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Measuring Long-Run Exchange Rate Pass-Through

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Abstract:
The paper discusses the issue of estimating short- and long-run exchange rate pass-through to import prices in euro area countries and reviews some problems with the measures recently proposed in the literature. Theoretical considerations suggest a 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. The main finding is that allowing for possible breaks around the formation of EMU and the appreciation of the euro starting in 2001 helps restore a long run cointegration relationship, where over the sample period the fixed component of the pass-through decreased while the variable component tended to increase.

JEL: F14, F31, F36, F42, C23
Keywords: Exchange rates; pass-through; import prices; panel cointegration; structural breaks

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1 Introduction

A large number of recent papers (see for example Campa and González Mínguez, 2006; Campa, Goldberg and González Mínguez, 2005; Frankel, Parsley and Wei, 2005; Marazzi et al., 2005) have investigated the issue of exchange rate pass-through (ERPT) to domestic prices. Studies of ERPT have been conducted both for the United States and for countries of the euro area, with a particular focus 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 monetary and financial shocks.

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, the accuracy of inflation forecasts, the impact of entering into a monetary union or structural reforms across the European Union. For the countries belonging to the euro area, the issues listed above are particularly relevant.

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 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, policy actions should in principle look beyond short-term inflation developments for a better understanding of the underlying forces.

Since it is commonly agreed that the time series considered are integrated, one way of defining the long run is in the sense of Engle and Granger (1987), henceforth EG, where the long run is given by the so-called cointegrating relationship. The reason for ignoring this long run, and substituting it by an ad hoc measure, is the failure to find evidence in the data for cointegration. The difficulty inherent in such a re-definition of the long run is two-fold, first the contradiction between a theoretical prediction of a steady state that cannot be found in the data, and, second, the ad hoc measure proposed being no more than an extended version of the estimate of the short-run (and, as we shall see below, strongly dominated by the estimated short-run). We therefore look for the long-run relationship using methods which allow for changes in the long run or use more powerful panel data methods.

Focusing on a specification of ERPT normally used in the literature, 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 sensitive to 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 them; 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 (e.g. Italy's re-entry into the ERM.) Not to take explicit account of such changes, which are evident in the data, could be to make mistakes in estimation and inference.

We begin in the next section with a 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, based on Campa and González Mínguez (2006), CG hereafter. We present the definitions of short- and long-run ERPT assumed by the empirical literature and assess their adequacy. Section 4 presents the data.

Section 5 provides our estimates of the Engle-Granger long run wherever these exist, also allowing for structural breaks in the cointegrating vector using methods developed by Gregory and Hansen (1996). We 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.

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.

A general discussion of our results is contained in Section 7. Section 8 concludes the paper.

2 Exchange Rate Pass-Through into Import Prices

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

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

In logarithms (depicted in lower case):

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

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

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

So that in logarithms we have:

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

Substituting for xp_t into equation (2) yields:

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

The literature on industrial organization yields insight into why the effect of a change in er_t on mp_t may differ from one, through markup determinants like competitive conditions

[1] This section is based on Campa, Goldberg and González Mínguez (2005), CGM hereafter.

that exporters have to face in the destination markets. Hence, the estimated pass-through elasticities are a sum of three effects:

- the unity translation effects of the exchange rate movement;
- the response of the markup in order to offset this translation effect;
- the changes in the marginal cost that is attributable to 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 (e. g. for commodities or oil), and lower if exporters compete for market shares (e. g. for manufactured goods), 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 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 markups 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 markup. Thus, markups in an industry are assumed to consist of a component specific to the type of good, independent of the exchange rate and a reaction to exchange rate movements:

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

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

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

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

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

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

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

equation can be re-written as follows:

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

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

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

The second issue concerns the fact that 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).

Importantly, much of the focus of the Stochastic Dynamic General Equilibrium 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 existence 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' markup. This issue is reviewed in more detail in Corsetti, Dedola and Leduc (2005). In the latter paper, the introduction of an intermediary sector which uses non-traded intermediate goods creates a long-run wedge between world prices in local currency and domestic prices.

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 (2005).

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 in the market, but also monetary policy, or expected future monetary policy (Taylor, 2000). Countries with lower inflation also experience less persistent exchange rate fluctuations, hence a lower exchange rate pass-through. The formation of the Economic and Monetary

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 CG 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-à-vis this 'world' currency. Here as in the CG case the extra-euro area imports denominated in US dollars are taken as a proxy for the world price. This might be seen as a strong assumption, but, by taking data from an homogeneous data base of IUVs for both import prices and world price, this avoids introducing additional measurement errors in the analysis.

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

Union, which occurs in the middle of the sample period used for the empirical exercise, is thus likely to have an important impact on ERPT (while the ERPT level itself may affect the strength, and exact timing of the break) 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), a number of empirical papers concerned with estimating pass-through equations are not able to reject the null hypothesis of the non-existence of a cointegrating relationship among the three series - i.e. they do not find evidence of a long-run relationship, and proceed to estimate equation (9) in differences.

A typical specification of a(n) (import) pass-through equation is therefore given by:

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

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

Table 1 presents the results from following such a methodology as reported by CG on their 1989-2001 sample. This is closest to our own data set, although our critique applies more widely.

Based particularly upon looking at the significance of the individual coefficients we have a number of comments to offer. We argue that the often encountered definition of the long run (see e.g. CG inter alia) may be inadequate for the purpose of enquiring about the actual long-run effect. For example, it is not clear why five lags are chosen. Furthermore, this measure does not take into account the significance of the coefficients on the individual lags, for which there is a great deal of evidence of multi-collinearity. 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 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 joint significance of the sum of the coefficient-estimates is generally indicative, but the high uncertainty surrounding the individual estimates does lead to difficulties in interpretation. The importance of our argument for inference can be illustrated further by taking the coefficients for SITC0 from Table 1: the five-lag 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 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). We also repeat the analysis by re-estimating the equation with four lags or six lags and reach similar conclusions. Details are available from us upon request.

Table 1: Estimates of Equation (10) - Coefficients and Standard Errors on the Lags of the Exchange Rate - Original CG Sample 1989-2001.

France

	Lag 0	Lag 1	Lag 2	Lag 3	Lag 4	CG 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. The last column reports the CG long-run estimate.

The fact that a cointegrated 'equilibrium' relationship cannot be found is 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 the final sections of this paper, following a brief description of the data in Section 4 below.

4 Data

In order to perform our estimations, we have the possible use of two data sets. The original sample, which covers the years 1989-2001, is the one used by CG and 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.

In our estimations we choose to focus on the results for a more up-to-date sample of 1995-2005, which is available from Eurostat. This data set has the advantage of extending further beyond the suspected break date related to the introduction of the euro. The construction of the variables follows CG, and is described in the Appendix A.[4]

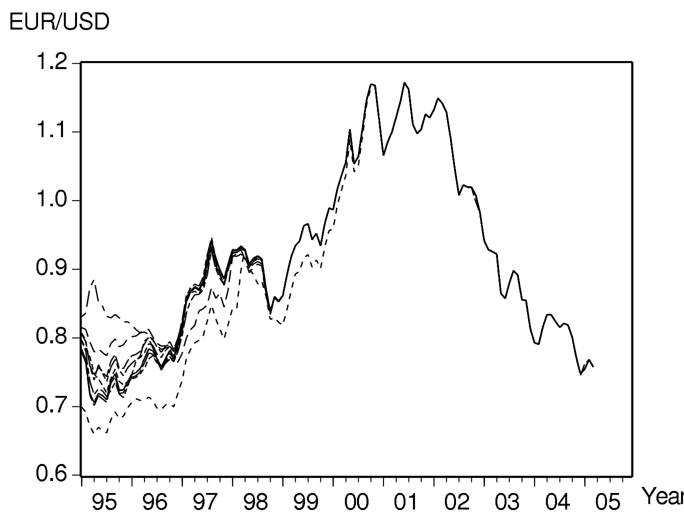

Figure 1: Monthly index of exchange rates of euro area currencies versus the USD. 1995-2005.

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 and frequency. 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

[4]Detailed results for the CG sample are available from the authors upon request. The conclusions reported in the text extend to this sample.

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 CG 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 such policy 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.

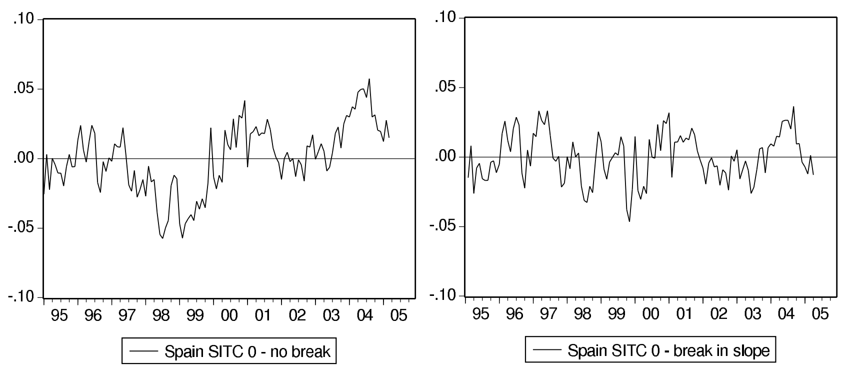

Figure 2: Residuals from the estimation of equation (9) without a break (left) and with a single estimated break (right) on the series for Spain, SITC0.

Anticipating to some extent our future results, on the left hand side of Figure 2 we show the errors from the estimation of the levels equation (9). 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 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.

Moreover the adoption of a common currency has changed the competitive conditions, by increasing the share of goods denominated in the (new) domestic currency, hence truly

[5]Greece joined the euro 2 years after the formation of the EMU, with effect on the 1^{st} of January 2001.

creating a single market for exporters.

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 with a fixed dollar price, depreciation of the euro vis-à-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 2) do not support the 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 a vast majority of the series (at 5% level). Therefore we can say that there is some evidence that a long-run relationship in levels, in the sense of Engle and Granger (1987), exists among 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 an estimated structural break. We test two alternative versions of the model proposed in equation (9). First, a break in the constant, thus a level shift:

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

thus we allow for a slope shift (i.e. a change in the elasticity of the long-run pass-through) in addition to a shift in the level:

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

In both cases d_s is a dummy variable equal to 0 if $t < s$ and equal to 1 otherwise. The GH algorithm 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 arguments 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 markup (denoted by the coefficient α) to fall rather than increase - due to potentially improved competition in the market arising from increased price transparency. Table 3 (for

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

Table 2: ADF Tests on the Errors from the OLS Regression of the ‘Long-run’ Equation (9). Sample: 1995-2005. (Test results with alternative lag-selection criteria are available from the authors.)

H_0: Unit root (no cointegration)

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

*ADF t-statistic, ***,**,* - the null hypothesis is rejected at 99%, 95% and 90% respectively.*

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

Table 3: Estimated Directions, Significance and Dates for Breaks from the GH Algorithm. Sample: 1995-2004.(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.)

	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%.

the GH single-equation tests[7]) shows clearly that in the specification that allows for only a break in the constant, the fixed component in the markup 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 3 shows that the more flexible specification of a break in slope and constant lead to the majority of the estimates pointing to a decrease or insignificant change in the fixed markup component. 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.[8] Overall the most important 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, break in slope). We treat this as strong evidence of the presence of a theory-backed long-run relationship in the data, which changes in response to key economic events.[9]

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 CG-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 estimation of equation (9) when allowing for a structural break in the entire cointegration vector (estimates (4) and (5) for

[7]In order to save space, we report only the directions, significance and dates of breaks estimated with the GH algorithm. These are reported in Table 3 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 5 for the break in constant specification and Table 6 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, 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.

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

[9]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 gradual change. This is unfortunately precluded in our study by the shortage of data.

the GH estimated break and (7), (8) for the imposed 1999m1 break) may, however, be more imprecise, especially if the estimated break happens to lie towards the beginning or end of the sample.

There are some country- and industry-specific differences 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. 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.

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

6 Panel cointegration tests

There are essentially three ways of proceeding in order to construct panels from the data sets, by creating: (1) 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 break in constant or in the cointegrating vector or both. 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, while allowing for an individual, estimated break date.[10]

The statistics for the Pedroni (1999) panel cointegration tests with no cross-sectional dependence and no breaks are displayed in the first row of Table 4. They show strong

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

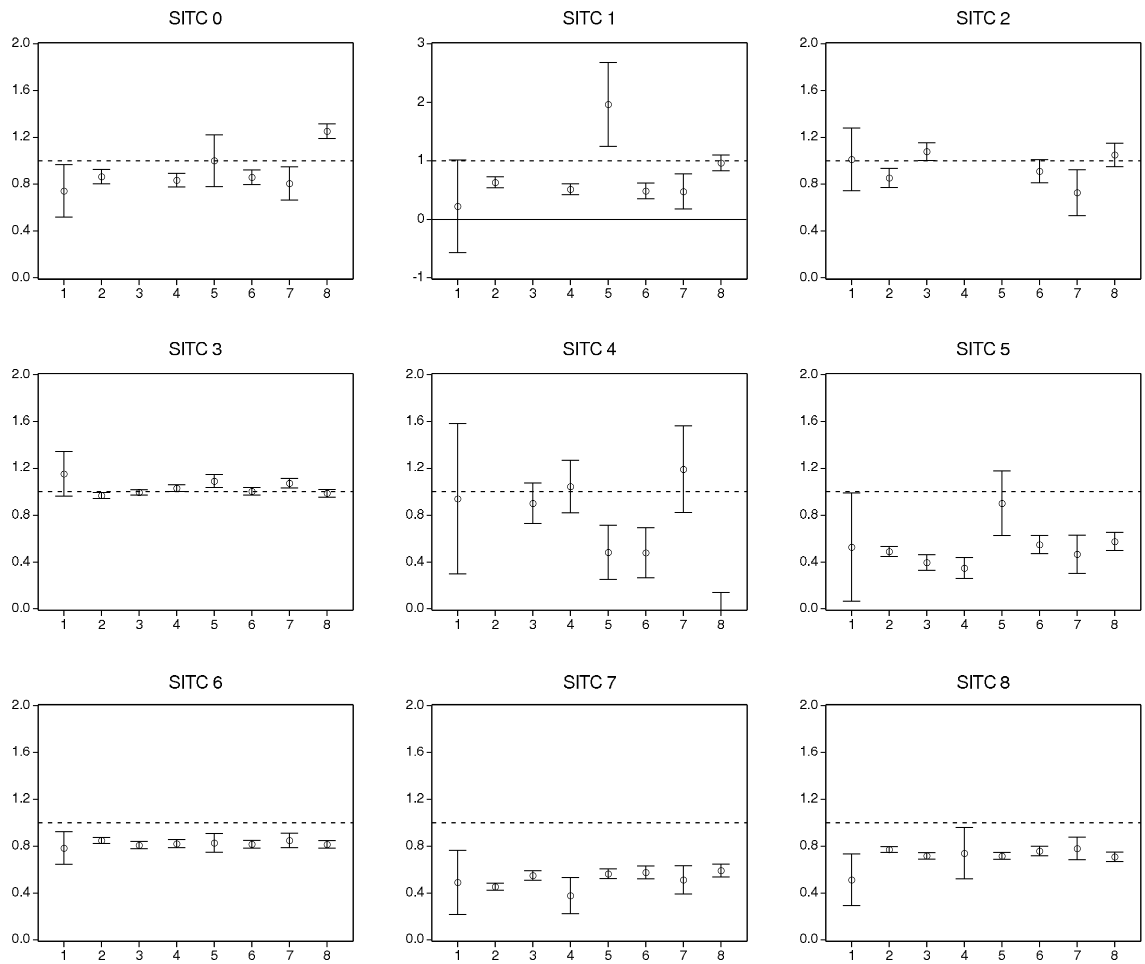

Figure 3: France - 'long-run' ERPT estimates with confidence intervals (95%). Individual industries, sample: 1995-2005. The estimators are presented in the following order: (1) CG long run, no cointegration, no break, (2) cointegrating long-run, no break, (3) cointegrating long run, break in constant (GH estimate), (4) cointegrating long run before break in slope (GH estimate), (5) cointegrating long run, after break (GH estimate), (6) cointegrating long run, break in constant (imposed in 1999m1), (7) cointegrating long run, before break in slope (imposed in 1999m1), (8) cointegrating long run, after break in slope (imposed in 1999m1). 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 from the authors. Dotted horizontal line at value of 1.

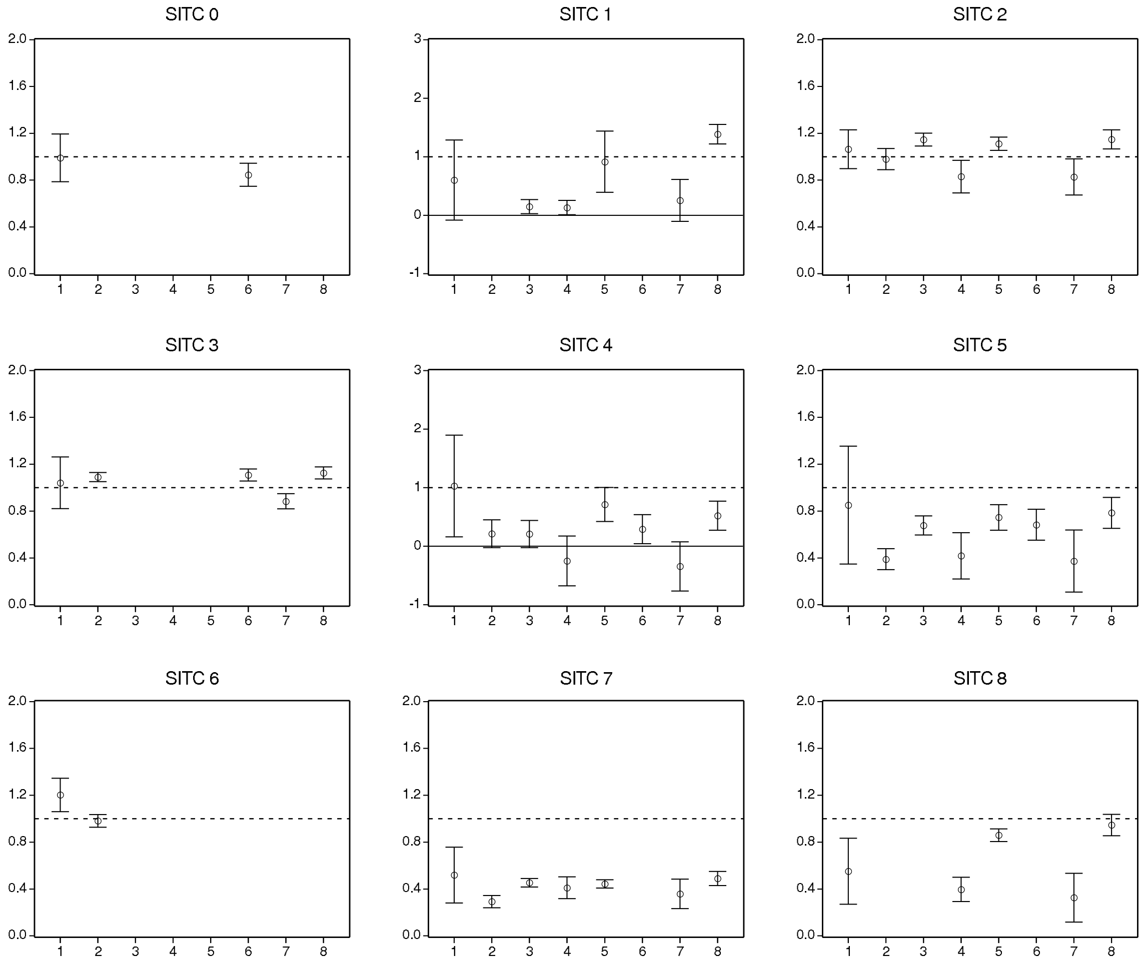

Figure 4: Italy - 'long-run' ERPT estimates with confidence intervals. Individual industries, sample: 1995-2005. Notes - see Figure (3) for explanations.

Table 4: Parametric Statistics for the Panel Cointegration Test. Sample: 1996-2004. See Appendix B for details.

Model	pseudo-t	pseudo-ρ
No break	-7.73	-35.45
Break in constant (eq. 12)	-22.15	-49.38
Break in constant and slope (eq. 13)	-23.26	-49.50

The null hypothesis is no cointegration.
Under the null hypothesis both statistics have
a N(0,1) distribution.
Full panel (N=90),
unit specific breaks, no cross-section dependence.

rejection of the hypothesis of no cointegration even when the alternative does not allow for a break. Since, as discussed above, we suspect the formation of the euro area constituted a shift in both competitive conditions and monetary policy which may have affected the long-run pass through, the (modified) Pedroni test statistics are computed allowing both for a break in the constant (equation (12)) and in the constant and slope (equation (13)). The results are reported in the second and third rows of Table 4 respectively. 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 5 and 6.

Within the context of the 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 6 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 slope and constant specification as being more coherent with the idea that the fixed component of the markup falls (a negative value of α_1) with reduced market power of "foreign currency denominated" products associated upon the introduction of EMU; while changes in the slope coefficient of the pass-through, observed for a number of sectors and countries, reflect the lower volatility of the exchange rate, as discussed below.[11] The estimated break dates for all the individual series are presented in Figure 5. There is some dispersion among the obtained dates, 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 5 and 6, in favor of both cointegration and structural change is unequivocally strong, a few qualifications are worth noting. First, the GH based algorithm here allows for only one, "strongest" break,[12] which is a serious

[11]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.

[12]In fact it does not touch upon the notion of the strength of evidence of the break. Generally the break

Table 5: 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 the Constant (equation (12)), Reported for Group Pseudo-t. Sample: 1996-2004.

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.
Cross-section: individual industry in an individual country. No cross-section dependence.

Table 5: Continued.

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 5: 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: Panel Cointegration (Pedroni, 1999, modified in Banerjee and Carrion-i-Silvestre, 2006, to account for breaks) Results from Specification as in Equation (13), Reported for Group Pseudo-t. Sample: 1996-2004.

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.
Cross-section: individual industry in an individual country. No cross-section dependence. Significance tests based on traditional t-stats.

Table 6: Continued.

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 6: 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.

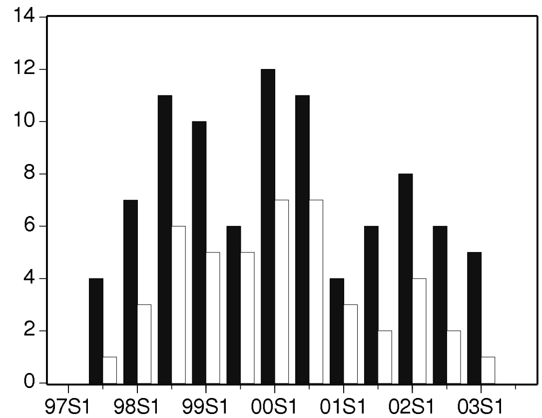

Figure 5: Distribution of estimated break dates in half-year intervals (1997s1-2003s2). Breaks in slope taken from Tables 6. Dark color - all breaks, light color - only breaks when long-run ERPT changed significantly (10%).

limitation as far as timing the (single) break allowed is concerned. Second, as noted earlier when referring to non-linear methods, the effect of the change in macroeconomic conditions on the ERPT may not have been either instantaneous or linear. Finally, there are other events in this period which may be 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 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 the estimated break points.

We consider the test results from the panel as sufficient evidence in favor of the existence of a long-run relationship between the variables, as implied by the theoretical underpinning - equation (9). Moreover, despite some variability in the estimated breaks in the individual series we can say that at least for some country/industry combinations there is evidence that the formation of the EMU led to a significant change in the equilibrium pass-through rate, be it directly upon its formation or indirectly by tying the currency to the euro, and thus seeing it appreciate against the dollar since about 2001. However, as given above, the failure of first generation panel cointegration tests to account for cross section dependence tends to oversize the tests and may lead to flawed inference on the existence of the long run relationship. Our final generalization of the testing framework, having already developed tests for cointegration with structural breaks, is to allow for a factor structure to model this type of dependence (as in Bai and Ng, 2004) and apply the test proposed by Banerjee and Carrion-i-Silvestre (2006) which allows for a unit specific estimated break in the series. In this second generation test, we test the null of no cointegration against an alternative hypothesis of cointegration (with up to r common factors modeling cross section dependence) with one common break date for all the series (1999m1) and for individual estimated break dates for each country. Due to the construction of the test, which is based on the extraction

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).

Table 7: Test Statistics for the Banerjee and Carrion-i-Silvestre (2006) Panel Co-integration Tests with Cross-section Dependence (Common Factors). Sample: 1996-2005.
H_0: unit root (no co-integration)

No. of factors	Pseudo-t - $ADF^c_{\hat{e}}(i)^*$				
	No break	Break (a)	Break (b)	Break (c)	Break (d)
(1)	-4.51	-3.72	-3.11	-4.51	-2.66
(3)	-3.71	-3.47	-2.60	-3.33	-2.59
(6)	-3.76	-2.95	-3.47	-4.44	-3.22

Under the null the statistics have the normal $N(0,1)$ distribution.
No break, (a) - common break in constant imposed in 1999m1,
(b) common break in the entire co-integrating relationship, imposed in 1999m1,
(c) - country-specific, estimated break as in equation (12),
(d) - country-specific, estimated break as in equation (13),
** - See Appendix B for description.*

of common factors from the error terms, the break and coefficient estimates are obtained with exactly the same first-generation tests conducted before and thus there is no need to repeat them. The results for 1-, 3- and 6-factor dependence structures are reported in Table 6 and we consider them as reconfirmation of the existence of a long-run or equilibrium relationship.

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 6 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 exchange rate regime, that occurred when these countries joined the euro area, which led foreign producers to expect more stable conditions, and as argued below 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 two years later, and in the run-up to the experienced a much more significant downward shift in inflation than the other countries. This may explain the particular pattern of the latter country.

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 the introduction of the euro for Italy, Greece and Spain. For the Netherlands, Ireland and Portugal the estimated breaks lie in mid 2002 and even in 2003. In the case of Ireland, an explanation for this may be provided by its 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. 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 a 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 that the pass-through in sector SITC3, 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 market power of foreign producers in this sector, who 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.

7 Discussion

Part of the literature, based primarily on US data (see for example Marazzi et al., 2005), has provided evidence for a fall in pass-through rates. This has been seen as support for Taylor (2000) who argued that a fall in the persistence of inflation leads to a fall in pass-through due to a decline in the pricing power of firms. Similarly we should expect ERPT to fall upon the introduction of the euro (see for instance Devereux, Engel and Tille, 2003, who contend 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 a significant rise in ERPT, especially for Italy and also Portugal and Spain, where the breaks coincide with the introduction of the euro. There may be several reasons for this finding.

First, the results referred to above concern 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, as indicated by Frankel, Parsley and Wei (2005), the effect of exchange rate volatility on the pass-through is often negative. 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 and volatile environment, producer-currency pricing may prove difficult. Faced with frequent and often temporary exchange rate changes, menu costs or costly pricing strategy reviews may lead to imported goods being 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 (2005) which generate high volatility of exchange rate associated with low ERPT.

Third, 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 single 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.

Fourth, the breaks in the vicinity of 2001 coincided 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 asymmetric 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 CG, 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. We treat the fact that in the cases when our estimated break point lies near 2001 the estimated ERPT rises, as support for the presence of an asymmetric ERPT.[13]

Our results are consistent with papers by Hellerstein et al. (2006) for the US and Campa and Goldberg (2006) for OECD member countries which challenge the idea of fall in ERPT as a robust finding. Hellerstein et al. (2006) mentions the role of intra-firm trade (that we cannot control for in our data base), as well as a commodity channel which creates a downward bias in the estimation of ERPT. Gopinath and Rigobon (2006) provide a careful analysis of ERPT using micro data on US export and import prices "at the dock" for the period 1994-2005. They find evidence of local currency pricing for US imports and producer currency pricing for US exports. The latter result should imply high levels of ERPT into import prices for US trading partners, in particular the European Union, which is consistent with our findings.

One final remark concerns the robustness of our methods in allowing for breaks in the long-run relationship. First, our models for both single-equation and panel methods allow for only a single break. This is a limitation imposed by the relatively short span of our sample. If there are, say, two breaks in the data, the algorithm may pick up only one of them, or estimate a break lying somewhere in between the two actual breaks. This may account for some of the heterogeneity reported in the tables. It should be emphasized however that the test for cointegration is robust to the presence of multiple structural breaks under the alternative as long as there are no breaks in trend (see Banerjee and Carrion-i-Silvestre, 2006). Moreover, the algorithm does not allow for a break under the null hypothesis, and the use of the infimum operator makes the test of the null hypothesis of no cointegration independent of the timing of the unknown break (under the alternative.) Thus, as in Gregory and Hansen's similar (but much simpler framework), consistent estimation

[13] Other suggested developments, arguably harder to date, include the increase in trade integration, ongoing trade liberalization, specific import composition of individual countries. Another development 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.

of the break date under the alternative is not needed to make consistent inference about cointegration (or no cointegration) under the null.

8 Conclusion

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.

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 unambiguous. Thus, even if one were not willing to accept the notion of 'detectable' structural change, as modeled 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.

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 introduction of the euro 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 it is more general. It provides estimates of the shift in constant that are more in line with the expected increase in competition arising out of trade integration and the introduction of the euro - i.e. the fixed component of the pass-through decreases while the variable component tends to increase.

Allowing for a structural break in the relationship we find that ERPT has generally increased in the time period near the introduction of the euro 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 asymmetric responses of the import prices. When we allow for a change in the long run relationship, we find that, towards the second part of our sample, i.e. 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.

- CG 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 CG data set (Belgium+Luxembourg excluded, Austria and Finland start 1996m1, Portugal and Austria stop 2004m12)

Appendix B - Descriptions of tests[14]

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 $x_{i,t}$, $i = 1, \dots, N$, $t = 1, \dots, T$. 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, $i = 1, \dots, N$. 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

[14]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 $(\hat{e}_{i,t}(\lambda_i))$:

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

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

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

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

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

The panel statistics on which they focus in order to test the null hypothesis are given by the $Z_{\hat{\rho}_{NT}}$ and $Z_{\hat{t}_{NT}}$ tests in Pedroni (1999, 2004), which can be thought as analogous to the residual-based tests in 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:

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

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

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

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

is the vector of estimated break fractions.

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

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

Panel cointegration with cross-sectional dependence: 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 (that is a break in the constant, as in equation (12) and a break in the constant and slope, as in equation (13)). 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 and we refer the reader to the original paper for details on the estimation strategy in the case of unknown, unit-specific breaks, which is used in our paper.
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:

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

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

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

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

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

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

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

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.

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