WORKING PAPER

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by

Antonia Arsova

University of Lüneburg
Working Paper Series in Economics

No. 384

February 2019

www.leuphana.de/institute/ivwl/publikationen/working-papers.html

ISSN 1860 - 5508

Exchange rate pass-through to import prices in Europe: A panel cointegration approach

Antonia Arsova*

Leuphana Universität Lüneburg

February 1, 2019

Abstract

This paper takes a panel cointegration approach to the estimation of short- and long-run exchange rate pass-through (ERPT) to import prices in the European countries. Although economic theory suggests a long-run relationship between import prices and exchange rate, in recent empirical studies its existence has either been overlooked, or it has proven difficult to establish. Resorting to novel tests for panel cointegration, we find support for the equilibrium relationship hypothesis. Exchange rate pass-through elasticities, estimated by two different techniques for cointegrated panel regressions, give insight into the most recent development of the ERPT.

Keywords: exchange rate pass-through, import prices, panel cointegration, cross-sectional dependence, common factors
JEL classification: C12, C23, F31

*E-mail: arsova@leuphana.de

1 Introduction

Exchange rate pass-through measures the extent to which import prices, expressed in the currency of the importing county, reflect changes in the exchange rate with its trading partners. Assuming that export prices are determined by a markup over marginal costs, the import price elasticity w.r.t. the exchange rate depends on the exporters' pricing strategies. If exporters choose to absorb exchange rate fluctuations into their markup, a strategy also known as local currency pricing (LCP) or pricing-to-market, then import prices remain largely unaffected by exchange rate shocks and the ERPT is said to be incomplete. On the other hand, if exporters choose not to adjust their markup, exchange rate fluctuations get reflected in full into import prices, which is known as producer currency pricing (PCP). The ERPT in this case is said to be complete. Under complete ERPT, depreciation of the importing country's currency translates into increase of import prices and may thus lead to inflation. Therefore, the degree and the determinants of exchange rate pass-through into import prices (a.k.a. first-stage pass-through) and subsequently into consumer prices (second-stage pass-through) are an important issue to policy-makers looking to stabilize inflation, especially in a monetary union such as the euro area.

In the ever growing body of empirical literature on ERPT one issue becomes apparent – namely, whether there exists a long-run equilibrium relationship between import prices, nominal exchange rate and other potential macroeconomic determinants of import prices. For example, Campa and Goldberg (2005), Ben Cheikh and Rault (2016) and Ben Cheikh and Rault (2017) find no or only weak evidence of cointegration and proceed to estimate an ERPT equation in first differences. De Bandt and Razafindrabe (2014) do not even consider the possibility of cointegrating relations and having established the nonstationarity of the model variables proceed to estimate a model in first differences as well.

On the other hand, De Bandt et al. (2008) and Brun-Aguerre et al. (2012) do establish a cointegrating relation and thus estimate error-correction (EC) models for the ERPT. However, De Bandt et al. (2008) allow for level shifts and structural breaks in the cointegrating relation, while Brun-Aguerre et al. (2012) employ individual-unit and first-generation panel cointegration tests, whose results might be compromised by unattended cross-sectional dependence. Delatte and López-Villavicencio (2012) and Brun-Aguerre et al. (2017) also find strong evidence for cointegration, but they focus on asymmetric ERPT – that is, allowing the effects of exchange rate appreciation or depreciation on import prices to differ. Consequently, they argue that imposing the restriction of symmetric ERPT may hinder revealing the long-run equilibrium. The evidence on the existence of a linear cointegration relationship is, therefore, inconclusive.

The presence or absence of cointegration determines the choice of estimation methodology and models which do not consider it have been criticized on two grounds. First, ignoring a significant error-correction (EC) term leads to omitting essential information and hence to inferior model performance (Brun-Aguerre et al., 2012). Second, by evading the notion of cointegration as long-run equilibrium, other ad-hoc measures of long-run ERPT need to be constructed, whose estimates strongly depend on the choice of other model parameters, e.g. the lag order, and can thus become unreliable (De Bandt et al., 2008). Therefore, the debate on whether cointegration underlies the ERPT not only constitutes an interesting econometric puzzle but has far-reaching consequences concerning the estimation results.

The contribution of this paper is twofold. First, employing novel second-generation panel cointegration tests it provides evidence on the existence of a long-run equilibrium relationship between import prices and nominal exchange rate for a panel of nineteen European countries. Contrary to some recent findings (e.g. De Bandt et al., 2008), cointegration emerges without the necessity to allow for structural breaks neither in the deterministic terms, nor in the cointegrating relation. The cointegrating relationship is shown to be driven by unobserved global stochastic trends. Second, taking cointegration and its driving forces into account, the paper presents estimates of the long-run and short-run pass-through elasticities at the panel level and for the individual countries using most recent data covering the period since the introduction of the Euro in 1999. The continuously updated fully modified (Cup-FM), and continuously updated bias-corrected (Cup-BC) estimators of Bai et al. (2009), and the dynamic common correlated effects (DCCE) estimator of Chudik and Pesaran (2015), all of which are robust to cross-sectional dependence induced by unobserved common factors, are employed. Despite the technical differences of these estimators, the results they yield are remarkably similar. Following a 1% depreciation of the exchange rate, the import prices are inclined to rise by 0.37% on average as estimated by the Cup-FM and Cup-BC estimators, and by 0.33% as estimated by the DCCE estimator. These results indicate only partial pass-through, rejecting both the LCP and PCP hypotheses for the panel as a whole.

The rest of the paper is organized as follows. Section 2 postulates the econometric model for the ERPT and describes the data used for the analysis. Section 3 presents the results of the unit root and cointegration analyses. Section 4 describes the econometric methodology for the estimation and discusses the empirical results, and Section 5 concludes. Auxiliary results are collected in the Appendix.

2 Model and data

2.1 Exchange rate pass-through into import prices

The analysis is based on the framework adopted by Campa and Goldberg (2005), which is commonly applied in the literature. For notational simplicity the model is written suppressing the dependence on the cross-sectional dimension i. It assumes that the import prices, P_t, equal the export prices of the country's trading partners, P_t^x, multiplied by the exchange rate, E_t, expressed per unit of foreign currency:

$$P_t = E_t P_t^x. \tag{1}$$

The export prices comprise the producers' marginal cost, C_t, and gross markup, M_t:

$$P_t^x = C_t M_t. \tag{2}$$

The marginal cost, in turn, depends on the wages in the exporting market, W_t, and on the demand conditions in the importing market, Y_t. Denoting the logarithms of all variables by lowercase letters, eq. (1) thus becomes

$$\begin{aligned} p_t &= e_t + c_t + m_t \qquad (3) \\ &= e_t + a_1 y_t + a_2 w_t + m_t. \end{aligned}$$

The markup is assumed to comprise both a fixed effect ϕ and a component depending the macroeconomic conditions, which may be reflected in the exchange rate and/or the demand conditions:

$$m_t = \phi + b_1 e_t + b_2 y_t. \tag{4}$$

Hence the general ERPT equation in log-linear form becomes

$$p_t = \phi + (1 + b_1)e_t + (a_1 + b_2)y_t + a_2 w_t, \tag{5}$$

or, more succinctly,

$$p_t = \beta_0 + \beta_1 e_t + \beta_2 y_t + \beta_3 w_t. \tag{6}$$

The primary focus of this paper is the pass-through elasticity given by the coefficient β_1 in eq. (6). If $\beta_1 = 1$, the pass-through to import prices is said to be complete. Exchange rate fluctuations are reflected one-to-one in the exporters' prices in the domestic market, and in this case producer currency pricing is present. If $\beta_1 = 0$, then exchange rate movements do not affect the prices in the importing market. Exporters do not adjust their prices abroad, but rather fully absorb the exchange rate fluctuations in their markup, and hence local currency pricing takes place.

2.2 Data description

The dataset comprises a balanced panel ($T = 113, N = 19$) with quarterly time series covering the period 1999Q1 – 2018Q1 for nineteen European countries: Austria, Belgium, Czech Republic, Denmark, Estonia, Finland, France, Germany, Italy, Lithuania, Luxembourg, Netherlands, Norway, Poland, Portugal, Spain, Sweden, Switzerland, and the United Kingdom.

The data on import prices are taken from the Main Economic Indicators (MEI) database of the OECD and reflect the prices of non-commodity imports of goods and services. Nominal effective exchange rate (NEER), weighted by the unit labour costs of a country's trading partners, is taken from the IMF International Financial Statistics (IFS) database for the model's exchange rate variable. It is defined in quantity notation such that an increase represents an appreciation of the domestic currency. This implies that the coefficient β_1 in (6) is expected to be negative, with $\beta_1 = -1$ indicating complete pass-through. Domestic demand is approximated by real GDP taken from the OECD Quarterly National Accounts database.

The choice of variable for the producers' costs is more involved, since there exists no directly observed variable which controls for the trade shares of the exporting countries. Therefore, a proxy for w has to be constructed from trade data. We follow Bailliu and Fujii (2004), who exploit the real effective exchange rate (REER) based on unit labour costs to create a trade-weighted measure of foreign producers' costs. Denoting the natural logarithm of REER by q, it can be represented as

$$q_t = e_t + ulc_t - ulc_t^*, \tag{7}$$

where ulc_t and ulc_t^* stand for the domestic and foreign unit labour costs in natural logarithms, respectively. REER is given in price notation, such that an increase reflects a worsening of the international competitive position, and e is given in quantity notation. Solving eq. (7) for ulc_t^* yields a trade weighted proxy for foreign producers' costs, which is then taken as w in the analysis. The unit labour costs series are obtained from the OECD MEI database, while REER and NEER are taken from IMF IFS.

3 Preliminary analysis

3.1 Testing for cross-sectional dependence

The first step of the analysis is to determine the degree and source of cross-sectional dependence in the panel. This is important in order to select the correct tools for analysing the integration and cointegration properties of the data and for the subsequent estimation of the ERPT. It is well-known that unattended strong cross-sectional dependence may result in oversized panel unit root and cointegration tests and biased estimates of the slope coefficients in eq. (6) (see, e.g., Banerjee et al., 2004 and Phillips and Sul, 2003, 2007).

For this aim the CD test of Pesaran (2015) is applied to the panel with country cross-sections for each variable in eq. (6). The test assumes weak[1] cross-sectional dependence under the null hypothesis, such as a spatial-type dependence or dependence driven by common factors affecting only a limited number of units as $N \to \infty$, for example. Rejection of the null is taken as evidence of the presence of strong cross-sectional dependence such as one caused by global (unobserved) common factors. The test statistic is computed as the standardized average of the pairwise correlation coefficients between the series in the panel and is normally distributed under the null hypothesis. To avoid spurious correlation arising from unit roots, the variables have been transformed into first differences. The results are presented in Table 1.

Table 1: Pesaran's (2015) CD statistic for the observed data

Variable	CD test statistic	p-value	$\overline{\hat{\rho}_{ij}}$	$\left\|\overline{\hat{\rho}_{ij}}\right\|$
Δp	42.72***	0.000	0.375	0.390
Δe	42.95***	0.000	0.377	0.483
Δy	52.45***	0.000	0.460	0.461
Δw	12.03***	0.000	0.106	0.172

Notes: $\overline{\hat{\rho}_{ij}}$ denotes the average pairwise correlation coefficient while $\left|\overline{\hat{\rho}_{ij}}\right|$ denotes the average absolute pairwise correlation coefficient over cross-sections.
*, ** and *** denote significance at the 10, 5 and 1% level, respectively.

The null of weak cross-sectional dependence is convincingly rejected for all variables. This is expected, given the tight economic and financial links between the European countries and the common currency and monetary policy in the euro area. Hence the analysis proceeds taking into account the presence of strong cross-sectional dependence.

3.2 Unit root and cointegration analysis

3.2.1 Unit root testing

Next the integration and cointegration properties of the time series are examined by second-generation panel unit root tests which are robust to cross-sectional dependence. In particular, the simple panel unit root test of Pesaran (2007) and the meta-analytic tests of Demetrescu et al. (2006) and Hanck (2013) are applied to the panel with country

[1]For definitions of notions of weak and strong cross-sectional dependence refer to Chudik et al. (2011).

cross-sections of each variable in eq. (6)[2]. Tables 8 and 9 in the Appendix summarize the results. The test of Pesaran (2007) rejects only for the exchange rate series at lags 1, 2, 3 and 4 and for the import price series at lag 1 (Table 9). On the other hand, a unit root at the panel level cannot be rejected for any variable in levels by the tests of Hanck (2013) and Demetrescu et al. (2006) (Table 8). All three tests reject the presence of a unit root in the first-differenced variables.[3] Hence there is prevailing evidence of the presence of unit roots in all variables in the model.

3.2.2 Cointegration testing

The next step in the analysis is to test the system of all four observed variables for cointegration. For this purpose the meta-analytic test of Arsova and Örsal (2019) is employed. Similarly to the panel unit root test of Hanck (2013), this test is too based on Simes' multiple testing procedure, where p-values from individual-unit likelihood-ratio (LR) cointegration rank tests of Saikkonen and Lutkepohl (2000) (SL) are used. Two versions of the latter test are considered, one allowing for a deterministic time trend both in the variables in levels and in the error-correction (EC) term, and one allowing for a trend only in the variables in levels only. Denoting the cointegrating rank of the system for country i by r_i, the null hypothesis of the test is $H_0 : r_i = r$, where $r = 0, 1, 2, 3$ denotes the common cointegrating rank in a sequential testing procedure. The alternative hypothesis is $H_1 : r_i > r$ for at least one i.

The results are presented in Table 2. As the smallest individual p-value for testing $H_0 : r = 0$ by the first variant of the SL test is lower than the corresponding Simes' critical value, while $H_0 : r = 1$ cannot be rejected, there is evidence of a single cointegrating relationship in the panel at the 5% significance level. In order to ensure that the long-run equilibrium connects not only a certain pair of variables, the test of Arsova and Örsal (2019) is applied to all eight different bi-variate systems. The null hypothesis of no cointegration is rejected for neither pair; the results are omitted for brevity. Hence the equilibrium relationship is more complex, involving at least three or all four of the variables in the system.

3.2.3 Analysis of the unobserved common and idiosyncratic components

Having established nonstationarity and the presence of a single long-run equilibrium relationship in the data, the analysis proceeds to uncover their driving forces. For this purpose the approach of panel analysis of nonstationarity in idiosyncratic and common components (PANIC) is employed, as set out in Bai and Ng (2004). The time series are decomposed into unobserved common and idiosyncratic components and their integration and cointegration properties are analyzed separately. The benefit of such analysis is that it provides better understanding of the interconnections among the variables in the system.

Unobserved dynamic common factors are extracted by the method of principal components from the panel for each variable with country cross-sections. Prior to the extraction the observed data are first-differenced. For the panels of import prices, GDP and

[2] More details on the computation of the tests by Hanck (2013) and Demetrescu et al. (2006) are given in the Appendix.

[3] The results of the tests by Hanck (2013) and Demetrescu et al. (2006) for the variables in first differences are omitted for brevity.

Table 2: Arsova and Örsal's (2019) intersection-type panel cointegration test

Trend in EC term				Trend orthogonal to EC term				Simes' crit. values	
Country	lag	$LR^{\mathrm{SL}}_{\mathrm{trace}}$	p-value	Country	lag	$LR^{\mathrm{SL}}_{\mathrm{trace}}$	p-value	5%	10%
				$H_0 : r = 0$					
Denmark	1	57.15	0.002**	Denmark	1	42.95	0.007	0.003	0.005
Poland	4	50.25	0.014	France	2	40.41	0.014	0.005	0.011
Sweden	2	49.50	0.017	Lithuania	2	39.00	0.021	0.008	0.016
France	2	44.62	0.059	Czech Republic	2	37.37	0.033	0.011	0.021
Lithuania	2	43.10	0.083	Estonia	4	34.92	0.062	0.013	0.026
Luxembourg	1	42.66	0.092	Sweden	2	34.08	0.077	0.016	0.032
Czech Republic	2	41.85	0.109	Portugal	1	33.36	0.091	0.018	0.037
Austria	2	41.77	0.111	Netherlands	2	32.45	0.111	0.021	0.042
Estonia	4	41.76	0.111	Germany	2	32.39	0.113	0.024	0.047
Germany	2	40.76	0.136	United Kingdom	2	31.62	0.133	0.026	0.053
United Kingdom	3	39.69	0.167	Spain	2	31.55	0.135	0.029	0.058
Norway	1	39.26	0.181	Belgium	3	30.72	0.161	0.032	0.063
Netherlands	2	39.23	0.183	Italy	2	30.43	0.171	0.034	0.068
Italy	2	38.86	0.195	Luxembourg	2	29.20	0.217	0.037	0.074
Switzerland	3	36.25	0.303	Finland	3	28.60	0.243	0.039	0.079
Portugal	1	35.13	0.358	Poland	4	28.52	0.247	0.042	0.084
Belgium	2	34.43	0.394	Austria	2	27.91	0.275	0.045	0.089
Spain	2	33.29	0.457	Switzerland	3	27.60	0.290	0.047	0.095
Finland	3	32.31	0.513	Norway	2	22.90	0.568	0.050	0.100
				$H_0 : r = 1$					
Czech Republic	2	27.73	0.063	Czech Republic	3	24.86	0.014	0.003	0.005
Netherlands	2	25.47	0.117	Lithuania	2	24.17	0.018	0.005	0.011
United Kingdom	2	24.62	0.146	United Kingdom	2	23.31	0.024	0.008	0.016
Spain	3	24.08	0.166	Luxembourg	1	18.94	0.093	0.011	0.021
Denmark	1	23.33	0.199	Switzerland	2	18.20	0.116	0.013	0.026
Sweden	2	23.00	0.214	Denmark	1	15.99	0.208	0.016	0.032
Germany	2	21.08	0.321	France	3	15.00	0.264	0.018	0.037
Italy	2	19.31	0.442	Germany	4	14.08	0.325	0.021	0.042
Switzerland	3	19.20	0.451	Spain	3	13.77	0.348	0.024	0.047
Luxembourg	3	18.53	0.501	Belgium	2	13.53	0.366	0.026	0.053
Estonia	4	17.79	0.558	Portugal	1	13.43	0.374	0.029	0.058
Portugal	1	16.69	0.643	Poland	4	13.37	0.378	0.032	0.063
Finland	4	15.75	0.714	Austria	3	12.49	0.450	0.034	0.068
Belgium	2	15.35	0.742	Finland	2	12.09	0.484	0.037	0.074
Austria	3	14.80	0.780	Italy	2	11.78	0.511	0.039	0.079
Poland	4	13.64	0.850	Netherlands	2	10.91	0.590	0.042	0.084
Norway	1	12.15	0.920	Estonia	2	9.06	0.754	0.045	0.089
Lithuania	4	11.72	0.935	Sweden	2	8.46	0.802	0.047	0.095
France	3	11.53	0.941	Norway	1	8.08	0.830	0.050	0.100

Notes: The lag order is selected according to the modified AIC criterion of Qu and Perron (2007). Results for each variable are sorted according to the p-values in ascending order for ease of comparison with the corresponding Simes' critical value.
*, ** and *** denote significance of the panel intersection test at the 10, 5 and 1% level, respectively.

producer's costs they are also demeaned to account for the observed time trend. The data are also standardized to have unit variance. The number of unobserved common factors for each panel is selected by the criterion of Onatski (2010); the maximum number allowed is six. The criterion picks two factors for the panels of p and e, explaining 57% and 75% of the variation in the data, respectively. For each of the panels of y and w a single factor is chosen, explaining 52% and 23% of the total variation, respectively.

Once the estimated variable-specific common factors are extracted and subtracted from the first-differenced (and potentially demeaned) observations, the remaining residuals are accumulated to yield estimates $\hat{e}^x_{i,t}$ of the idiosyncratic components for each variable and cross-sectional unit, $x \in \{p, e, y, w\}$. The estimated idiosyncratic components are then tested for unit roots by the P_a, P_b, and $PMSB$ tests proposed by Bai and Ng (2010). Table 3 presents the results. The null hypothesis of a unit root cannot be rejected for either panel.

Table 3: Bai and Ng's (2010) panel unit root tests for the estimated idiosyncratic components

Idiosyncratic component	Avg. volatility	P_a	P_b	$PMSB$
$\hat{e}^p$	0.037	−0.005	−0.005	0.027
$\hat{e}^e$	0.058	1.221	1.825	2.893
$\hat{e}^y$	0.034	−0.758	−0.682	−0.585
$\hat{e}^w$	0.035	0.945	1.092	1.266

Notes: Trend is included in the test regressions for p, y and w, while only a constant is considered for e. All three test statistics have a $N(0,1)$ distribution under the null hypothesis of a unit root with a rejection region in the left tail of the distribution. The average volatility is computed over all cross-sections.

Next, the cointegration properties of the idiosyncratic components are examined by the PSL^J_{def} test of Arsova and Örsal (2018) and the $P^*_{\Phi^{-1}}$ test of Örsal and Arsova (2017). The first test computes the panel test statistic as the standardized average of the individual LR trace statistics of Saikkonen and Lutkepohl (2000) computed from defactored data, while the second one combines the p-values of these statistics by the inverse normal method. Both statistics have a limiting $N(0,1)$ distribution under the null hypothesis of a common cointegrating rank $H_0 : r_i = r, \forall i$, whereas the rejection region for the panel-SL test is in the right tail, and for the $P^*_{\Phi^{-1}}$ test in the left tail, respectively. Örsal and Arsova (2017) show that the $P^*_{\Phi^{-1}}$ exhibits better finite-sample properties than the panel-SL test in some situations. The value of the PSL^J_{def} test statistic under the null of no cointegration is 0.2, while that of the $P^*_{\Phi^{-1}}$ is −1.33, which is significant at the 10% level. As neither test rejects the null of cointegrating rank one ($PSL^J_{def} = -2.13$ and $P^*_{\Phi^{-1}} = 4.99$ in this case), we conclude that there is some, albeit not very strong, evidence of a single cointegrating relationship among the idiosyncratic components, matching the result for the observed variables.

We next turn our attention to the extracted and accumulated common factors. They are denoted as $F^p_1, F^p_2, F^e_1, F^e_2, F^y$, and F^w, with the superscript signifying the variable-specific panel they have been extracted from and the subscript denoting the factor number. A graph of the factors is displayed in Figure 1. It reveals how they all capture the effects of the Global Financial Crisis, reacting mostly simultaneously and with similar turns in the dynamics. Such behaviour hints at possible cointegration among them, which could lead to cross-unit cointegration of the observed variables.

Testing for unit roots in the extracted common factors is carried out by a standard ADF test[4]. The results, presented in Table 4, indicate that the unit root null hypothesis cannot be rejected for any individual factor. The test statistic of the modified inverse normal panel test of Demetrescu et al. (2006) is −0.504, supporting this conclusion.

It is interesting to note the enormous difference between the volatility of the estimated common factors and that of the idiosyncratic components, displayed in Tables 3 and 4, respectively. Even the smallest volatility among those of the factors (0.52 for $\hat{F}^p_2$) is about ten times greater than the largest average volatility of the idiosyncratic components (0.058 for $\hat{e}^e$). Hence we conclude that it is the unobserved common factors

[4]Bai and Ng (2004) show that the limiting distributions of the ADF test statistics, computed for common factors extracted from first-differenced or first-differenced and demeaned data, coincide with the usual limiting distributions of the ADF test with a constant only or a constant and linear time trend, respectively.

Figure 1: Extracted common factors from the panel of each model variable

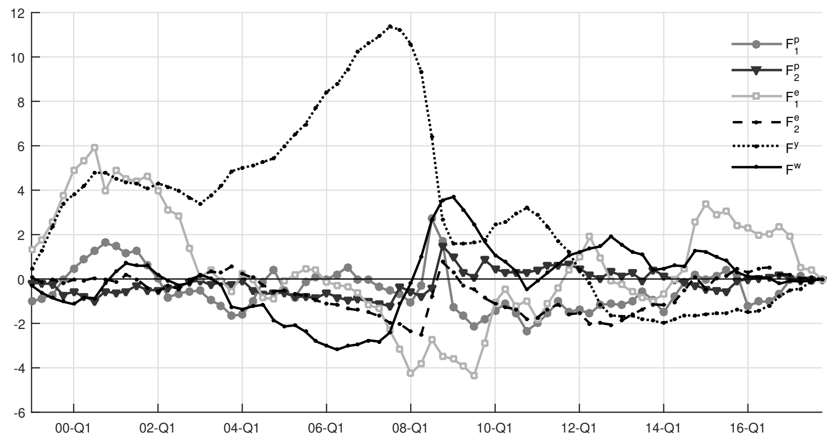

Table 4: Unit root tests for the estimated common factors

Factor	Volatility	deterministic term	lag order	ADF_τ	p-value
F_1^p	0.97	trend	2	-3.05	0.125
F_2^p	0.52	trend	4	-2.67	0.254
F^y	2.42	trend	1	-1.89	0.651
F^w	0.83	trend	2	-2.13	0.522
F_1^e	3.80	const	2	-2.17	0.496
F_2^e	1.53	const	1	-3.03	0.131

Notes: ADF_τ denotes the Augmented Dickey-Fuller test statistic. The lag order is selected according to the modified AIC (MAIC) criterion of Ng and Perron (2001). The p-values are computed as in MacKinnon (1996); the author is grateful to Christoph Hanck for providing the GAUSS code.

which to a large extent determine the behaviour of the observed variables, while the idiosyncratic components have only a minor impact.

Having established the presence of global stochastic trends, we next assess whether they exhibit any cointegration. For more reliable results the SL test of Saikkonen and Lutkepohl (2000) is employed for each pair of estimated factors, as it is known that the LR cointegrating rank tests become less powerful in larger systems (see, e.g. Saikkonen and Lutkepohl (2000)). Table 5 displays the results.

At first glance there seems to exist a cointegrating relationship between almost any pair of factors considered when allowing for no trend in the cointegrating relation. However, the results of these tests are highly correlated, and one must take the nature of such multiple testing into account. In order to select only the meaningful rejections, Hommel (1988) proposes a procedure which controls the family-wise error rate at a chosen significance level α. Details on Hommel's procedure can be found in Hanck (2013), whose exposition is briefly reproduced here for convenience. Let the ordered p-values of n tests be $p^*_{(1)} \leq \ldots \leq p^*_{(n)}$ and $\mathbb{N}_n$ denote the set of all natural numbers between 1 and n. Selecting the meaningful rejections by the Hommel's procedure is then carried out in two steps: (A) Compute $j = \max\{i \in \mathbb{N}_n : p^*_{(n-i+k)} > \frac{k\alpha}{i}, \forall k \in \mathbb{N}_i\}$, and (B) If $p^*_{(n)} \leq \alpha$, reject all $H_{i,0}$; else, reject those $H_{i,0}$ for which $p^*_i \leq \frac{\alpha}{j}$.

Following this procedure, $j = 10$ is computed, and the corresponding Hommel's critical values at the 5% and 10% significance levels are 0.005 and 0.01, respectively. Hence, only the first five rows in the second panel of Table 5 can be considered genuine rejections at the 10%-level; at the 5%-level it would only be the first three. We may

Table 5: SL cointegration tests for the estimated common factors

Trend in EC term				Trend orthogonal to EC term			
Factors	Lag order	$LR^{\text{SL}}_{\text{trace}}$	p-value	Factors	Lag order	$LR^{\text{SL}}_{\text{trace}}$	p-value
F^p_1, F^p_2	2	20.81	0.006***	F^p_2, F^w	4	19.59	0.001***
F^p_2, F^w	4	18.01	0.020**	F^p_1, F^w	3	16.81	0.002***
F^p_1, F^w	3	16.90	0.032**	F^p_1, F^p_2	2	16.60	0.003***
F^e_2, F^w	3	13.75	0.105	F^y, F^w	4	14.53	0.007***
F^e_1, F^y	3	12.52	0.159	F^e_2, F^w	3	13.69	0.010***
F^p_2, F^e_1	4	12.45	0.163	F^e_1, F^w	2	12.88	0.014**
F^e_1, F^w	2	12.43	0.163	F^e_1, F^y	3	12.50	0.016**
F^p_1, F^e_1	3	11.73	0.205	F^P_1, F^y	3	11.91	0.021**
F^p_2, F^y	4	11.15	0.244	F^p_2, F^e_1	4	11.59	0.024**
F^p_1, F^e_2	2	10.63	0.284	F^p_2, F^y	4	11.45	0.026**
F^P_1, F^y	3	10.12	0.327	F^p_1, F^e_2	2	10.82	0.034**
F^y, F^w	4	9.96	0.342	F^p_1, F^e_1	3	10.03	0.048**
F^e_2, F^y	3	7.76	0.574	F^e_2, F^y	3	7.78	0.120
F^p_2, F^e_2	4	5.99	0.776	F^p_2, F^e_2	4	6.34	0.208
F^e_1, F^e_2	1	5.14	0.858	F^e_1, F^e_2	1	5.50	0.282

Notes: The lag order is selected according to the modified AIC criterion of Qu and Perron (2007). Results for each variable are sorted according to the p-values in ascending order for ease of comparison with the corresponding critical value of Hommel's (1988) procedure.
*, ** and *** denote significance at the 10, 5 and 1% level, respectively.

therefore conclude that two global stochastic trends exist among the extracted common factors: one which is shared by F^p_1, F^p_2, F^y, F^w and F^e_2, and one driving F^e_1.

By analyzing the factor loadings (see Table 10 in the Appendix), F^e_1 may be identified as the Euro-exchange-rate factor, which is perhaps not surprising, as the dataset features both countries in and outside the euro area. On the other hand, F^e_2 can be thought of the factor influencing more the dynamics of the exchange rates of the non-euro area countries (including the newest members of the euro area like Lithuania, for example). Relating these results to those from the cointegration testing of the observed variables (Table 2), we conclude that there is much more evidence in favour of a long-run equilibrium relationship in the ERPT for non-euro area countries than it is for euro area ones. One explanation for this phenomenon may lie in the fact that the import prices in euro area countries, whose principal share of imports come from other euro area countries, react much less to aggregate exchange rate fluctuations because these are basically zero between the one and the same currency. This leads us to believe that the ERPT estimates would be lower for the older member-countries of the euro area than they would be for the newer ones or the countries outside the euro area.

The results of the unit root and cointegration analysis can be summarized as follows. All variables in the log-linear ERPT relationship in eq. (6) are integrated of order one. There is evidence of a single cointegrating relationship at the panel level linking the observed variables, suggesting that the average long-run elasticity of the exchange rate is different from zero. It is worth noting that, contrary to the results of De Bandt et al. (2008), this relationship emerges without the necessity to consider structural breaks, neither in the deterministic components, nor in the long-run equilibrium. This is so because of the present cross-unit cointegration driven by unobserved common factors. These factors capture the major exogenous shocks such as the Global Financial Crisis which, in turn, force the observed variables to react more or less simultaneously and in a similar fashion. Although the data dynamics are mostly determined by six unobserved common factors (two for the panel of import prices, two for the panel of nominal exchange rate and one for each of the domestic demand and producer's cost proxy panels), the driving forces behind them are only two distinct global stochastic trends. One of them is shared by the import prices panel, the doemestic demand panel, the producer's costs panel, and by the exchange rate data for countries outside the euro area as well as newer member-countries of the euro area. The second global stochastic trend can be viewed as a Euro-nominal-exchange-rate factor, influencing mostly the exchange rate series of the euro area countries. The idiosyncratic components of the data, although with much less impact than the common components, are also non-stationary and cointegrated by a single relationship. These findings lead us to expect more significant ERPT elasticities for non-euro area countries than for euro area ones.

4 ERPT estimation

Having established the presence of a long-run equilibrium relationship at the panel level, the next step is to estimate the ERPT equation (6). However, the presence of cross-sectional dependence, depending on its nature, may yield the results of earlier panel regression estimators either biased, inconsistent or inefficient (see, e.g. Phillips and Sul (2003, 2007) and Moon and Weidner (2017)). Further, cross-unit cointegration has also been shown by Urbain and Westerlund (2006) to pose an issue in pooled ordinary least squares estimation. Hence an estimator which takes into account both cointegration and cross-sectional dependence induced by global stochastic trends is needed.

Two suitable approaches have recently been proposed in the literature. The first one, put forward by Bai et al. (2009), features two estimators: the continuously-updated bias-corrected (Cup-BC) and the continuously-updated fully-modified (Cup-FM) estimator. They estimate level relationships in panel cointegration models where unobserved common factors drive the dependence in the regression errors and which may also be correlated with the regressors. This methodology has been widely applied in the recent empirical panel data studies, employed by e.g. Bodart et al. (2015) for estimation of a long-run relationship between real exchange rates and commodity prices, and by Örsal (2017) for estimation of a long-run money demand relation.

The second approach, using a common correlated effects (CCE) mean-group (MG) estimator, is due to Chudik and Pesaran (2015). They extend earlier work of Pesaran (2006) to panel data models allowing for lagged dependent variables and weakly exogenous regressors. The residual dependence induced by the unobserved common factors is captured by cross-sectional averages of the observed variables included as additional regressors in the individual equations. Details on the estimation by each estimator are briefly outlined next, while the empirical results are discussed in Section 4.3.

4.1 The Cup-BC and Cup-FM estimators of Bai et al. (2009)

The ERPT equation (6) can be written in the Bai et al. (2009) estimation framework as

$$p_{it} = \beta_0 + \beta_1 e_{it} + \beta_2 y_{it} + \beta_3 w_{it} + u_{it}, \tag{8}$$

$$u_{it} = \lambda_i' f_t + \varepsilon_{it}. \tag{9}$$

The errors ε_{it} are assumed to be stationary and only weakly cross-sectionally dependent, while the unobserved common factors in the $(r \times 1)$-vector f_t are allowed to be $I(0)$, $I(1)$ or a mixture of the two. They are treated as parameters and estimated together with the common slope coefficients $\beta = (\beta_1, \beta_2, \beta_2)$ in an iterative procedure. Albeit consistent, the resulting $\hat{\beta}_{\text{Cup}}$ estimator has been shown to be asymptotically biased, hence a bias correction is necessary. The Cup-BC and the Cup-FM estimators differ with regard to when this bias correction takes place. With the Cup-BC it is applied only once at the final iteration, while with the Cup-FM the correction is made at each iteration. $\hat{\beta}_{\text{Cup}}$ is shown to be at least T-consistent regardless of the integration order (zero or one) of the factors or that of the regressors. Being pure panel estimators, however, both the Cup-FM and the Cup-BC assume homogeneity of the coefficients across cross-sections, and hence do not produce individual-unit results, which may be viewed as a drawback in practice.

In order to account for the trending behaviour of the variables p, y and w, eq. (9) is estimated with demeaned and detrended series, as suggested by Bai et al. (2009). The number r of residual common factors is selected by the criterion of Onatski (2010). It picks two factors which account for 59% of the variance of the first-stage residuals $\hat{u}_{it}$.

The results are presented in Table 6. The actual estimates of the elasticity parameters are quite similar across the two estimators. The nominal exchange rate elasticity, which is the only statistically significant coefficient, is estimated by both the Cup-FM and the Cup-BC as $\hat{\beta}_1 = -0.37$. This implies that a 1% depreciation in the exchange rate would lead to an average increase of 0.37% in the import prices. The results are discussed in more detail in Section 4.3.

Table 6: ERPT estimation results by the Cup-BC and Cup-FM estimators

Variable	Cup-BC	Cup-FM
Nominal exchange rate elasticity $\hat{\beta}_1$	-0.372^{***}	-0.369^{***}
	[0.029]	[0.029]
Domestic demand elasticity $\hat{\beta}_2$	-0.041	-0.009
	[0.044]	[0.042]
Producers' costs elasticity $\hat{\beta}_3$	-0.011	0.007
	[0.035]	[0.035]

Notes: Standard errors are presented in brackets.
*, ** and *** denote significance at the 10, 5 and 1% level, respectively.

Figure 2 in the Appendix presents a graph of the estimated residual common factors[5]. The Global Financial Crisis manifests itself in the two spikes in 2008Q4 and 2009Q1, respectively. Analysis of the Cup-FM[6] model residuals, depicted in Figure 3 in the Appendix, reveals that the two factors adequately capture the effects of the crisis, as no further common shocks can be observed. Applying Demetrescu et al.'s (2006) panel unit root test to the estimated residuals yields a value of -5.25 for Hartung's test statistic with $\kappa = 0.2$, which points to their stationarity. This leads us to the conclusion that no model assumptions have been violated.

4.2 The dynamic CCE estimator of Chudik and Pesaran (2015)

The second approach considered for the estimation of the ERPT equation (6) is the panel autoregressive distributed-lag (ARDL) model with multifactor error structure proposed by Chudik and Pesaran (2015). This framework differs from the specification of Bai et al. (2009) in that it allows for (a) lagged values of the dependent

[5]These factors, common to the first-stage residuals of the Cup-FM and Cup-BC models, are not to be confused with the common factors extracted from the panel formed by each variable.

[6]Results for the residuals of the CUP-BC estimation are very similar and omitted for brevity.

and independent variables as additional regressors and (b) heterogeneous coefficients $\beta_{j,i}, (i = 1, \dots, N,\ j = 0, \dots, 3)$, for each unit, which are then combined in a mean-group (MG) estimator. As in Bai's framework, unobserved common factors in the residuals drive the strong cross-sectional dependence. The common factors, however, are not explicitly estimated from the data. Instead, they are approximated by cross-sectional averages of the observed model variables and the resulting regressions are estimated individually for each unit by ordinary least squares. A necessary condition for the validity of the resulting CCE MG estimator is that the number of unobserved common factors be no more than the observed variables in the system. This assumption is likely to be satisfied in our case, as Onatski's (2010) criterion picks two factors in the residuals of the panel regression in eq. (9). Initially proposed for stationary factors (Pesaran, 2006), the CCE approach has been proved to be valid for integrated factors as well (Kapetanios et al., 2011). With regard to the assumed weak exogeneity of the regressors, also necessary for the validity of the DCCE MG estimator, we note that the preceding analysis is valid upon the assumption that changes in the import prices do not contemporaneously affect exchange rates, domestic demand or producers' costs. This assumption is commonly made in the empirical ERPT literature and is not a restrictive one, given the quarterly frequency of the data; for a more detailed discussion we refer to Brun-Aguerre et al. (2017).

To cast the ERPT model (6) into the CCE-framework, we begin with eq. (9), allowing for heterogeneous coefficients[7]:

$$p_{it} = \beta_{0,i} + \beta_{1,i} e_{it} + \beta_{2,i} y_{it} + \beta_{3,i} w_{it} + u_{it}, \tag{10}$$

$$u_{it} = \lambda_i' f_t + \varepsilon_{it}. \tag{11}$$

Taking the cointegrating relationship explicitly into account, we then put it in an error-correction form:

$$\begin{aligned} \Delta p_{it} = {} & \beta_{0,i} + \rho_i \left(p_{i,t-1} - \beta_{1,i} e_{i,t-1} - \beta_{2,i} y_{i,t-1} - \beta_{3,i-1} w_{it} - \lambda_i' f_{t-1} \right) \\ & + \gamma_{1,i} \Delta e_{it} + \gamma_{2,i} \Delta y_{it} + \gamma_{3,i} \Delta w_{it}^x + \gamma_{i,f} \Delta f_t + \varepsilon_{it}, \end{aligned} \tag{12}$$

so that, by re-arranging, we get

$$\begin{aligned} \Delta p_{it} = {} & \pi_{0,i} + \pi_{ec,i} p_{i,t-1} + \pi_{1,i} e_{i,t-1} + \pi_{2,i} y_{i,t-1} + \pi_{3,i} w_{i,t-1} + \pi_{f,i} f_{t-1} \\ & + \pi_{4,i} \Delta e_{it} + \pi_{5,i} \Delta y_{it} + \pi_{6,i} \Delta w_{it} + \varepsilon_{it}. \end{aligned} \tag{13}$$

The long-run parameters $\beta_{j,i}, (i = 1, \dots, N,\ j = 1, 2, 3)$, can be recovered from the coefficients of eq. (13) as $\beta_{j,i} = -\pi_{j,i}/\pi_{ec,i}$, while the short-run parameters $\pi_{j,i}, j = 4, 5, 6$, are estimated directly. The term $\pi_{ec,i}$ describes the speed of adjustment to equilibrium, and its statistical significance may be viewed as an additional evidence of the presence of cointegration.

For the estimation, the unobserved common factors f_t in eq. (13) are replaced by cross-sectional averages of the observed variables:

$$\begin{aligned} \Delta p_{it} = {} & \pi_{0,i} + \pi_{ec,i} p_{i,t-1} + \pi_{1,i} e_{i,t-1} + \pi_{2,i} y_{i,t-1} + \pi_{3,i} w_{i,t-1} \\ & + \pi_{4,i} \Delta e_{it} + \pi_{5,i} \Delta y_{it} + \pi_{6,i} \Delta w_{it} \\ & + \pi_{1,i}^* \Delta \bar{p}_t + \pi_{2,i}^* \bar{p}_{t-1} + \pi_{3,i}^* \bar{e}_{t-1} + \pi_{4,i}^* \bar{y}_{t-1} + \pi_{5,i}^* \bar{w}_{t-1} \\ & + \pi_{6,i}^* \Delta \bar{e}_t + \pi_{7,i}^* \Delta \bar{y}_t + \pi_{8,i}^* \Delta \bar{w}_t + \varepsilon_{it}. \end{aligned} \tag{14}$$

[7]The exposition is similar to that of Eberhardt and Presbitero (2015), whose Stata code for the dynamic CCE MG estimator has been used for the estimation.

So far, eq. (14) constitutes the model for the standard CCE MG estimator of Pesaran (2006). As Chudik and Pesaran (2015) show, finite-sample bias arises in the dynamic panel model with weakly exogenous regressors and recommend the inclusion of sufficient number (s) lagged values of the cross-sectional averages to mitigate it. Their suggested rule of thumb, $s = \text{int}(T^{1/3})$, gives $\hat{s} = 4$ in our case. Hence the complete model to be estimated by the dynamic CCE (DCCE) estimator becomes

$$\begin{aligned}\Delta p_{it} = \pi_{0,i} &+ \pi_{ec,i} p_{i,t-1} + \pi_{1,i} e_{i,t-1} + \pi_{2,i} y_{i,t-1} + \pi_{3,i} w_{i,t-1} \\ &+ \pi_{4,i} \Delta e_{it} + \pi_{5,i} \Delta y_{it} + \pi_{6,i} \Delta w_{it} \\ &+ \pi^*_{1,i} \Delta \bar{p}_t + \pi^*_{2,i} \bar{p}_{t-1} + \pi^*_{3,i} \bar{e}_{t-1} + \pi^*_{4,i} \bar{y}_{t-1} + \pi^*_{5,i} \bar{w}_{t-1} \\ &+ \pi^*_{6,i} \Delta \bar{e}_t + \pi^*_{7,i} \Delta \bar{y}_t + \pi^*_{8,i} \Delta \bar{w}_t \\ &+ \sum_{l=2}^{4} \pi^*_{9,l,i} \Delta \bar{p}_{t-l} + \sum_{l=1}^{4} \pi^*_{10,l,i} \Delta \bar{e}_{t-l} \\ &+ \sum_{l=1}^{4} \pi^*_{11,l,i} \Delta \bar{y}_{t-l} + \sum_{l=1}^{4} \pi^*_{12,l,i} \Delta \bar{w}_{t-l} + \varepsilon_{it}.\end{aligned} \tag{15}$$

The results from the estimation at the panel level are listed in Table 7, while results from the individual country models are available in Table 11 in the Appendix.

Table 7: ERPT estimation results by the CCE MG estimator

Variable		Standard MG estimator		DCCE estimator		Augmented DCCE estimator	
		const	trend	const	trend	const	trend
e	LRA	−0.473***	−0.479***	−0.420***	−0.315***	−0.441***	−0.325***
		[0.069]	0.082	[0.121]	[0.095]	[0.154]	[0.131]
	SRA	−0.363***	−0.370***	−0.305***	−0.221**	−0.240***	−0.187*
		[0.073]	0.062	[0.071]	[0.104]	[0.081]	[0.106]
y	LRA	0.473***	0.337***	−0.063	0.083	−0.003	0.112
		[0.089]	0.089	[0.115]	[0.170]	[0.117]	[0.158]
	SRA	−0.074	−0.163	0.000	0.105	0.069	0.106
		[0.138]	0.137	[0.131]	[0.131]	[0.091]	[0.116]
w	LRA	−0.124	−0.163***	0.026	0.132*	0.031	0.056
		[0.080]	0.044	[0.065]	[0.072]	[0.092]	[0.094]
	SRA	0.116**	0.092*	0.022	0.028	0.029	0.024
		[0.047]	0.049	[0.070]	[0.071]	[0.067]	[0.068]
avg. EC coefficient $\bar{\rho}$		−0.363***	−0.428***	−0.650***	−0.729***	−0.666***	−0.748***
		[0.035]	0.035	[0.065]	[0.064]	[0.052]	[0.045]
Implied half-life		1.54	1.24	0.66	0.53	0.63	0.50
RMSE		0.022	0.021	0.012	0.012	0.010	0.010
CD test		34.87	34.13	4.50	3.12	1.90	0.52
Trends share			0.47		0.42		0.37

Notes: Standard MG estimator refers to the MG estimator of Pesaran and Smith (1995). Augmented DCCE estimator denotes the DCCE estimator augmented with a dummy variable accounting for the Great Recession in 2008. LRA and SRA denote long-run average and short-run average coefficients, respectively. Standard errors are presented in brackets. Implied half-life is computed as $\ln(0.5)/\ln(1+\bar{\rho})$. CD test denotes Pesaran's (2015) test for weak cross-sectional dependence of the regression residuals; it has $N(0,1)$ distribution under the null hypothesis. Trends share stands for the share of group-specific trends significant at the 5% level.
*, ** and *** denote significance at the 10, 5 and 1% level, respectively.

The first two columns present the results from the estimation with the standard mean-group panel estimator of Pesaran and Smith (1995), included to illustrate what the effect of unattended cross-secitonal dependence would be. The residual correlation as measured by Pesaran's (2015) CD test is highly significant and the resulting estimates are therefore likely to be biased and inconsistent (Moon and Weidner, 2017).

The middle two columns of Table 7 present the results from the DCCE estimator with a constant or a constant and a linear time trend, respectively. The CD test statistic has much lower values, indicating that the inclusion of the cross-sectional averages controls well for the dependence. Nevertheless, it is significant at the 1% level. Analysis of the regression residuals has identified a common shock due to the Great Recession as a possible reason for the elevated correlation. Hence the model equation (15) is augmented by a dummy variable which has the value one in 2008Q3, 2008Q4 and 2009Q1, and zero otherwise. It is then included as an observed common factor in the estimation.

The results from this augmented DCCE model are presented in the last two columns of Table 7. The CD test statistic (1.90) is significant at the 10% level in the constant only case, showing that the residual cross-sectional correlation has not been eliminated completely even after the inclusion of the Great Recession dummy variable. Such correlation might, however, be due to omitted incidental trends. Turning to estimation results with trend included (last column), we note that this is indeed the case. The value of the CD statistic is 0.52, meaning that there is no strong cross-sectional dependence left among the residuals. Also more than a third (37%) of the time trend coefficients are significant at the 5% level, suggesting that the trending behaviour of the variables must be accounted for. The insignificant value of the CD test statistic serves also as evidence that the number of unobserved common factors is less than the number of observed variables in the system, thus rendering the DCCE estimator valid in this regard.

The residuals of the augmented DCCE model with trend are analyzed as a diagnostic check. Demetrescu et al.'s (2006) panel unit root test yields a value of -17.64 for Hartung's inverse normal test with $\kappa = 0.2$, convincingly rejecting the unit root null hypothesis. A plot of the estimated residuals, presented in Figure 4 in the Appendix, reveals no anomalies which could have resulted from potentially unattended structural breaks.

The next section compares the results from the Bai et al.'s (2009) estimators with those from the augmented DCCE estimator with trend and discusses their implications.

4.3 Discussion

The results produced by the Cup-BC, Cup-FM and the DCCE estimators are remarkably similar. The point estimate of the ERPT by both the Cup-BC and Cup-FM estimators is -0.37 with a standard error of 0.03, while the average long-run ERPT by the DCCE MG estimator is -0.33 with a standard error of 0.13. Hence their 95% confidence bands overlap, and neither includes the borderline values zero or one. Therefore, there is no evidence of PCP or LCP in the long-run, but rather of incomplete and low ERPT at the panel level. The average short-run ERPT coefficient estimated by the DCCE model is -0.19, which is significantly different from zero at the 10% level, but not at the 5%. These values are lower than the average exchange rate elasticity of 0.54 repored by Ben Cheikh and Rault (2016) for twelve euro-area countries in the period 1990Q3 - 2012Q4. Thus our results provide further evidence that ERPT has been declining over time, as found in the recent literature (Campa and Goldberg, 2005, Ben Cheikh and Rault, 2016). This may be attributed to the fact that our data comprises a longer period since the creation of the monetary union, so that a greater degree of convergence to more stable macroeconomic conditions has taken place in most countries of the panel. Another reason, however, might be the significant share of intra-euro area trade, which biases the aggregate ERPT estimates downwards (see Blagov, 2018).

Turning our attention to the average error-correction coefficient $\bar{\rho}$, reported in the last column of Table 7, we see that it is highly significant, once again highlighting the presence of cointegration in the system. Its value is -0.748, which implies high speed of adjustment to equilibrium – the aggregate implied half-life[8] is just 0.5 quarters.

The importing country's demand and the producers costs proxy are not statistically significant in either model. Domestic demand y_{it} not being an informative regressor for import prices despite predictions by economic theory is a result found also by other empirical studies, see, e.g., Campa and Goldberg (2005) and Beirne and Bijsterbosch (2009). The insignificance of the producers costs elasticity, on the other hand, could be based on unit labour costs not being a sufficiently good approximation for the producers' actual costs due to their slowly changing nature, for example. In a single-equation framework such results could raise questions regarding the validity of the ERPT results, as the estimation might be subject to omitted variable bias. In the panel framework, however, the unobserved common factors become a remedy for such problem. Being allowed to be $I(0)$, $I(1)$ or mixture of the two by both models, they capture the effects of any common shocks that influence the import prices and are not contained in the regressors. Hence besides controlling for cross-sectional dependence, the inclusion of unobserved common factors in the regression equations (9) and (13) has the added benefit of guarding against omitted variable bias. Therefore, the results for the exchange rate elasticity remain valid.

Finally, we discuss the individual countries' results from the DCCE estimator, presented in Table 11 in the Appendix. The estimated long-run exchange rate coefficients are quite heterogeneous with relatively large standard errors, which is the price to pay for the inclusion of current and lagged values of the cross-sectional averages in the individual equations. The long-run exchange rate elasticity is insignificantly different from zero for most of the older euro area member countries (Austira, Belgium, Finland, France, Luxembourg, Netherlands, Spain), and also for Switzerland and Estonia. Notable exceptions are Germany, Italy, and Portugal. The estimated long-run pass-through coefficients for these countries are -0.51, -1.08 and -1.26, respectively. On the other hand, the long-run ERPT coefficients are significant and with the expected negative sign for all non-euro area countries except for Switzerland: that is, for the Czech Republic (-0.37), Denmark (-1.46), Lithuania (-1.17), Poland (-0.30), Sweden (-0.82) and the United Kingdom (-0.47).

5 Summary and outlook

This paper establishes the presence of a long-run equilibrium relationship underlying the ERPT in a panel of nineteen European countries by employing new second-generation panel cointegration tests. Taking into account that unobserved global stochastic trends drive the cointegrating relationship and induce cross-sectional dependence in the data, aggregate long- and short-run exchange rate elasticities are estimated by novel panel estimators from most recent data. The results support the findings of earlier studies which report declining ERPT over time.

Future work could broaden the scope of the present analysis in several directions. Firstly, other data on import prices, which distinguish between imports from inside and

[8]The period of time needed for deviations in import prices to decline by half following a unit shock of the exchange rate.

outside the euro area, could be considered. As argued in Section 3, since a significant share of a Eurozone country's imports comes from other Eurozone contires, this may introduce a downward bias in an aggregate ERPT estimate. Monthly data on intra- and extra-euro area import prices is available from Eurostat, however sufficiently long time series exist for only six countries and the euro area as a whole. Hence the large-N panel estimators in this study would not be applicable and other estimation techniques would be needed. Secondly, ERPT to import prices could be investigated at an industry level, as considerable heterogeneity in ERPT elasticities has been reported across different industries (De Bandt et al., 2008; Blagov, 2018). Finally, the question of asymmetric ERPT as in Brun-Aguerre et al. (2017) seems a promising avenue to explore, in particular with regard to providing a better understanding of the pricing behaviour of exporters w.r.t. currency appreciation or depreciation and its implications for policy makers.

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A Appendix

A.1 Simes' (1986) intersection test

Let $p_1^* \leq \ldots \leq p_N^*$ be the p-values of N individual test statistics. Ordering them as $p_{(1)}^* \leq \ldots \leq p_{(N)}^*$, the joint null hypothesis $H_0 := \bigcap_i H_{i,0}, \;\; i = 1, \ldots, N$, (that all individual null hypotheses are simultaneously true) is rejected by Simes' test at significance level α if

$$p_{(i)}^* \leq \frac{i\alpha}{N} \;\; \text{for any}\, i = 1, \ldots, N. \tag{16}$$

Simes (1986) shows that the test is conservative under independence of the individual test statistics, that is

$$P_{H_0}\left\{p_{(i)}^* \geq \frac{i\alpha}{N}, i = 1, \ldots, N\right\} \geq 1 - \alpha. \tag{17}$$

Simes' intersection method has been introduced to testing for panel unit roots by Hanck (2013) and to testing for panel cointegration in dependent panels by Arsova and Örsal (2019).

A.2 Demetrescu et al.'s (2006) panel unit root test

Demetrescu et al. (2006) employ the modified inverse normal method of Hartung (1999), in which the dependence is captured by a single correlation coefficient ρ_t, which can be interpreted as a "mean correlation approximating the case of possibly different correlations between the transformed statistics" (Hartung, 1999). The statistic has a $N(0, 1)$ distribution under the null hypothesis. It is computed as [9]

$$t\left(\hat{\rho}_{\mathrm{t}}^*, \kappa\right) = \frac{\sum_{i=1}^{N} t_i}{\sqrt{N + (N^2 - N)\left(\hat{\rho}_{\mathrm{t}}^* + \kappa \cdot \sqrt{\frac{2}{(N+1)}}(1 - \hat{\rho}_{\mathrm{t}}^*)\right)}}. \tag{18}$$

Here $t_i = \Phi^{-1}(p_i^*)$ denote the probits and the variance of the denominator is augmented with an estimator of the correlation between the individual probits $\hat{\rho}_{\mathrm{t}}^*$: $\hat{\rho}_{\mathrm{t}}^* = \max\{-\frac{1}{N-1}, \hat{\rho}_{\mathrm{t}}\}$, where $\hat{\rho}_{\mathrm{t}} = 1 - \frac{1}{N-1}\sum_{i=1}^{N}\left(t_i - \frac{1}{N}\sum_{i=1}^{N} t_i\right)^2$.

The correction term $\kappa\sqrt{\frac{2}{(N+1)}}(1 - \hat{\rho}_{\mathrm{t}}^*)$, which simply scales the standard deviation of $\hat{\rho}_{\mathrm{t}}$ by a factor κ, aims to avoid a systematic underestimation of the denominator in eq. (18). For the κ parameter Hartung suggests two alternative values: $\kappa_1 = 0.2$ and $\kappa_2 = 0.1 \cdot \left(1 + \frac{1}{N-1} - \hat{\rho}_{\mathrm{t}}^*\right)$, where κ_2 is suitable mainly for smaller $\hat{\rho}_{\mathrm{t}}^*$.

The test statistic $t\left(\hat{\rho}_{\mathrm{t}}^*, \kappa\right)$ has a $N(0, 1)$ limiting distribution under the null hypothesis with a rejection region in the left tail.

[9] For simplicity the computation of the statistic is presented with unit weights for all cross-sections, as in its current implementation.

A.3 Auxiliary results

Table 8: Hanck's (2013) and Demetrescu et al.'s (2006) panel unit root tests

Country	lag	ADF_τ	p-value	Country	lag	ADF_τ	p-value	Simes' crit. value	
	p				e			5%	10%
Italy	2	−3.97	0.014	Poland	2	−3.13	0.029	0.003	0.005
Austria	2	−3.66	0.031	Sweden	1	−2.92	0.047	0.005	0.011
Belgium	2	−3.41	0.057	Denmark	4	−2.01	0.282	0.008	0.016
Luxembourg	4	−3.11	0.110	Austria	1	−1.95	0.308	0.011	0.021
Denmark	5	−3.08	0.119	Lithuania	4	−1.86	0.347	0.013	0.026
Portugal	2	−2.75	0.221	Germany	1	−1.86	0.347	0.016	0.032
Poland	3	−2.72	0.234	Italy	1	−1.84	0.357	0.018	0.037
France	2	−2.67	0.250	France	1	−1.71	0.425	0.021	0.042
United Kingdom	1	−2.60	0.282	Belgium	1	−1.68	0.436	0.024	0.047
Sweden	1	−2.56	0.298	Finland	1	−1.68	0.440	0.026	0.053
Finland	5	−2.27	0.443	Netherlands	1	−1.67	0.442	0.029	0.058
Germany	2	−2.13	0.521	Czech Republic	2	−1.66	0.446	0.032	0.063
Estonia	6	−2.12	0.525	Luxembourg	1	−1.66	0.449	0.034	0.068
Lithuania	3	−2.04	0.572	Spain	1	−1.62	0.466	0.037	0.074
Netherlands	3	−1.86	0.664	Portugal	1	−1.61	0.473	0.039	0.079
Czech Republic	2	−1.84	0.673	Norway	1	−1.38	0.588	0.042	0.084
Norway	1	−1.84	0.677	United Kingdom	1	−1.22	0.661	0.045	0.089
Switzerland	4	−1.63	0.771	Switzerland	1	−0.65	0.853	0.047	0.095
Spain	2	−1.48	0.827	Estonia	1	−0.23	0.929	0.050	0.100
	y				w			5%	10%
Sweden	3	−3.44	0.054	Italy	2	−3.91	0.016	0.003	0.005
Italy	1	−3.03	0.132	France	1	−3.24	0.084	0.005	0.011
France	2	−3.03	0.132	Norway	3	−3.15	0.102	0.008	0.016
Germany	1	−3.02	0.134	Spain	2	−2.92	0.162	0.011	0.021
Luxembourg	2	−2.94	0.157	Germany	2	−2.75	0.220	0.013	0.026
Estonia	3	−2.92	0.161	Switzerland	2	−2.73	0.230	0.016	0.032
United Kingdom	1	−2.89	0.170	Netherlands	3	−2.72	0.231	0.018	0.037
Switzerland	1	−2.89	0.171	Czech Republic	4	−2.70	0.240	0.021	0.042
Belgium	1	−2.73	0.226	Estonia	4	−2.63	0.267	0.024	0.047
Austria	1	−2.64	0.265	Portugal	2	−2.62	0.274	0.026	0.053
Finland	3	−2.57	0.294	Finland	2	−2.57	0.296	0.029	0.058
Lithuania	3	−2.54	0.309	Austria	4	−2.35	0.401	0.032	0.063
Spain	6	−2.49	0.330	Poland	1	−2.33	0.411	0.034	0.068
Netherlands	2	−2.40	0.378	Belgium	1	−2.29	0.433	0.037	0.074
Denmark	2	−2.39	0.381	Lithuania	3	−2.11	0.530	0.039	0.079
Poland	4	−2.27	0.446	Luxembourg	2	−2.02	0.580	0.042	0.084
Portugal	1	−2.14	0.513	United Kingdom	5	−1.94	0.624	0.045	0.089
Czech Republic	1	−2.06	0.561	Denmark	2	−1.89	0.650	0.047	0.095
Norway	5	−2.04	0.571	Sweden	1	−1.12	0.919	0.050	0.100

Variable	Hartung's κ_2 test statistic
p	-0.804
e	-0.279
y	-0.719
w	-0.668

Notes: ADF_τ denotes the Augmented Dickey-Fuller test statistic. A linear time trend is included in the test regressions for p, y and w, while only a constant is considered for e. The lag order is selected according to the modified AIC (MAIC) criterion of Ng and Perron (2001). The p-values are computed as in MacKinnon (1996); the author is grateful to Christoph Hanck for the GAUSS code. Results for each variable are sorted according to the p-values in ascending order for ease of comparison with the corresponding Simes' critical value.

Table 9: Pesaran's (2007) $CIPS$ panel unit root test

lag	p	e	y	w	Δp	Δe	Δy	Δw
6	-2.037	-1.662	-2.322	-1.548	-3.216^{***}	-3.708^{***}	-2.550^{***}	-2.659^{***}
5	-2.022	-1.734	-2.210	-1.503	-3.378^{***}	-4.134^{***}	-2.807^{***}	-3.071^{***}
4	-2.049	-2.221^{**}	-2.061	-1.571	-3.971^{***}	-4.442^{***}	-3.084^{***}	-3.751^{***}
3	-2.218	-2.460^{***}	-2.267	-1.851	-4.315^{***}	-3.880^{***}	-3.768^{***}	-4.300^{***}
2	-2.517	-2.305^{**}	-2.116	-2.050	-5.283^{***}	-4.236^{***}	-4.316^{***}	-4.520^{***}
1	-2.702^{**}	-2.278^{**}	-2.213	-1.857	-5.798^{***}	-5.200^{***}	-5.168^{***}	-5.196^{***}

Notes: Trend is included in the test regressions for p, y and w, while only a constant is considered for e. The 10%, 5% and 1% critical values for the model with constant only are -2.11, -2.2 and -2.36, and -2.63, -2.7 and -2.85 for the model with trend, respectively.
*, ** and *** denote significance at the 10, 5 and 1% level, respectively.

Table 10: Estimated factor loadings

Country	$\hat{\Lambda}_1^p$	$\hat{\Lambda}_2^p$	$\hat{\Lambda}_1^e$	$\hat{\Lambda}_2^e$	$\hat{\Lambda}^y$	$\hat{\Lambda}^w$
Austria	1.19	0.57	-1.26	-0.13	1.14	1.76
Belgium	1.25	-0.20	-1.27	-0.15	1.13	1.47
Czech Republic	1.06	0.23	-0.46	-1.71	1.15	0.17
Denmark	0.71	-1.87	-1.23	0.49	0.81	1.20
Estonia	0.94	-0.90	-0.83	1.52	0.95	0.84
Finland	1.08	-0.18	-1.25	0.25	1.11	-0.32
France	1.37	-0.08	-1.27	-0.28	1.22	1.36
Germany	1.35	-0.26	-1.27	-0.17	1.11	0.35
Italy	1.32	-0.04	-1.27	-0.20	1.23	0.57
Lithuania	0.56	1.78	-0.15	1.32	0.93	-0.06
Luxembourg	0.09	-2.32	-1.26	0.17	0.64	0.84
Netherlands	0.88	0.72	-1.27	-0.01	1.16	1.12
Norway	0.80	-0.14	-0.22	-1.52	0.38	1.11
Poland	0.30	1.60	-0.09	-2.05	0.36	0.38
Portugal	1.20	-0.79	-1.27	-0.06	0.94	1.09
Spain	1.24	-0.41	-1.27	-0.07	1.06	0.96
Sweden	0.98	0.72	-0.25	-1.90	1.03	-0.72
Switzerland	0.62	0.70	0.22	0.50	1.06	1.68
United Kingdom	0.86	0.84	0.60	-1.07	1.01	0.63

Figure 2: Estimated common factors from the Cup-FM model

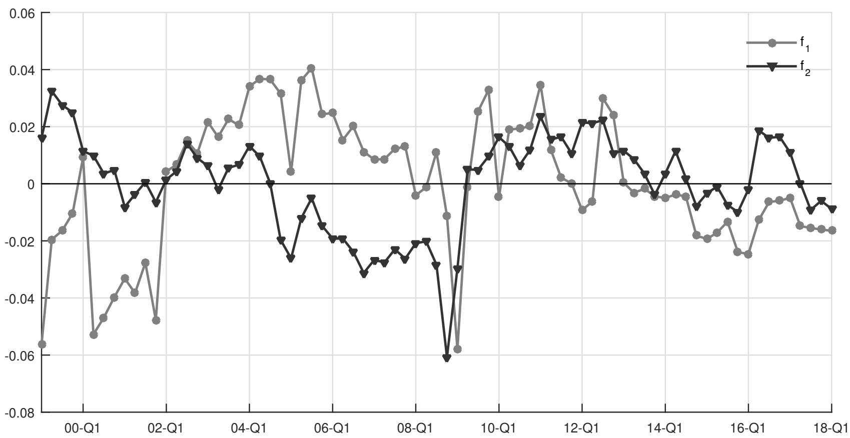

Figure 3: Estimated residuals from the Cup-FM model

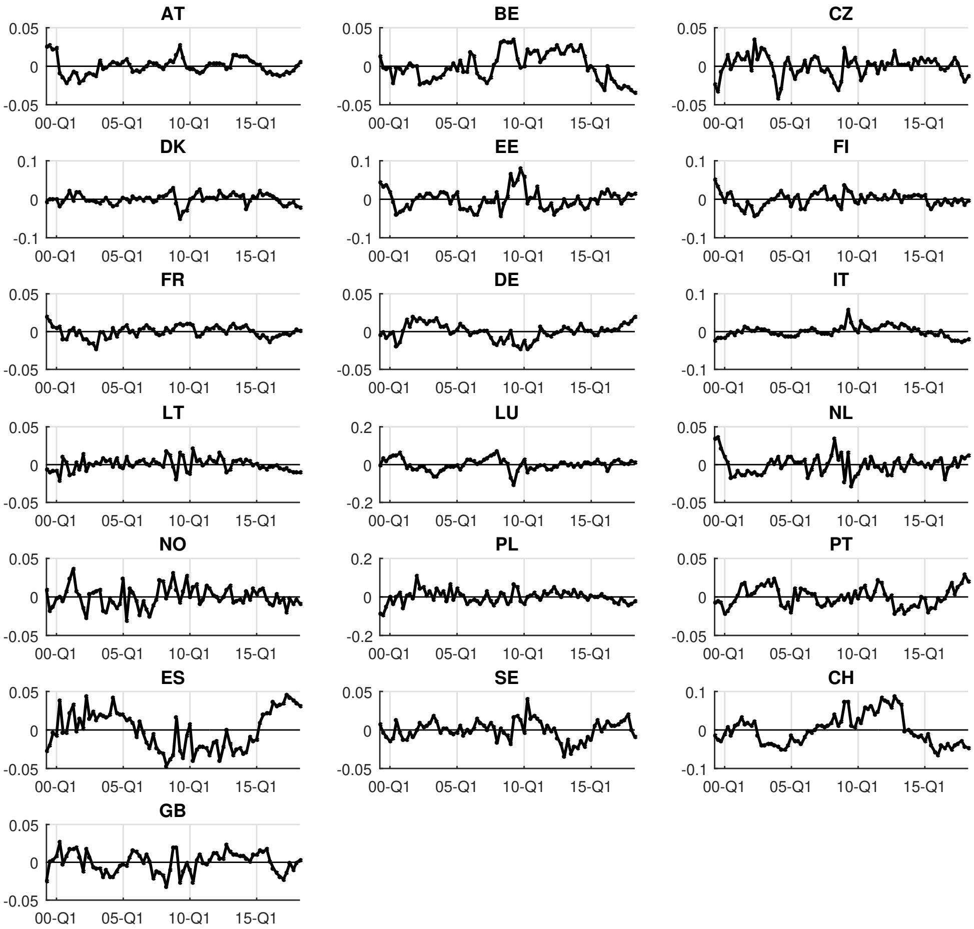

Figure 4: Estimated residuals from the augmented DCCE model with trend

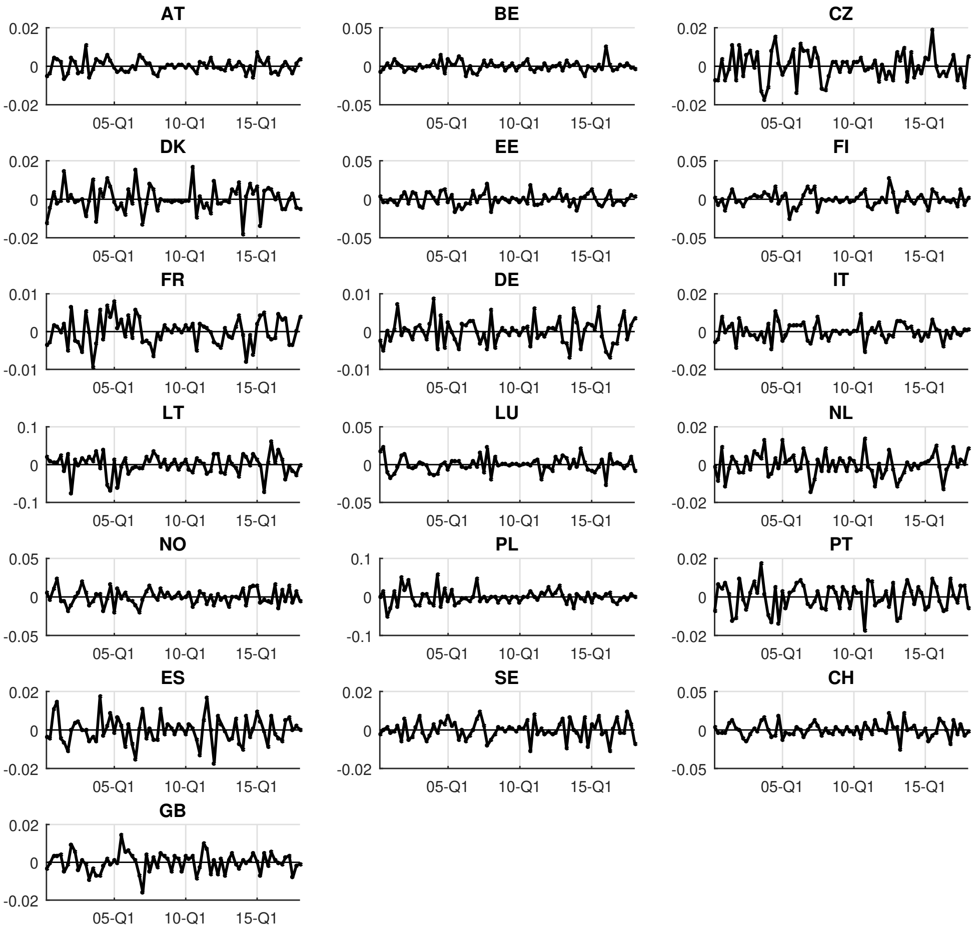

Table 11: ERPT estimation results for the individual countries by the DCCE estimator

Country	Exchange rate e long-run	Exchange rate e short-run	Domestic demand y long-run	Domestic demand y short-run	Producers' costs w long-run	Producers' costs w short-run	EC coefficient ρ_i
Austria	−0.11	−0.20	−0.15	0.34	0.01	0.06	−0.68***
	(0.17)	(0.20)	(0.25)	(0.23)	(0.12)	(0.13)	(0.19)
Belgium	0.08	−0.19	1.27	0.12	0.32	−0.04	−0.47***
	(0.49)	(0.37)	(0.89)	(0.63)	(0.38)	(0.27)	(0.15)
Czech Republic	−0.37***	−0.35***	−0.34	0.79	−0.46*	0.00	−0.83***
	(0.14)	(0.11)	(0.34)	(0.52)	(0.24)	(0.19)	(0.18)
Denmark	−1.46***	−1.11***	−0.42	0.30	−0.50***	0.01	−1.04***
	(0.39)	(0.30)	(0.30)	(0.34)	(0.15)	(0.21)	(0.17)
Estonia	−0.05	0.26	0.43**	0.32*	0.09	0.11	−0.88***
	(0.09)	(0.18)	(0.20)	(0.19)	(0.12)	(0.20)	(0.15)
Finland	−0.12	0.30	0.06	0.47	0.33	0.41	−0.80***
	(0.27)	(0.38)	(0.35)	(0.45)	(0.31)	(0.19)	(0.15)
France	0.03	0.29	0.40	0.00	0.21	0.05	−0.81***
	(0.14)	(0.21)	(0.47)	(0.42)	(0.16)	(0.17)	(0.18)
Germany	−0.51**	−0.44***	−0.33*	−0.06	0.10	0.22*	−0.64***
	(0.24)	(0.14)	(0.20)	(0.17)	(0.13)	(0.12)	(0.15)
Italy	−1.08**	−0.72***	1.00**	0.02	0.20	−0.04	−0.61***
	(0.45)	(0.23)	(0.44)	(0.35)	(0.28)	(0.14)	(0.16)
Lithuania	−1.17***	−0.98***	−0.62	−1.71**	0.52	−0.49	−0.82***
	(0.37)	(0.33)	(0.71)	(0.73)	(0.36)	(0.39)	(0.15)
Luxembourg	−0.16	−0.26	0.80	0.67*	−0.03	0.49*	−0.61***
	(0.67)	(0.68)	(0.55)	(0.35)	(0.26)	(0.26)	(0.13)
Netherlands	−0.03	−0.18	−0.53	−0.27	0.07	0.01	−0.87***
	(0.21)	(0.29)	(0.40)	(0.43)	(0.12)	(0.15)	(0.16)
Norway	−0.46*	−0.22**	−0.62	−0.18	−0.21	0.19	−0.57***
	(0.24)	(0.11)	(0.82)	(0.40)	(0.35)	(0.30)	(0.19)
Poland	−0.30**	−0.25	−0.08	−0.27	−0.75*	−0.53	−1.13***
	(0.13)	(0.19)	(0.31)	(0.62)	(0.42)	(0.34)	(0.20)
Portugal	−1.26**	0.20	1.18***	0.42	0.46*	0.26*	−0.71***
	(0.55)	(0.61)	(0.36)	(0.35)	(0.24)	(0.15)	(0.14)
Spain	0.64	0.48	0.63**	2.31***	0.41	−0.15	−0.84***
	(0.43)	(0.54)	(0.27)	(0.67)	(0.37)	(0.21)	(0.15)
Sweden	−0.82***	−0.41***	1.10***	−0.15	0.00	−0.16*	−0.80***
	(0.17)	(0.07)	(0.40)	(0.28)	(0.06)	(0.09)	(0.12)
Switzerland	0.65	−0.11	−2.61	1.38*	−12.25	1.04	−0.10
	(1.13)	(0.12)	(4.30)	(0.82)	(14.60)	(0.92)	(0.10)
United Kingdom	−0.47***	−0.31***	−0.38	−0.45	−0.32	−0.23*	−0.55***
	(0.18)	(0.06)	(0.36)	(0.36)	(0.24)	(0.12)	(0.15)

Notes: Results from the DCCE estimator with trend, augmented with a Great Recession dummy variable. Standard errors are presented in brackets.
*, ** and *** denote significance at the 10, 5 and 1% level, respectively.

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Leuphana Universität Lüneburg
Institut für Volkswirtschaftslehre
Postfach 2440
D-21314 Lüneburg
Tel.: ++49 4131 677 2321
email: brodt@leuphana.de

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