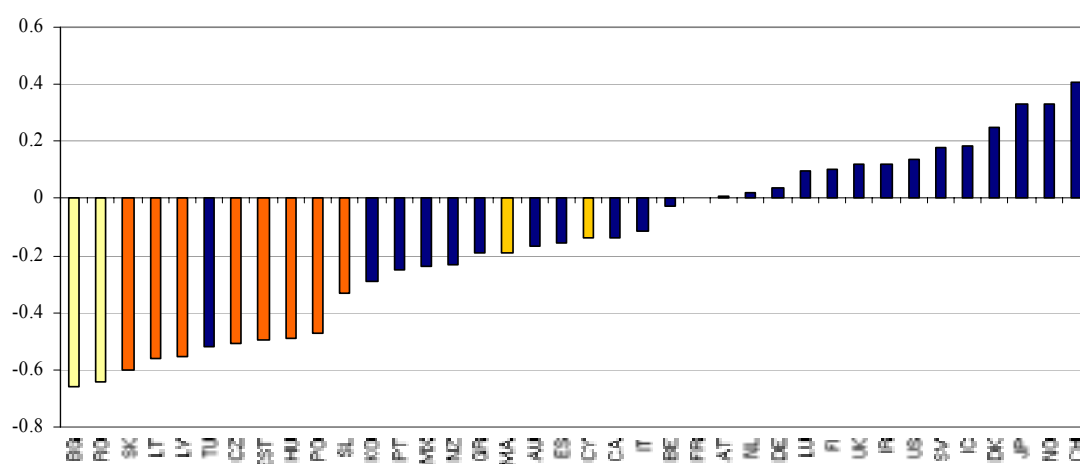
Overall, the "exchange rate gap" (*egap* henceforth) should be farther in negative territory, the lower the level of productivity and the stage of a country's economic development. To account for the Balassa-Samuelson and demand effects, two explanatory variables, which should show a positive link with the exchange rate gap, have been constructed. The first is economic development, measured as GDP per capita in PPP terms (*ypc*)[13] and the second is a productivity variable, measured as GDP in PPP terms divided by the number of persons

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[13]   This follows seminal work by Kravis et al. (1982) and it is what Samuelson (1996) called the "Penn effect". Using per-capita income may also better reflect the fact that the relative price between non-traded and traded goods is also influenced by demand-side factors.

employed (*prod*).[14] Average labour productivity (ALP) and per capita income are highly correlated by construction and cannot be used in the same regression to disentangle the relative importance of supply and demand side factors. In order to avoid the influence of the transition trend present in acceding countries' data on the estimation of the equilibrium, the regression was run by excluding the acceding countries.[15] The OLS cross-section estimation results for 2002 are summarised below (t-statistics in parentheses).

$$\ln(egap) = \underset{(-7.78)}{-1.56} + \underset{(7.62)}{0.50} \ln(ypc) \qquad N = 25, \; R^2 = 0.65 \tag{1.1}$$

$$\ln(egap) = \underset{(-4.62)}{-1.93} + \underset{(4.57)}{0.48} \ln(prod) \quad N = 24, \; R^2 = 0.36 \tag{1.2}$$

There is a significant positive relationship between the exchange rate gap and per-capita income as well as productivity. Overall, the results using per-capita income are stronger and more robust than those based on ALP as the t-statistics and the goodness-of-fit (measured by the adjusted R^2) are markedly higher.[16] The results of this exercise depend to a significant extent on the group of countries included. While the elasticities presented in DeBroek and Sloek (2001) are broadly in line with these results, Coudert and Couharde (2003), for instance, find a lower goodness-of-fit and elasticity (0.25) in a similar regression based on income per capita including 120 developing and emerging countries. Experimenting with a broader country sample from the World Developments Indicators, we found that the inclusion of poor countries – particularly African countries – tends to generate lower elasticies. Nonetheless, it could be argued that the integration of these countries in international financial and goods markets is fairly limited so that it might be more sensible to include only countries with a reasonable degree of global integration.

While such a cross-section analysis provides a simple indication of the relationship between the exchange rate and income/productivity across countries, it has the drawback that it ignores the time series information contained in the data. In order to incorporate this information, one needs to adopt a panel data modelling approach. The analysis of the time series characteristics

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[14] On purely theoretical grounds, total factor productivity (TFP), which signals improvement in the overall efficiency of the economic process, is the favoured measure of productivity. From a practical point of view, however, TFP cannot be measured directly and is difficult to estimate, particularly for accession countries (see Dobrinsky 2001, p. 13 for a critical discussion and computation of TFP for acceding countries). As regards inputs in ALP, limited data availability for many countries rules out the use of output per hour worked. Therefore, output per person employed is commonly used in this literature. In order to account for diverging trends in productivity in the traded and in the non-traded goods sectors, several studies (see Fischer (2002), DeBroek and Sloek (2001)) have employed productivity measures in different sectors of the economy, e.g. agriculture, industry, and services. These data, however, are not consistently available for a broad cross-section of countries, so that we essentially assume that productivity advances materialise primarily in the traded goods sector and focus on economy-wide productivity measures.

[15] The following 25 countries have been included: Australia, Austria, Belgium, Canada, Germany, Denmark, Finland, France, Greece, Ireland, Iceland, Italy, Japan, Korea, Mexico, Netherlands, Norway, New Zealand, Portugal, Spain, Sweden, Switzerland, Turkey, United Kingdom, United States. Since productivity data for Mexico was not readily available from the AMECO database, the second regression comprises only 24 countries.

[16] Including the acceding countries in the regression increases the estimated elasticities significantly. In the cross-section regression excluding the acceding countries an increase of per-capita income relative to the euro area of 1% translates into an appreciation of almost 0.5%. In the regression including acceding countries the elasticity rises to almost 0.7%. For productivity, the corresponding elasticity is somewhat lower if the acceding countries are included and higher if they are excluded.

of the data indicates the presence of unit roots. As a consequence, the bivariate relationship between per-capita income and the exchange rate gap is estimated in a panel cointegration framework. Subsequently, we discuss the inclusion of other macroeconomic fundamentals in the model.

4. Taking a panel cointegration perspective

4.1 Econometric strategy

The panel data perspective adds the time series dimension to the cross-section analysis of the previous section. While this increases the information set, it also complicates the analysis, as the time series characteristics have to be taken properly into account. It has been widely recognised that the real exchange rates as well as the underlying fundamentals are mostly non-stationary variables, which must be modelled in a suitable econometric framework in order to avoid drawing conclusions based on spurious results. Accordingly, we first test for unit roots in order to confirm that the variables are indeed integrated. We then test for cointegration and estimate the long-run parameters.

4.1.1 Unit root tests

Testing for unit roots in panel data instead of individual time series entails the advantage of increasing the power of the test by exploiting simultaneously cross-section and time series information. Furthermore, the test statistics conveniently converge asymptotically to the standard normal distribution. We carried out panel unit root tests on the basis of two standard test procedures: the Im et al. (2003) test (IPS t-test), which tests the null hypothesis of a unit root, and the test proposed by Hadri (2000), which has stationarity as the null hypothesis.

The structure of the IPS t-test is based on N augmented Dickey-Fuller regressions:

$$\Delta y_{it} = \rho_i y_{it-1} + \sum_{j=1}^{p_i} \varphi_{ij} \Delta y_{it-j} + \alpha_i + \gamma_i t + \varepsilon_{it} \qquad \text{for t} = 1\dots \text{T}; \quad \text{i} = 1, \dots \text{N}, \quad (1.3)$$

where T is the length of the sample, N is the cross-section dimension, y_{it} is the variable under consideration, the term following the sum includes lagged dependent variables with country-specific lag length p_i, α_i and γ_i are country-specific intercepts (fixed effects) and trend parameters, respectively. The error term ε_{it} is distributed as a white-noise random variable, with possibly different variance for each member of the panel. The crucial coefficient for testing for a unit root is ρ_i. In contrast to the test proposed earlier by Levin et al. (2002), the IPS t-test allows for heterogeneity in the value of ρ_i under the alternative hypothesis of stationarity. This implies that the mean convergence dynamics are allowed to vary across the cross-section members. Hence the null hypothesis is H$_0$: $\rho_i = 0$ for all i against the alternative hypothesis H$_a$: $\rho_i < 0$. The IPS t-bar statistic is defined as the average of the individual ADF statistics.[17] After a suitable normalisation of the statistics using simulated values tabulated in Im et al. (2003), which account for the number of lags, the distribution of the test is standard normal.

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[17] Alternatively, Im et al. (2003) also propose an asymptotically equivalent LM-bar statistic. The results in this paper are, however, independent of the use of the t-bar or the LM-bar statistic so that the LM-bar is not discussed further.

The test proposed by Hadri (2000) is a residual-based Lagrange Multiplier test (LM) which – in the spirit of the KPSS test suggested by Kwiatkowski et al. (1992) – has a reverse null hypothesis, i.e. that the time series for each cross-section unit is stationary around a deterministic level or trend, against the alternative hypothesis of a unit root. It is based on the following regression:

$$y_{it} = \alpha_i + \gamma_i t + \sum_{t=1}^{T} u_{it} + \varepsilon_{it}$$

(1.4)

where the deterministic terms are defined as in (1.3) above, and the error term has two components: ε_{it}, which is white noise, and $\sum_{t=1}^{T} u_{it}$, which is a random walk. The test is based on the fact that under the null hypothesis of stationarity the variance of the random walk component (σ_u^2) is zero. The test statistic takes the form $\dfrac{\sigma_u^2}{\sigma_\varepsilon^2}$, which has a standard normal distribution under the null hypothesis.

4.1.2 Cointegration tests

The cointegration tests proposed by Pedroni (1999) have become a standard workhorse in panel data econometrics.[18] He proposes seven residual-based tests based on the null hypothesis of no cointegration. The starting point is the group-by-group estimation of the proposed long-run relationship:

$$y_{it} = \alpha_i + \gamma_i t + \theta_t + \beta_1 x_{1it} + ... + \beta_K x_{Kit} + \varepsilon_{it}$$

(1.5)

where K is the number of regressors and β_k are the elasticities. The deterministic elements (α_i and γ_i) are defined as above and θ_t are common time effects. This formulation allows for considerable heterogeneity in the panel since fixed effects, individual-specific deterministic trends and different error variances are all allowed. Some of the Pedroni tests (the so-called group tests) also allow for heterogeneous slope coefficients, as the elasticities are estimated by averaging the individual β_k instead of pooling the long-run information. There is no requirement for exogeneity of the regressors since the dynamics are determined jointly for both y_i and all x_{ki}.

The seven tests proposed by Pedroni follow asymptotically a standard normal distribution after a suitable normalisation. A particularly important correction takes into account the heterogeneity of the cross-section units, by adjusting for the group-specific long-run variance of cointegration residuals. Four of the tests are based on pooling along the "within"-dimension of the panel and three are based on averaging along the "between"-dimension.

Based on Monte Carlo experiments for a case with one dependent variable (Pedroni (1997)), the tests seem to have distorted size and low power for sample sizes below T=100 (see also Banerjee et al. (2003)). Overall, Pedroni (1997) suggests that the panel-ρ statistic seems to be the most reliable when T is large enough; for small T, the parametric group-t statistic and panel-t statistic appear to have the highest power, followed by the panel-ρ statistic. In view of

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[18] As an alternative, the tests by Kao (1999) and McCoskey and Kao (1998) could be employed. However, the tests proposed by Pedroni allow for heterogeneous variances across the countries in the panel and some form of dependence across the countries at each point in time.

this, we carried out some simulations to derive approximate small-sample critical values for the tests (see Annex 1).

4.1.3 Estimation and inference in dynamic heterogeneous panel models

We consider three main approaches to estimating a long-run relationship among integrated variables in a panel framework: (1) the error correction (EC) mean-group and pooled mean-group estimators proposed by Pesaran et al. (1999), (2) the fully-modified OLS (FMOLS) and (3) the dynamic OLS (DOLS) estimators. The latter two estimators have been associated with Pedroni (2000), Mark and Sul (2001) and Kao and Chiang (2000). The basic starting assumptions in all cases are that the variables are I(1), that they are cointegrated with the real exchange rate, but not among themselves, and there is only one cointegration vector.

In the present context, the issue of slope homogeneity is particularly relevant. The strategy of estimating the relationships on the basis of a sample (N_1) that excludes acceding countries and then extrapolating to the sample of acceding countries (N_2), requires slope homogeneity between the two samples. Since this cannot be tested, a weaker requirement is to test whether slope homogeneity holds at least for the countries in the N_1 sample. This can be done by computing both pooled and mean-group estimators. The pooled estimator estimates the long-run parameters jointly, thereby maximising the degrees of freedom. The mean-group estimator, by contrast, implies estimating the parameters country-by-country and then averaging them across countries. It provides consistent estimates of the mean of the long-run slope coefficients (though it suffers from a lagged dependent variable bias for small T), but it is inefficient if slopes are homogeneous. Under the null hypothesis of homogeneity the pooled estimator is consistent and efficient, while it is inconsistent under the alternative hypothesis. Using the fact that the mean-group estimator is always consistent, a Hausman test can be constructed to test for slope homogeneity.[19]

An important assumption of all panel cointegration tests considered in this paper is the absence of cross-sectional correlation. The possible presence of contemporaneous correlation can be addressed either by using time dummies or by subtracting the cross-sectional mean from the data. If the slopes are homogeneous, the two methods are equivalent. In the present paper, demeaned data were used for the pooled mean group estimator and time dummies were included for the other estimators.

Dynamic estimation: the error correction model.

Pesaran, Shin and Smith (1999) propose a pooled-mean group estimator (PMGE), which constrains the long-run coefficients to be identical in an error correction framework, but allow

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[19] Pesaran et al. (1999) also provide a likelihood ratio test for equality of error variances and/or slopes and claim that these tests usually reject the null of homogeneity. Therefore they suggest the Hausman test as an alternative. Pesaran et al. (1996) present two different ways to compute a Hausman test depending on how the variance of the mean group coefficient is obtained: either using a parametric or a non-parametric estimator. For the case of dynamic models, they find that the latter has better size properties than the former. However, using the non-parametric variance estimator does not guarantee a positive definite value for the difference between the variance of the mean group coefficients and the variance of the pooled ones. We are not aware of Monte Carlo studies about the properties of the different version of the Hausman test when other methods than the error correction are used.

the short-run coefficients and error variances to differ across groups. They propose estimating the following autoregressive distributed lag (ARDL) model of order (p_i, q_i):

$$\Delta y_{it} = \phi y_{it-1} + \beta_i' x_{it} + \sum\nolimits_{j=1}^{p_i-1} \lambda_{ij} \Delta y_{it-j} + \sum\nolimits_{j=0}^{q_i-1} \delta_{ij}' \Delta x_{it-j} + \alpha_i + \gamma_i t + \varepsilon_{it} \qquad (1.6)$$

where y_{it} is the dependent variable, x_{it} is a *m x 1* vector of explanatory variables, α_i and γ_i represent the country-specific intercepts and time trend parameters respectively, λ_{ij} and δ_{ij} include the country-specific coefficients of the short-term dynamics, ε_{it} is a white noise error term. The long-run coefficients β are defined to be the same across countries. If ϕ_i is significantly negative, there exists a long-run relationship between y_{it} and x_{it}. The equation is then estimated using the maximum likelihood procedure to get the PMG estimator. This regression can also be estimated with individual specific β_i which are then averaged over N to obtain a mean-group estimator (MGE) which is the natural background to test for the presence of slope homogeneity based on a Hausman test.

Static estimation: FMOLS and DOLS

The starting equation to be estimated with these methods is the static regression (1.5). In a country-by-country set-up, this corresponds to the Engle-Granger procedure, which generates a consistent estimator of the long-run parameters. However, in the panel set-up, the long-run parameters are biased and Kao et al. (1999) show, on the basis of Monte Carlo experiments, that correcting the coefficients for this bias does not improve over the uncorrected OLS. This leads to using alternative methods, such as the FMOLS and the DOLS.

- ### *FMOLS in panel data.*

The FMOLS takes into account the presence of the constant term and the possible correlation between the error term and the differences of the regressors. To adjust for these factors, non-parametric adjustments are made to the dependent variable and then to the estimated long-run parameters obtained from regressing the adjusted dependent variable on the regressors. Accordingly, the FMOLS long-run coefficient estimators are defined as:

$$\hat{\beta}_i = \left(\sum\nolimits_{t=1}^{T} x_{it}' x_{it} \right)^{-1} \sum\nolimits_{t=1}^{T} \left(x_{it}' y_{it}^* - T \hat{\lambda}_i \right) \qquad (1.7)$$

where y_{it}^* are the regressands adjusted for the covariance between the error term and the Δx_t and T $\hat{\lambda}$ is the adjustment for the presence of a constant term. The associated statistic for testing the significance of the parameters needs to be similarly adjusted. In the panel setting, the mean-group FMOLS long-run coefficients are obtained by averaging the group estimates over N: $\hat{\beta}_{MG}^{FMOLS} = N^{-1} \sum\nolimits_{i=1}^{N} \hat{\beta}_i$, and the corresponding t-statistic converges asymptotically to a standard normal distribution: $t_{MG}^{FMOLS} = N^{-1/2} \sum\nolimits_{i=1}^{N} t_i \rightarrow N(0,1)$.

The pooled FMOLS coefficients can be computed in two different ways: weighted and unweighted. In the first case each group is weighted by the components of the long-run covariance of the group residuals and the right-hand-side variables in differences; in the second case these components are averaged. Pedroni (2000) shows that the weighted statistics require prior knowledge of the estimated parameters. Therefore, to compute a feasible weighted statistics several authors have proposed different starting values. Pedroni (2000)

uses the values estimated under the null hypothesis, while Kao and Chiang (2000) employ the parameters from a static fixed-effects models. Mark and Sul (2001), for the case of DOLS, suggest to use parameters from an uncorrected DOLS estimation. Other alternatives proposed in the applied part of this paper include using as a preliminary step the estimate of the unweighted version of the tests and the mean group estimate, which is consistent but not efficient.

- ***DOLS in panel data***

The starting point of the DOLS estimator is also equation (1.5). In order to obtain an unbiased estimator of the long-run parameters, DOLS involves a parametric adjustment to the errors of the static regression. The correction is achieved by assuming that there is a relationship between the residuals from the static regression and first differences of the leads, lags and contemporaneous values of the regressors in first differences:

$$\varepsilon_{it} = \sum_{j=-q}^{q} c_{ij}\Delta x_{it-j} + \varepsilon_{it}^{*} \qquad (1.8)$$

Substituting (1.8) into (1.5) yields:

$$y_{it} = \alpha_{i} + \gamma_{i}t + \theta_{t} + \beta_{1i}x_{1it} + \ldots + \beta_{1i}x_{1it} + \sum_{j=-q}^{q} c_{ij}\Delta x_{it-j} + \varepsilon_{it}^{*} \qquad (1.9)$$

A simple OLS regression provides superconsistent estimates of the long-run parameters. The t-statistic is based on the long-run variance of the residuals instead of the contemporaneous variance, which is commonly used in OLS regressions. The group-mean DOLS are obtained in a similar fashion as the group-mean FMOLS. Similarly, weighted and unweighted versions of the DOLS estimator can be derived.

The decision to select one method over the other depends to some extent on the length of the sample (see Pedroni (2000)). In principle, FMOLS requires fewer assumptions and tends to be more robust. However, for the case of a single regressor, Kao and Chiang (2000) conclude that weighted DOLS has a smaller bias than weighted FMOLS. This result depends on the fact that both the number of lags used for kernels (in FMOLS) and the number of lags and leads in DOLS are fixed. Mark and Sul (2001) focus on DOLS and study the small sample performance of weighted and unweighted panel estimators and also of the mean group estimator. Most of the discussion relates to a model with only one regressor, although one specification with two regressors is also included. They suggest that (1) both panel methods outperform the mean group in terms of precision, and (2) the unweighted estimator tends to be more precise and displays smaller size distortion than the weighted estimator. Finally, Pedroni (2000) finds that group mean FMOLS has satisfactory size and power properties even for small panels if T is larger than N.

Since there is currently no paper available studying intensively and conclusively the small-sample properties of these tests based on Monte Carlo experiments, all methods will be employed in section 4.2 to ensure that the results are robust to the method chosen.

4.2 Estimation results

4.2.1 Data and unit root tests

The analysis is carried out on the basis of an annual balanced panel data set spanning the sample period 1975-2002 and includes the countries listed in footnote 15. The exchange rate gap (*egap*) is defined as the purchasing power parity exchange rate divided by the market exchange rate, both in national units per ecu/euro. Per capita income is the GDP per head (*ypcr*) for the total economy in the sample country (in PPP terms) relative to an euro area aggregate (see data Annex for details). Both variables are in logs.

Table 1 summarises the unit root test results. For relative per-capita income, the IPS test fails to reject the null hypothesis of non-stationarity and, correspondingly, the Hadri test strongly rejects stationarity for this variable for specifications with and without trend. Accordingly, this variable is clearly non-stationary. For the exchange rate gap, the tests are also in favour of a unit root process, although the results are not as unequivocal. While the Hadri test again strongly rejects the null of stationarity for the exchange rate gap, the IPS test rejects the null of non-stationarity if there is no deterministic trend included in the regression. On the other hand, when a trend is included in the ADF regression, the IPS test fails to reject the null of a unit root. In the absence of information on whether the data generating process includes a non-zero constant term, one should include a deterministic trend in the ADF regression to ensure asymptotic similarity of the ADF t-statistic with respect to the deterministic coefficients. As a consequence, it seems reasonable to conclude on the basis of this result, and given the strong rejection of the Hadri test, that also the exchange rate gap follows a random walk.

Table 1: Unit root tests

	IPS t-bar-test				Hadri test			
Variable	no trend		Trend		no trend		trend	
Per capita income (ypcr)	-0.29		-2.03		14.25	**	8.22	**
Exchange rate gap (egap)	2.95	**	0.39		14.41	**	8.36	**
**/* Significant at the 1%/5% level. In the IPS-test, two lags have been imposed. The Hadri-test accounts for the presence of autocorrelated errors and employs the finite sample critical values following Hadri and Larsson (2002).								

4.2.2 Panel data results – bivariate regressions

In view of the non-stationarity of the time series, it is important to employ a panel cointegration framework in order to avoid spurious regression problems, which – as shown by Entorf (1997), via simulation, and analytically by Kao (1999) – may lead to highly misleading statistics and potentially invalid conclusions. In the first step, the cointegration tests proposed by Pedroni (1999) are used to verify whether there is a long-run relationship between the exchange rate gap and per capita income. The results are summarised in Table 2.

In the majority of the cases the results point to cointegration between the exchange rate gap and per capita income. The null hypothesis is comfortably rejected in most cases, the exceptions being the non-parametric group-t statistic, which is, however, still fairly close to the 10% level, and the group-ρ statistic, which clearly fails to reject the null. This needs to be

seen against the background of simulation evidence provided by Banerjee et al. (2003) showing that for panels of this dimension all tests suffer from significant size and power distortions, and in particular the group-ρ statistic is grossly undersized. However, the tests performing best according to our Monte Carlo evidence – the panel-v and the panel-ρ statistic – suggest that the variables are cointegrated.

Table 2: Pedroni cointegration tests

Variables: egap, ypcr. N=25, T=28		
	test statistic	p-value
Panel v-Statistic	1.913	0.028
Panel ρ-Statistic	-2.336	0.010
Panel t-Statistic (non-parametric)	-2.310	0.010
Panel t-Statistic (parametric)	-3.333	0.000
Group ρ-Statistic	0.261	0.603
Group t-Statistic (non-parametric)	-1.014	0.155
Group t-Statistic (parametric)	-2.872	0.002

Given the evidence in favour of cointegration among the variables, the panel cointegration method discussed in the previous section can be employed to estimate the long-run parameters. Table 3 summarises the results.[20]

Table 3: Estimation results: (P)MGE, DOLS and FMOLS

	FMOLS		DOLS		(P)MGE	
Dep. var: egap	unweighted	weighted	unweighted	weighted	pooled	mean group
ypcr	0.634 *(28.4)*	0.484 *(33.0)*	0.386 *(20.4)*	0.571 *(29.3)*	0.436 *(4.9)*	0.740 *(3.3)*
φ					-0.313 *(-14.7)*	-0.373 *(-14.8)*
Hausman-test *p-value*	0.85 *(0.357)*	-0.04 *na*	1.21 *(0.271)*	0.00 *(0.991)*	2.20 *(0.14)*	
Note: The PMG was estimated using a ARDL (2,1) specification, for DOLS, 2 lags have been included. t-statistics in parentheses if not indicated otherwise.						

The results broadly confirm those of the cross-section analysis. All estimators suggest that there is a highly significant and positive relationship between the exchange rate gap and per capita income. The coefficients are fairly close to those of the cross-section analysis, suggesting that an increase in per capita income of one percent feeds into a real appreciation – and thus (for lower income countries) a narrowing of the exchange rate gap – of between 0.39 and 0.63%. The significantly negative coefficients of the adjustment term (φ) in the PMGE

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[20] For the lag selection in the EC-PMGE, we first estimated a ARDL(2,2) specification and then removed insignificant lags from the dynamic specification in order to get a more parsimonious model. The DOLS estimates employ two lags and two leads. Overall, the results were very robust with respect to the choice of the lag structure. A country-specific constant has been incorporated while a time trend was not included.

strongly suggests mean reversion of the exchange rate gap to a long-term equilibrium schedule, and, thus, strengthens the evidence of cointegration among the variables. The magnitude of the coefficient implies a half-life of deviations from equilibrium of almost two years.

The results of the Mean-Group estimator (MGE) proposed by Pesaran et al. (1999) are reported in the right-hand-side column. These estimates provide indirect information about parameter heterogeneity in the sample and, thus, poolability of the data. The coefficients are broadly in line with those of the PMG estimator in terms of sign and magnitude. More formally, the Hausman tests cannot reject that the sample is sufficiently homogeneous to be pooled.[21] As can be expected under these conditions, the significance of the coefficients increases noticeably when long-run homogeneity is imposed. The strong evidence in favour of slope homogeneity is quite important in view of the feasibility of the proposed two-step procedure. In-sample homogeneity could be seen as a necessary condition for extrapolation: it does not guarantee that the slopes would also be homogeneous if the sample of countries was enlarged, but indications of heterogeneity would clearly make the extrapolation untenable. Furthermore, even if the long-run parameters were not homogeneous between the transition and non-transition samples in the past, due to transition factors, it could still be appropriate to apply "post-transition" estimates to assess the present level of equilibrium exchange rates in transition countries.

In the country-specific regressions, the long-run elasticities of per capita income are somewhat dispersed, which is partly due to the limited degrees of freedom precluding an efficient estimation of the coefficients. For the MGE, they range from roughly –2 (insignificant) in the case of the United States to +3.2 in the case of Australia. In all but five cases, however, the coefficient has the correct sign and there is not a single case in which the coefficient is significantly negative. The goodness-of-fit (R^2) is commonly below 50% and in all but four cases above 10%. At the 5% significance level, there is evidence for serial correlation in only one out of the 25 countries included in the sample. In three equations there seem to be some problems with the functional form, but there is not a single equation having non-normal errors and only for one country heteroscedastic errors seem to be present. The fact that 21 out of 25 country-specific equations show no evidence of misspecification is reassuring. The results for the FMOLS and DOLS are only slightly worse, as more groups have the wrong sign (seven and nine, respectively), while the dispersion of the parameter estimates across countries diminishes somewhat.

4.2.3 Additional fundamentals

Apart from productivity differentials, an array of variables has been included in model specifications for estimating equilibrium exchange rates.[22] In the following, we pragmatically draw on the findings of this literature in the selection of additional variables.

First of all, given the increasingly free flow of capital across borders, indicators reflecting the *international economic environment* such as the real interest rate differential and terms-of-

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[21] The LR tests reject the restrictions including homogeneity at the 5% significance level, but not at the 1% level, which is rather encouraging since Pesaran et al. (1999) claim that it is almost impossible not to reject this hypothesis.

[22] For a more thorough discussion of variables that can be included in a BEER framework, see MacDonald 2000, this section draws on Osbat and Schnatz (2002).

trade shocks may be relevant. Both variables are, however, subject to measurement problems and data deficiencies, which are even more binding for economies in transition. For instance, data for terms of trade as well as for long-term interest rates are rather scarce. Moreover, a measure of real interest rates would need to include an assumption on inflation expectations, which is difficult to employ, particularly for emerging markets and the acceding countries.[23] Whenever such variables were employed in empirical work related to acceding countries, the findings also were not very encouraging. For example, DeBroek and Sloek (2001) and Fischer (2002) include the terms of trade and various indices of commodity prices but generally fail to find a robust link with the real exchange rate. Moreover, commodity price shocks may affect countries differently, contingent on their resource dependency, making a pooled estimation of the terms-of-trade effects more difficult.

Regarding the domestic economic environment, variables related to the *fiscal stance and monetary policy considerations* are frequently included as determinants of the real exchange rate. Government spending relative to GDP is one variable typically included in such regressions. In practice, government spending may have different short-run and long-run effects on real exchange rate movements. In the short to medium run, a positive impact of government spending on the real exchange rate can be conceived, if the private sector lowers its demand for goods less than the increase in government spending, or if the marginal propensity of the public sector to spend on non-traded goods is higher. The latter possibility appears plausible, given the government sector's spending on infrastructure, for instance, which is mainly satisfied by domestic inputs. Therefore, higher government spending could affect the real exchange rate positively via higher demand for non-traded goods. In the long term, however, higher government spending, in particular if deemed unbalanced, may lead to distortions in an economy and undermine the market's confidence in a currency. On the other hand, particularly in acceding countries in recent years, public spending affected, to a sizeable extent, growth-enhancing effects such as human capital formation, infrastructure building and establishing an institutional environment supporting a proper functioning of the market economy. In the available empirical estimates, such growth-enhancing effects, in combination with the demand-side considerations, appear to dominate.[24] On the financial sector side, two variables have been employed in the literature: money balances and credit growth. Given their ambiguous impact (as they may reflect expansive monetary shocks as well as financial deepening), neither DeBroek and Sloek (2001) nor Begg, Halpern and Wyplosz (1999) find a significant effect of such variables.

Several authors also suggest that an increase in *economic openness* should have an influence on the real exchange rate. Similar to the government spending variable, there are arguments in favour of a positive or a negative impact on the real exchange rate. On the one hand, the more open to trade a country is, the less it relies on protection and distortions to its external accounts: hence increasing openness should enhance the country's economic performance and lead to an appreciation of the real exchange rate. To some extent, this channel may already be captured by including a proxy for productivity developments. On the other hand, Edwards (1994) and Elbadawi (1994) derive an equilibrium exchange rate model for developing countries where it is shown that a greater openness to foreign trade leads to a real depreciation

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[23]   However, omitting the real long-term interest differential might not be a crucial issue as, theoretically, this should be stationary, and thus, it should have a non-permanent impact on the BEER.

[24]   See Kim and Korhonen (2002), Halpern and Wyplosz (1997), Fischer (2002) as well as Jazbec (2002).

as a result of lower tariffs on imports or taxes on exports. The empirical evidence for this variable is broadly in favour of a significant negative impact on the real exchange rate.[25]

Against this background, two variables were added (in logs) to the model that links the exchange rate gap to per capita income: government spending as a ratio of GDP relative to the EU average (*gov*), and openness defined as the average of exports and imports of goods as a ratio of GDP (*open*), again relative to the EU average. The estimation procedure is the same as in the bivariate case.

4.2.4 Panel data results – multivariate regressions

Tests for the order of integration of government spending and openness clearly suggest that both are random walks (see Table 4).

Table 4: Unit root tests

	IPS-t-bar-test				Hadri-test			
Variable	no trend		trend		no trend		trend	
Gov. spending to GDP (gov)	-0.35		0.62		14.73	**	7.61	**
Openness (open)	0.44		-0.05		12.77	**	5.34	**
**/* Significant at the 1%/5% level. In the IPS-test, two lags have been imposed. The Hadri-test accounts for the presence of autocorrelated errors and employs the finite sample critical values following Hadri and Larsson (2002)								

Pedroni cointegration tests, on balance, also confirm that there is cointegration among the variables (see Table 5). While the Group-ρ statistic again fails to reject the null of no cointegration, the previous section argued that this test is subject to excessive size and power distortions for panels of this size, which make it unreliable. By contrast, the group-t statistics and the panel-t statistics strongly reject the null hypothesis of no cointegration. As a result, it seems reasonable to proceed under the assumption that the variables are cointegrated.

Table 5: Pedroni cointegration tests

egap, ypcr, gov, open	test statistic	p-value
Panel v Statistic	0.450	0.326
Panel ρ Statistic	-0.575	0.282
Panel t Statistic (non-parametric)	-2.837	0.002
Panel t Statistic (parametric)	-2.433	0.007
Group ρ Statistic	1.023	0.847
Group t Statistic (non-parametric)	-2.355	0.009
Group t Statistic (parametric)	-3.344	0.000

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[25] Kim and Korhonen (2002) and DeBroek and Sloek (2001) find strong evidence for a significantly negative impact of this variable on the real exchange rate, while Begg et al. (1999) confirm this result only at the margin. Kemme et al. (2000) find a significant explanatory power of openness for the CPI real effective exchange rate of Poland.

Using three right-hand side variables, the results for the mean group estimator suffer from a lack of degrees of freedom for panels of this dimension, particularly if they are based on DOLS and the PMGE. Therefore, the parameters of the static fixed-effects model have been used as starting values for the PMGE.[26] The results are summarised in Table 6.

The results suggest again that the signs and magnitudes of the coefficients are robust to the econometric approach chosen. Per-capita income has the expected positive sign in all specifications, suggesting that higher per-capita income is associated with an appreciation of the real exchange rate. Both for FMOLS and DOLS, the coefficients based on the weighted estimates are slightly lower than for the unweighted regressions and the PMGE. Overall, the estimated elasticities are broadly in line with the results of the bivariate model and the cross-section analysis. The coefficient of openness has a negative sign, implying that greater openness is associated with a depreciation of the real exchange rate. In line with the findings of other papers, government spending has a positive impact in the real exchange rate, suggesting that the demand channel seems to outweigh the confidence channel for government spending for the period under review. Accordingly, higher government spending leads to an appreciation of the domestic currency in real terms. Again, the coefficients are robust to the use of different estimators.

Table 6: Estimation results: (P)MGE, DOLS and FMOLS

	FM-OLS		DOLS		PMGE
Dep. var: egap	unweighted	weighted	unweighted	weighted	
ypcr	0.402 *(23.5)*	0.235 *(19.6)*	0.360 *(10.6)*	0.185 *(9.7)*	0.569 *(7.5)*
open	-0.174 *(-7.7)*	-0.324 *(-16.7)*	-0.119 *(-2.8)*	-0.189 *(-5.2)*	-0.379 *(-7.2)*
gov	0.313 *(16.4)*	0.369 *(27.0)*	0.219 *(6.2)*	0.267 *(11.0)*	0.393 *(6.5)*
φ					-0.317 *(-7.8)*
Note: The PMG was estimated using a ARDL (2,1,1,1) specification. The DOLS estimation employs one lead and one lag.					

5. Out-of-sample extrapolation ("simulation") stage: Some methodological considerations

In the panel estimation approach, the value of the constant term is bound to affect the level of the fitted value for the computation of an equilibrium exchange rate. In the estimation stage, the fixed effects estimated for each country in the panel regression provide estimates of such intercepts for each country in the sample. However, since the acceding countries are not included in the panel, their own specific constant term, that could be applied to derive the equilibrium exchange rate, cannot be estimated. Incidentally, including these countries in the

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[26] This also affects the Hausman tests for homogeneity, which was rejected in some cases. In a intermediate step, however, two models with three variables each have been estimated. Both included the exchange rate gap and per-capita income as well as government spending or openness. In both cases we did not find any evidence that the assumption of poolability is violated, so that we assume that it also holds in the case when all variables are included simultaneously.

regression does not solve the problem.[27] As explained in more detail in section 2, this may distort particularly the constant term, if one does not account for the initial episode of undervaluation of acceding countries' currencies and their gradual adjustment to more "reasonable" levels. As a result, one needs to estimate acceding-country-specific intercepts indirectly. This amounts to *choosing* a constant term that can be rationalised. Should such an approach be adopted, given its *ad hoc* nature, a thorough robustness check of the results becomes imperative.

The available literature may provide some guidance in this regard: Halpern and Wyplosz (1997) and Begg et al. (1999), for instance, derive their estimates of the equilibrium dollar wage – as they define the equilibrium real exchange rate – by: (1) including a planned economy dummy (capturing, however, mainly observations under the planned economy period), (2) setting this dummy to zero to obtain a dollar wage consistent with "normal" economic conditions, and (3) adding the effect of an OECD dummy, because converging to OECD averages can be considered to be the ultimate goal of the acceding countries.[28] They present different "equilibrium" exchange rates for each country, which they interpret as a kind of lower and an upper bound. However, the range implied by these estimates is very wide. Compared with the specification without any dummy, belonging to the planned economies group implies a discount of 33%, while being a member of the OECD leads to an increase in the "equilibrium" exchange rate by 50% (in log terms). Based on data for 1997, Begg et al. (1999) find that all countries have been below the upper bound, i.e. the "OECD equilibrium". The actual exchange rates of the Czech Republic and Slovakia were even below the lower bound of this range, suggesting that these currencies were significantly undervalued at that time. The dollar wages of Lithuania and Estonia were between the lower bound and the specification without the dummy, suggesting still some potential to appreciate; Poland and Hungary were close to these levels, while both Slovenia and Latvia were above these levels. Krajnyak and Zettelmeyer (1998) broadly confirm these results, finding that all currencies were significantly undervalued in 1995.

Kim and Korhonen (2002) derive equilibrium exchange rates on the basis of the common intercepts provided by the PMGE.[29] As regards the US dollar based exchange rates, they find that most acceding countries' currencies were in line with their equilibrium level in 1999, with the exception of the Slovenian tolar, which appeared undervalued. By contrast, on the basis of real effective exchange rates, most of the acceding countries' currencies appeared overvalued by the end of the 1990s, indirectly suggesting – in combination with the finding of equilibrium against the US dollar – a strong overvaluation of these currencies vis-à-vis the euro. A striking feature of these results is that the equilibrium exchange rates of these countries should have depreciated in the 1990s instead of appreciating, as the catching-up hypothesis would suggest. Moreover, according to these estimates the Hungarian forint was continuously overvalued from 1991 onwards.

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[27] In the present data set, an inclusion of these countries is practically not possible since this would lead to a severe lack of degrees of freedom in the derivation of the country-specific constant term, but it may be considered as an option in a data set using higher frequency data.

[28] Obviously, some of the accession countries are already formal members of the OECD, yet they have not achieved the level of economic development of most other members.

[29] The authors show some robustness checks suggesting that country-specific intercepts derived on the basis of the per-capita income of these countries biases the results towards a higher overvaluation of these currencies.

The results of Kim and Korhonen, which seem to depend quite importantly, for some countries, on the degree of openness, also point to another important methodological issue, which relates to the importance of considering the level of the data appropriately. For instance, the analysis presented in section 3 suggests that the level of per-capita income is important. The panel data analysis showed that also its development over time is relevant. By contrast, cross-section regressions suggest that for other variables, such as relative openness and governments spending, the level of the variables itself does not have an impact on the exchange rate gap. Accordingly, in order to avoid an impact of the different levels of these variables on the country-specific constant term, it would seem appropriate to demean these variables for each country (across T_i). While taking deviations from the mean does not affect the estimated elasticities, it has an impact on the constant terms (fixed effects). Ignoring this issue may have important consequences for the derived exchange rate equilibrium. For instance, many CEE countries are very open to international trade. Since the coefficient of openness is negative, imposing a high level of openness shifts the equilibrium exchange rate downwards and could likely lead to finding overvalued exchange rates.

The issue of choosing a constant term for deriving equilibrium exchange rates for acceding countries in an "out-of-sample" approach, such as that of Kim and Korhonen, leads necessarily to an *ad hoc* choice. In this respect, several alternative strategies could be pursued.

- First of all, the sample average of all the constants could be used, either from the full set of countries or for subsets of countries, chosen on the basis of relevant characteristics. One such choice could be based on the **aspirations** of the acceding countries to become an integral part of the euro area. Accordingly, the average constant term of a subset of euro area countries that share some characteristics with acceding countries could be used.

- A second possible approach would be to derive a constant term on the basis of **episodes** when countries in the sample were **similar,** in terms of **per capita income,** to the acceding countries. As a result, data mainly from countries at the lower end of the income spectrum can be used. Then, the average of the used observations could be employed to calculate a constant term.

- Finally, if an overvalued currency is considered to raise more problems than an undervalued currency in view of the higher risk of having a currency crisis, a more **conservative approach** could be employed. Assuming that the correct constant for the acceding countries is within the range of the estimated country-specific constants, the lowest constant term of the euro area countries could be chosen. This, *ceteris paribus*, would most likely lead to finding a currency overvaluation.

As remarked above, any of these choices is obviously *ad hoc*, and this underscores the importance of conducting a detailed robustness check. An advantage of the suggested methodology would be that combining the point estimates of different econometric methodologies and the different reasonable constant terms should give an array of equilibrium exchange rates for each country, against which the level of the actual exchange rate could be gauged. Accordingly, rather than being able to provide a precise estimate of an equilibrium exchange rate, such a procedure may mainly help to identify major exchange rate misalignments.

6. Conclusions and outlook

This paper provides a discussion of fundamental methodological issues and a selected overview of the academic literature on equilibrium exchange rates of acceding countries, and it applies different methods to estimate the long-term relationship between exchange rates and a set of selected macroeconomic fundamentals. The analysis in the paper suggests that estimating equilibrium exchange rates is not straightforward for these countries, in view of data limitations and of the rapid structural change they have been undergoing in the transition phase. Consequently, it cannot be expected that any statistical approach can deliver a precise value for the equilibrium exchange rate. On the other hand, many approaches may be useful in providing a qualitative indication of the extent to which the exchange rates of acceding countries' currencies are in line with economic fundamentals.

This overview suggests that from a conceptual point of view, a two-step approach, where an equilibrium exchange rate model is estimated for non-acceding countries in the first step and then calibrated in the second step for acceding countries on the basis the estimated relationships, could appear to be a promising avenue. This paper focuses on the first step, estimating a long-run relationship between the exchange rate and fundamentals for a sample including annual data for 25 OECD countries between 1975 and 2002. Building on a cross-section analysis of the relationship between real exchange rate levels and productivity developments, the analysis is extended using a panel cointegration model, which is estimated using various recent econometric techniques. A broader range of macroeconomic fundamentals is also considered in the panel data model of the real exchange rate. The results indicate that in the medium to long term the real exchange rate depends on developments in relative per-capita income (as a measure of productivity), relative government spending and relative openness. While higher per-capita income and government spending give rise to an equilibrium appreciation of the currency, an increase in openness is associated with equilibrium depreciation.

The "out-of-sample" calibration of the equilibrium real exchange rate for acceding countries requires additional assumptions. As discussed in the last section of this paper, where we present a few methodological considerations in this respect, the most important problem in this context is the choice of the constant term. We discuss a few plausible strategies in this regard. Although not purged of shortcomings, given its relatively *ad hoc* nature, the two-step approach has the advantage that combining the different econometric methodologies and the different reasonable constant terms should give an array of point estimates for the equilibrium exchange rate for each country.

ANNEX I

SMALL-SAMPLE PROPERTIES OF PEDRONI TESTS

In order to evaluate the performance of the Pedroni tests for the null hypothesis of cointegration in the small sample used in this paper, a simple Monte Carlo simulation experiment was conducted. The data were simulated under the null hypothesis, with N = 25 and T = 28. The empirical critical values, which should be compared with the standard normal ones, are reported in Table A1.

Table A1: Small-sample empirical critical values of Pedroni tests

	99%	97.5%	95%	90%	10%	5%	2.5%	1%
Panel ν-Statistic	-1.61	-1.39	-1.20	-0.95	1.15	1.51	1.81	2.18
Panel ρ-Statistic	-2.03	-1.68	-1.41	-1.06	1.07	1.32	1.54	1.78
Panel t-Statistic (non-par.)	-2.85	-2.53	-2.23	-1.87	0.68	1.03	1.33	1.67
Panel t-Statistic (parametric)	-3.01	-2.67	-2.38	-2.03	0.55	0.89	1.17	1.55
Group ρ-Statistic	-0.97	-0.67	-0.38	-0.10	1.87	2.13	2.32	2.57
Group t-Statistic (non-par.)	-2.96	-2.54	-2.18	-1.77	1.04	1.42	1.77	2.19
Group t-Statistic (parametric)	-3.27	-2.85	-2.50	-2.08	0.75	1.15	1.51	1.93

In order to gauge the small-sample size distortion of each test at T = 28 and T = 25, Table A2 reports the empirical rejection frequencies corresponding to each nominal size.

Table A2: Small-sample empirical size of Pedroni tests

|	99%	97.5%	95%	90%	10%	5%	2.5%	1%
Panel ν-Statistic	0.9999	0.9984	0.9915	0.9616	0.0793	0.0374	0.0173	0.0071
Panel ρ-Statistic	0.9991	0.9955	0.9824	0.9436	0.0646	0.0279	0.0119	0.0042
Panel t-Statistic (non-par.)	0.9982	0.9959	0.9892	0.9723	0.2477	0.1431	0.0843	0.0405
Panel t-Statistic (parametric)	0.9986	0.9965	0.9927	0.9799	0.2931	0.1787	0.1113	0.0563
Group ρ-Statistic	0.9755	0.9192	0.8314	0.6830	0.0041	0.0004	0.0001	0.0000
Group t-Statistic (non-par.)	0.9932	0.984	0.9689	0.9347	0.2014	0.1219	0.0739	0.0376
Group t-Statistic (parametric)	0.9964	0.9912	0.9825	0.9611	0.2897	0.1881	0.1204	0.0666

Looking at the four rightmost columns, Table A2 shows that most tests are oversized, i.e. they tend to reject the null hypothesis of no cointegration too often. The panel-v and panel-rho test have the least size distortions, while the group-rho test is severely undersized.

In order to also analyse the deviation of the distribution of each test statistic from the standard normal distribution, their quantiles are compared to the standard normal quantiles in Charts A1 and A2. Chart A1 shows the empirical distribution of each test along with a standard normal distribution, while Chart A2 shows the plot of the empirical quantiles against the standard normal quantiles. It can be seen that the panel-rho and the two panel-t tests display the most "non-normal" distributions, having a greater probability mass on the left tail.

Chart A1. Empirical ditributions vis-à-vis the standard normal distribution.
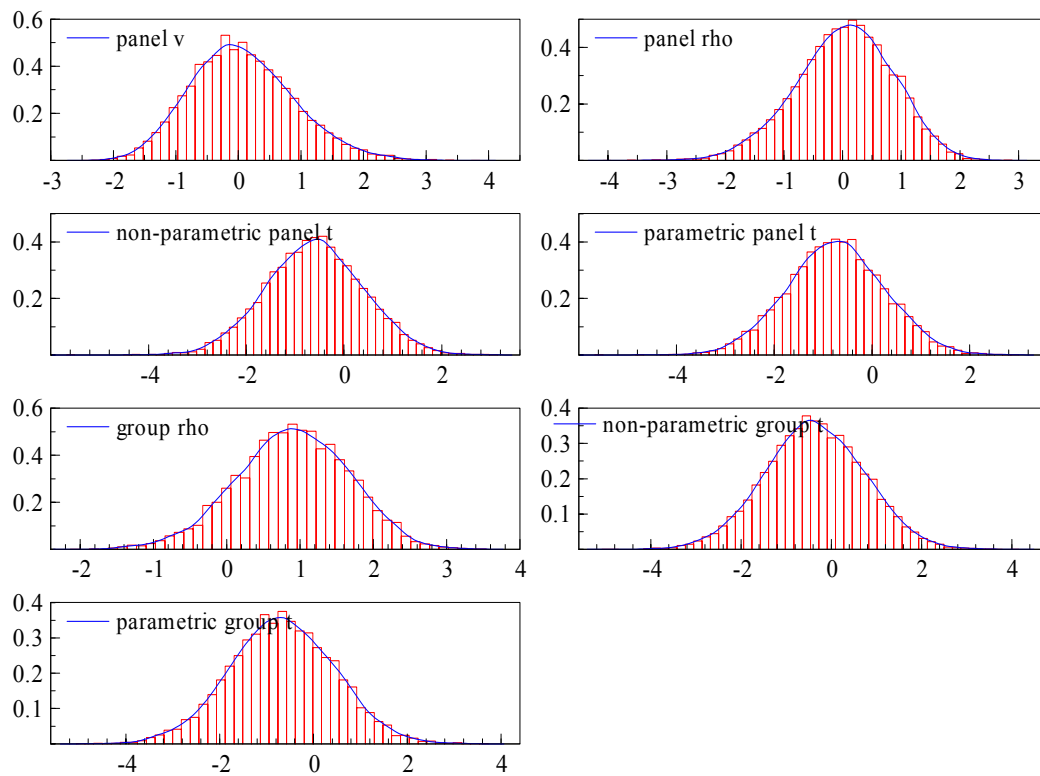

Chart A2. QQ plots of the empirical quantiles vis-à-vis the quantiles of the standard normal distribution.
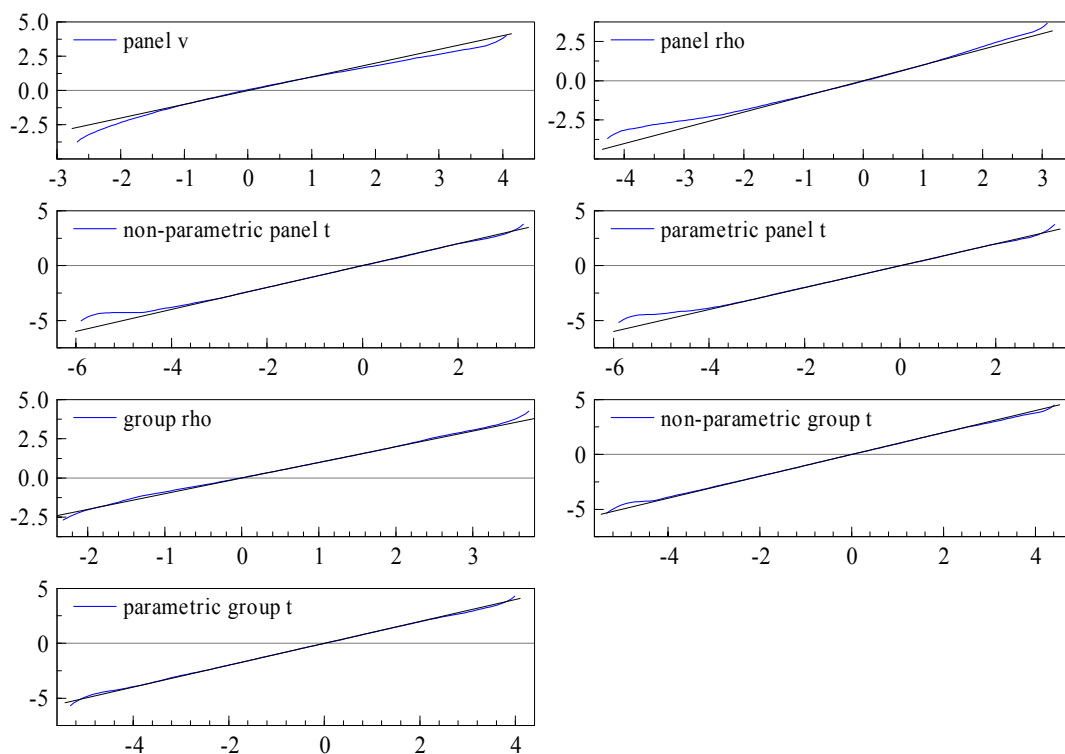
In order to also gauge the power of the tests, a simple further Monte Carlo experiment was run, using the moments of the empirical distributions to "standardise" further the seven Pedroni tests, in order to at least partially adjust for size. This adjustment does not take into account the deviations of the distributions from the standard normal, but it was found that after this further adjustment, the empirical quantiles corresponded quite closely to the nominal ones. Using this correction, power curves were calculated, with different values for the autoregression coefficient of the residual under the alternative hypothesis. Chart A3 shows these power curves.

Chart A3: Power curves at size = 5%, N = 25, T = 28
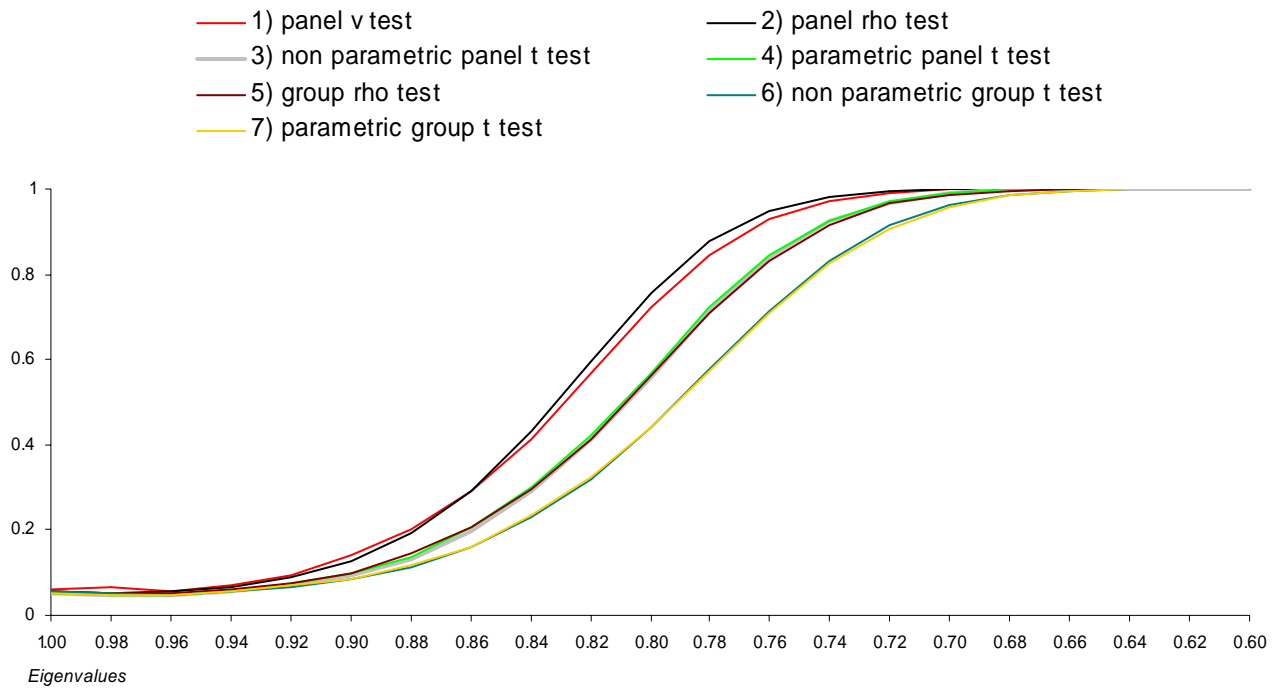
ANNEX II

DATA SOURCES:

- **Exchange rate gap:** Purchasing power parity exchange rates – AMECO: 1 0 212 0 KNP in national units per PPS. Market exchange rates – AMECO: 1 0 99 0 XNE in national units per ECU/EUR. Eurostat interpolates the data which are compiled at a three year frequency from 1993 and a five year frequency in the 1970s and 1980s. Data for the first part of the period is based on ICP 1979 and on ICP 1995 thereafter. Accordingly, methodological consistency can not be fully ensured over the full sample period.

- **Gross domestic product per head, total economy:** AMECO: 1 0 212 0 HVGDP, The euro area aggregate chain-links in 1997 the series AME.A.d11.1.0.212.0.hvgdp and AME.A.e11.1.0.212.0.hvgdp. For Germany, the data for West Germany was chain-lined with data for FR Germany in 1997.

- **Gross domestic product per person employed, total economy (productivity):** AMECO: AME.A.d11.1.0.212.0.HVGDE. The euro area aggregate chain-links in 1997 the series AME.A.d11.1.0.212.0.hvgde and AME.A.e11.1.0.212.0.hvgde. For Germany, the data for West Germany was chain-lined with data for FR Germany in 1997.

- **Average of exports and imports of goods to GDP (openness):** AMECO: 1 0 310 0 DAGT. The euro area aggregate was proxied by data for the EU-15. AME.A.e65.1.0.310.0.dagt lined with AME.A.d65.1.0.310.0.dagt in 1991. For Germany, the data for West Germany was also chain-lined with data for FR Germany in 1991.

- **Total final consumption expenditure of general government to GDP (government spending)**: AMECO: 1 0 310 0 UCTG. The euro area aggregate was proxied by data for the EU-15. AME.A.e65.1.0.310.0.uctg lined with AME.A.d65.1.0.310.0.uctg in 1991. For Germany, the data for West Germany was also chain-lined with data for FR Germany in 1991.

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