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.

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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 (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 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 4 presents the data.

Section 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 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 7 where we discuss whether we should reconsider the traditional way of computing the long run pass-through.[1]

[1] Detailed results for single equation estimates with and without breaks, using the Gregory and Hansen (1996) algorithm, together with all tables and graphs reproduced for the CM sample 1989-2001 and detailed descriptions of all tests used are available in the working paper version of this paper, De Bandt, Banerjee and Kozluk (2006).

2 Exchange Rate Pass-Through into Import Prices

By definition[2] 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):

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

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^j into equation (2) yields:

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

when looking at an individual sector in a country, we can simplify the notation by dropping superscript j.

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. Generally, a larger share of imports in total industry supply, higher degree of price discrimination or a larger share of imported inputs in the production in the destination country leads to a higher predicted pass-through. ERPT may be higher if the ratio of exporters relative to local competitors is high, and lower if exporters compete for market share, even if nominal exchange rate variability is high. 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

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

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 c_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)[3]. Thus the final equation can be re-written 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.[4]

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'

[3]The integrated market hypothesis in CM 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 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 denominated in US dollars are taken as a proxy for the world price.

[4]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'.

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 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 the Economic and 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.

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), CM 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 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. Similar patterns of fluctuation of the coefficient estimates are also observed for some other sectors in Table 1 (and for other countries).

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.

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 5, following a brief description of the data in Section 4 below.

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 2 of De Bandt, Banerjee and Kozluk (2006). 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.

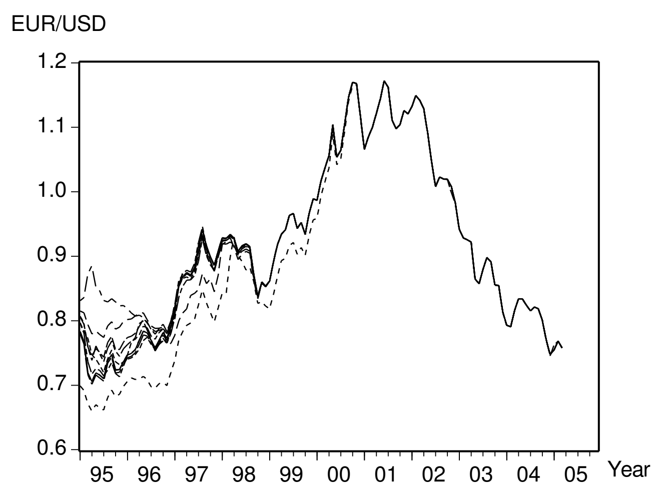

Figure 1: Monthly index of exchange rates of euro area currencies versus the USD. 1995-2005.

The indicator we use for import prices, the index of import unit values (IUV) has a series of caveats concerning their 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.

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.

Further following from our discussion in Section 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.[5] 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.

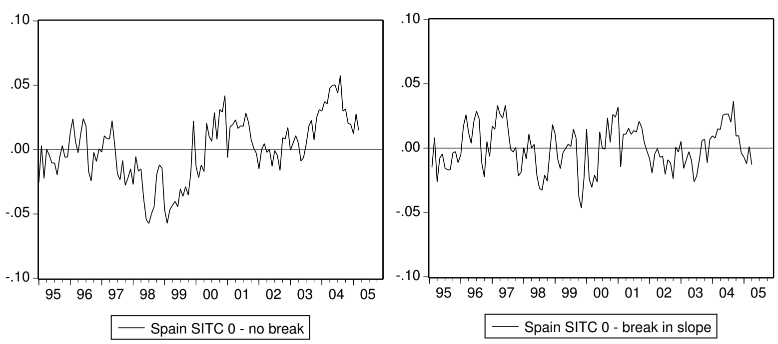

Figure 2: Residuals from the estimation of equation (9) without a break (left) and with a single estimated break (right) on the series for Spain, SITC0.

Anticipating to some extent our future results, on left hand side of Figure 2 we show the errors from the estimation of the levels equation (9), for which as we will see in Section 5.1 it is quite hard 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 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

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

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	0.64	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	0.34	0.89	0.68	0.94	0.95	0.93	0.87	0.92	0.94	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	0.44	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	0.44	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.

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 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%) 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. 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.

Finally, looking at the exchange rates of current euro area currencies in Figure 1 we see 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.

5 Results

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 Table 3) only partially 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 over 50% of the series (at 5% level). Moreover, as we see in Appendix 2 of De Bandt, Banerjee and Kozluk (2006), 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 about three quarters 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.

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

H_0: Unit root (no cointegration)

Country	SITC0	SITC1	SITC2	SITC3	SITC4	SITC5	SITC6	SITC7	SITC8
France	**-3.61***	**-3.37**	**-3.03**	**-5.87***	-1.84	**-5.49***	**-3.73***	-2.41	**-5.16***
Netherlands	**-3.23**	**-3.14**	**-4.48***	**-2.96**	**-4.06***	**-3.11**	-2.42	**-3.4**	**-6.45***
Germany	-1.45	-2.28	-2.3	**-3.85***	**-3.73***	**-3.77***	-2.38	**-3.65***	**-5.68***
Italy	-1.91	-1.73	-1.49	**-5.59***	**-4.66***	**-3.4**	**-2.95**	-1.98	-2.52
Ireland	0.33	**-4.86***	**-4.96***	**-5.24***	-2.14	**-9.28***	**-2.97**	**-5.04***	**-6.82***
Greece	-1.81	**-3.67***	-2.26	**-2.71***	**-2.58***	**-5.15***	**-2.9**	**-3.46**	**-5.22***
Portugal	-2.46	-1.82	**-2.83***	**-4.3***	-1.58	**-9.09***	-2.27	**-4.18***	**-4.27***
Spain	**-2.59***	**-3.57***	**-3.63***	-2.3	**-6.16***	**-3.19**	-2.55	**-3.6***	**-5.53***
Finland	-0.8	**-3.49**	**-2.78***	**-2.69***	**-3.85***	**-3.04**	**-3.46**	**-2.74***	**-6.33***
Austria	**-6.46***	**-7.73***	**-5.45***	**-4.63***	-2.43	**-8.1***	**-3.14**	**-3.02**	**-6.41***

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

Specification: 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 test results with alternative lag-selection criteria and for the CM 1989-2001 sample we refer the reader to the working paper De Bandt, Banerjee and Kozluk (2006).

break. We test two alternative versions of the model proposed in equation (9). First, a break in the constant, thus a levels shift:

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

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

$$mp_t = \widehat{\alpha} + \widehat{\alpha}_1 * d_s + \widehat{\beta} \cdot er_t + \widehat{\beta}_1 \cdot er_t * d_s + \widehat{\gamma} \cdot fp_t + \widehat{\gamma}_1 \cdot fp_t * d_s + \upsilon_t \qquad (13)$$

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.[6] 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, 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 4 (for the GH single equation tests[7]) 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 4 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.[8] In more detail, when we allow for the more general break as in equation (13), for the GH single equations for 40 out of 61 series α_1's are negative (of which 24 are significant at 10%), leaving only 7 positive and significant (at 10%).

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.[9] Overall the most important

[6]Brief details of the procedure are contained in Appendix B.

[7]In order to save space, we report only the directions, significances and dates of breaks estimated with the GH algorithm. These are reported in Table 4 if the null hypothesis of unit root (i.e. no cointegration) can be rejected at 10%. As for the coefficient estimates, they are essentially very similar to the ones in Table 6 for the break in constant specification and Table 7 for the break in entire cointegrating vector specification, as the Banerjee and Carrion-i-Silvestre (2006) adaptation of the Pedroni (1999) test to account for breaks builds on the GH specification. However, as due to data availability the single-equation GH estimation is based on a longer (by one year 1995) series for all countries except Finland and Austria, we refer the interested reader to De Bandt, Banerjee and Kozluk (2006) for detailed results of the GH specification.

[8]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.

[9]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,

	France		Netherlands		Germany		Italy		Ireland		Greece		Portugal		Spain		Finland		Austria	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
SITC_0		6/00 - +		2/99 + -*					6/03 +***	4/02 -*** +*	2/02 +***	8/01 - +	6/03 +***	7/02 -*** +***		9/99 -*** +***	2/03 -***	7/99 + -	4/98 +	2/98 - +
SITC_1		8/02 -*** +***	7/98 -***	1/98 + -**			10/00 +***	10/00 -*** +***	11/00 -***	2/03 + -				1/00 -*** +***	9/03 +***		1/03 -	9/01 - +		
SITC_2	1/97 -***					2/99 - +***	3/97 -***	11/97 +** +***	9/03 +***	9/03 - +			*1/99* -***	*1/99* + -**	3/98 -***	3/98 - +			1/98 +***	*1/97* +* -
SITC_3	1/01 -***	1/01 -*** +*			8/96 -				6/00 +***	5/03 - -	10/02 +***	4/00 +*** -***	*10/02* -***	11/98 - +**					12/97 +***	12/96 - +
SITC_4	2/00 +***	2/00 +** -***	1/99 -***	4/99 -*** +***	*12/00* -***	12/00 + -	10/96 -**	7/99 -*** +***				12/00 - +	12/02 +***	12/02 + -			9/97 +***	9/96 - +		
SITC_5	2/01 +***	1/01 -*** +***	12/02 -***	9/02 -*** +***	*1/03* -***		4/97 -***	9/98 + +***	12/97 +	12/97 - +	2/03 -***	11/98 + +	3/98 -**	8/98 -*** +**	6/97 -***	2/97 - +	1/03 -***	4/02 -** +	4/98 -***	2/98 -** +*
SITC_6	4/01 +***	*7/01* + +	11/01 -***		*3/98* +***				5/98 +***	8/02 -*** +***	*4/02* +***		*8/02* -***		11/01 -***	11/01 -*** +***		12/99 -*** +	9/00 +***	9/99 - +**
SITC_7	7/98 -***	7/98 -** +**	9/98 -***	9/98 + -	1/01 +**	10/00 -*** +***	12/97 -***	1/98 +* +				*1/98* - +	6/03 -***	1/02 + -					3/02 -***	*5/01* + -
SITC_8	7/96 +***	7/96 + +***	9/01 -***	8/01 -** +***				1/98 -*** +***	7/96 +***	*5/00* -*** +***			4/99 -***	*5/99* -** +***	11/98 -***	12/99 - +***	6/97 +***	4/99 + +***	9/97 +**	

For each country industry combination column (1) gives: first row - date of estimated break in constant, second row: direction and significance of the break in constant.
Column (2) gives: first row - date of estimated break, second row: direction and significance of the break in constant, third row: direction and significance of the break in slope.
*+ or - indicate whether the change is positive or negative, *,**,*** indicate whether it is significant (t-stat of $\widehat{\alpha}_1$ and $\widehat{\beta}_1$ respectively) at 10%, 5% and 1%.*
Value not reported if the hypothesis of unit root (no cointegration) cannot be rejected at 10%, (Italics) if it cannot be rejected at 5% but can be rejected at 10%.

Table 4: Estimated directions, significances and dates for breaks from the GH algorithm. For each country/industry combination column (1) represents the specification of break in constant and column (2) represents the specification of break in the entire cointegrating vector. Sample: 1996-2004.

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.[10]

A selection of the single equation results for the "long run" ERPT are presented in Figures 3 and 4. 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 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 $9 \times 10 \times 110$

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.

[10]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.

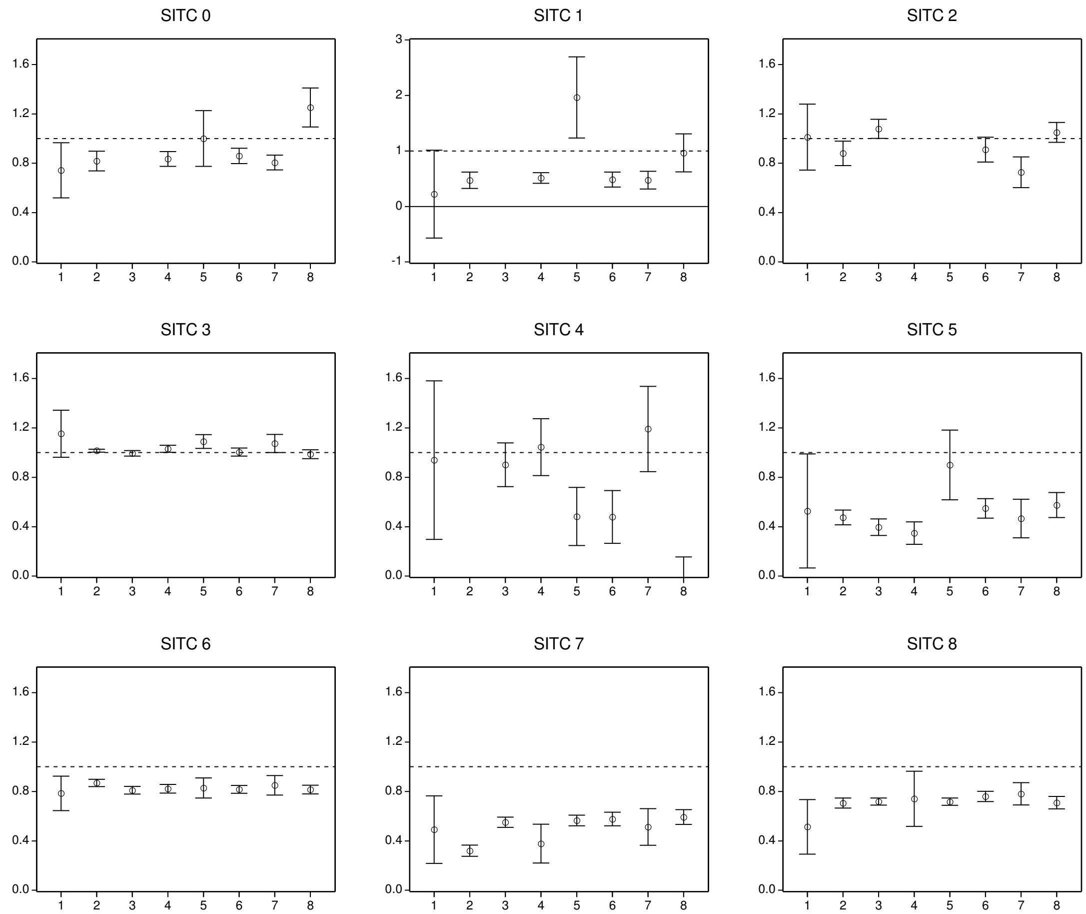

Figure 3: France - "long run" ERPT estimates with confidence intervals (95%). Individual industries, sample: 1995-2005, entire sample analysis. The estimators are presented in the following order: (1) CM long run, no cointegration, no break, (2) cointegrating long-run, no break, (3) cointegrating long run, break in constant (estimated, GH), (4) cointegrating long run before break in slope (estimated, GH), (5) cointegrating long run, after break (estimated, GH), (6) cointegrating long run, break in constant (imposed on 1998/99), (7) cointegrating long run, before break in slope (imposed on 1998/99), (8) cointegrating long run, after break (imposed on 1998/99). In (3)-(5) values extracted from GH algorithm, ADF*. Values not reported if no cointegration (ADF). Break dates estimated with the GH are available together with equivalent graphs for Germany and Portugal in De Bandt, Banerjee and Kozluk (2006). Dotted horizontal line at value of 1.

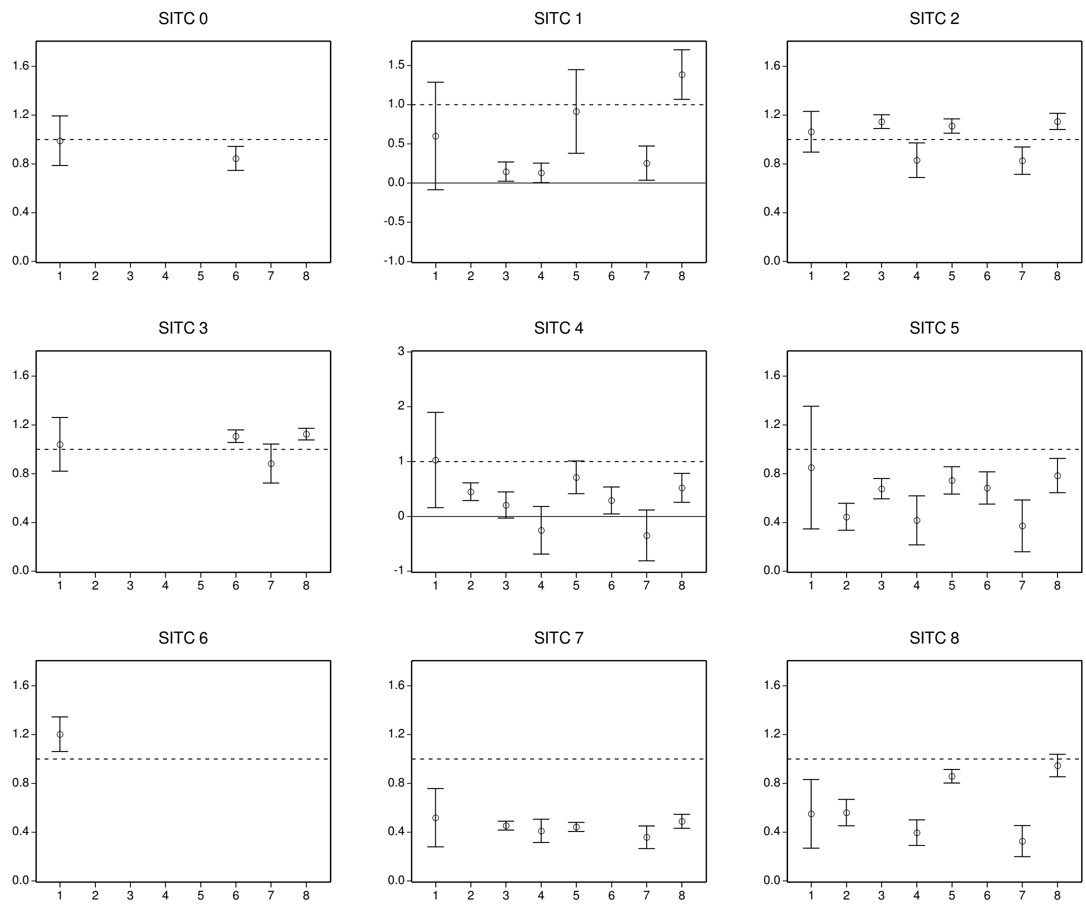

Figure 4: Italy - "long run" ERPT estimates with confidence intervals. Individual industries, sample: 1995-2005. Notes - see Figure (3) for explanations.

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.

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 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.[11] 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

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 5: Pseudo-t and Pseudo-ρ parametric statistics from De Bandt, Banerjee and Kozluk (2006). The null hypothesis is no cointegration. Sample: 1996-2004, full panel (N=90), unit specific breaks.

cointegration tests with no cross-sectional dependence and no breaks are displayed in the first row of Table 5. 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

[11]Brief details of these tests are contained in Appendix B.

mentioned, we suspect the formation of the euro area constituted a shift in both competition conditions and monetary policy which may 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 6 and 7.

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 7 that out of 90 series, 60 have a negative estimated α_1 coefficient, of which in 35 cases they are significant, while only for 10 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.[12] The estimated break dates for all the individual series are presented in Figure 5. There is

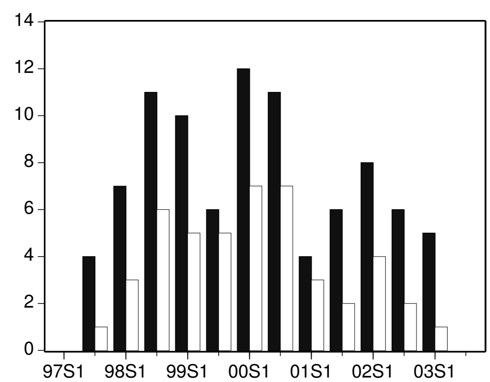

Figure 5: Distribution of estimated break dates by semester (1997s1-2003s2). Breaks in slope taken from Tables 7. Dark color - all breaks, light color - only breaks when long run ERPT changed significantly (10%).

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).

Although the evidence, as presented in Tables 6 and 7, in favor of both cointegration and structural change is unequivocally strong, a few qualifications are worth noting. First,

[12]It is worth noting that also in the case of the specification with the break in constant only, 52 out of 90 of the estimated changes in the constant are negative (51 of which significant at 10%) though admittedly many more are significantly positive (38) than in the case of the more flexible specification (break in constant and slope) which we treat as an argument in favor for the latter. This is in line with the argument that the formation of the EMU may have increased the power of 'domestic currency denominated' products, or reduced uncertainty associated with the exchange rate leading to a fall in the constant mark-up component charged by 'foreign currency denominated' products.

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)	(0.04)	–***	(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 6: 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 6: 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 6: 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 7: 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 7: 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 7: continued.

the GH based algorithm here allows for only one, ”strongest”[13] 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 instanteous 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 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

H_0: unit root (no co-integration)

No. of factors	Pseudo-t - $ADF^c_{\hat{e}}(i)^*$		
	No break	Break (a)	Break (b)
(1)	-4.51	-3.72	-3.11
(3)	-3.71	-3.47	-2.60
(6)	-3.76	-3.00	-3.47

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

Table 8: Test statistics for Banerjee and Carrion-i-Silvestre (2006) panel co-integration tests with cross-section dependence (common factors). Sample: 1996-2005.

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

[13] 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).

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 8 and we consider them as reconfirmation of the existence of a long run equilibrium relationship.

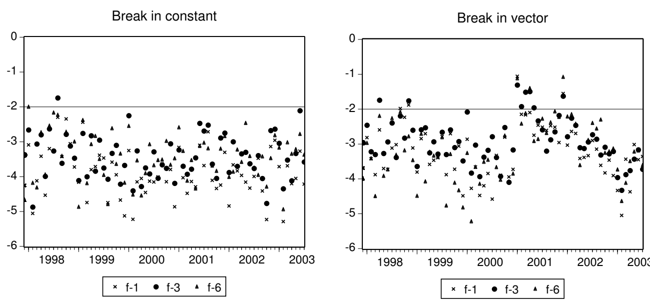

Figure 6: Distribution of the test statistics from Banerjee and Carrion-i-Silvestre (2006) test. The null hypothesis is no cointegration, the alternative cointegration with a common break in constant (left hand side) as in equation (12), and cointegration with common break in cointegrating vector(right 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.

The results in Table 8 are computed with the common break date imposed at 1999m1. More general results are plotted in Figure 6 where we report the Banerjee and Carrion-i-Silvestre (2006) test statistic for all possible choices of common break dates (subject to 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.

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[14], which we discussed in Section 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 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

[14]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.

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.[15]

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 7 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 (significant 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

[15] 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.

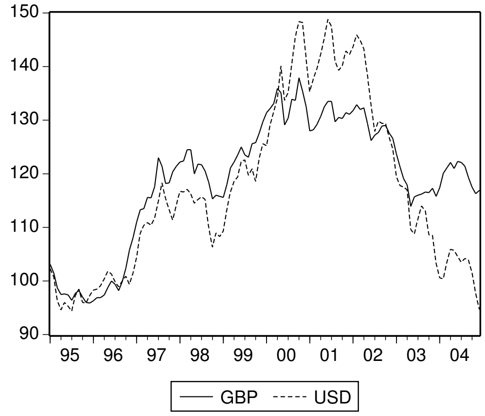

Figure 7: Evolution of EUR(ECU)/USD and EUR(ECU)/GBP exchange rates(1996=100).

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 7). 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 on 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 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 first of all it is more general, and moreover in this case the estimated shift in the constant tends to be more 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

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.
Sources: Eurostat (COMEXT). 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 - Descriptions of tests[16]

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)

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 $T_{bi} = \lambda_i T$, $\lambda_i \in \Lambda$, where Λ is a closed subset of $(0, 1)$, for the applications considered Λ is given by $[0.15, 0.85]$, 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 EG and GH

[16] This Appendix is an 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.

but in the panel data framework developed in Pedroni (1999, 2004). In fact, GH propose both of the specifications given by models 1 and 4 above.

The Banerjee and Carrion-i-Silverstre (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 $(\hat{e}_{i,t}(\lambda_i))$:

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

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

$$\hat{T}_{bi} = \arg\min_{\lambda_i \in \Lambda} t_{\hat{\rho}_i}(\lambda_i); \quad \hat{T}_{bi} = \arg\min_{\lambda_i \in \Lambda} T\hat{\rho}_i(\lambda_i), \quad (22)$$

$\forall i = 1, \ldots, N$. 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 EG. 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:

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

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

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: Banerjee and Carrion-i-Silvestre (2006)

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 test described here is for a common (across all the units), known structural break point although this can be generalized in some cases (see Banerjee and Carrion-i-Silvestre, 2006).
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

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

We denote by $ADF_{\tilde{e}}^{c}(i)$ the pseudo t-ratio ADF statistics for testing $\alpha_{i,0} = 0$ in (35). As in (24) the individual ADF statistics $ADF_{\tilde{e}}^{c}(i)$ for the idiosyncratic disturbance terms can be pooled to define a panel data cointegration test. Thus, one can define

$$N^{-1/2} Z_{\hat{t}_{NT}}^{e} - \Theta_2^{e}\sqrt{N} \Rightarrow N\left(0, \Psi_2^{e}\right), \tag{36}$$

where the superscript e denotes the idiosyncratic disturbance. The moments Θ_2^e and Ψ_2^e are the same as the ones for the statistics in Bai and Ng (2004), where these do not depend on the break fraction λ.