Exchange Rate and Consumer Prices in the Euro Area: A Cointegrated VAR Analysis

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Abstract

This paper analyzes the exchange rate pass-through (ERPT) into consumer prices for 12 EA countries within a CVAR framework. Using the Johansen cointegration procedure, results indicate the existence of one cointegrating vectors at least for each EA country of our sample. When measuring the long-run effect of exchange rate changes on consumer prices, we found a wide dispersion of ERPT elasticities, especially between "peripheral" and "core EA" economies. For instance, consumer prices rise by 84% in Portugal following one percent depreciation of exchange rate, while for the German economy the extent of pass-through is not exceeding 0.20%. Besides, the loading factors point out a very slow adjustment of consumer prices towards their long-run equilibrium across EA countries. This would explain the weakness of ERPT estimates in the short-run.

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1 Introduction

In this paper we examine the overall effect of exchange rate changes on consumer prices, an issue which is most relevant for monetary policy in the euro area. Movements in the exchange rate can have a significant influence on inflation dynamics, both in terms of their direct effect on prices and their indirect effect through changes in the aggregate demand and wages. Thorough knowledge of the underlying behavior behind pass-through is a key input to determine the proper monetary policy responses. Policymakers must be able to prevent the changes in relative prices - such as those stemming from exchange rate movements - which may fuel a continuous inflationary process.

In order to provide reliable estimates, many recent empirical studies has adopted vector autoregressive (VAR) models (see e.g. MCCARTHY, 2007; CA'ZORZI, KAHN, and SÁNCHEZ, 2007; CHOUDHRI, FARUQUEE, and HAKURA, 2005; FARUQEE, 2006; ITO and SATO, 2008, to name but a few). The advantage of simultaneous equation approach allows for potential and highly likely endogeneity between the variables of interest, i.e exchange rate and price series. Ignoring such simultaneity would result in simultaneous equation bias. In a single-equation pass-through regression, for example, the fact that domestic inflation may affect the exchange rate is ignored. However, an important drawback regarding VAR-based approaches, is that the time-series properties of the data - particularly non-stationarity and cointegration issues - was neglected. To our knowledge, few are studies who deals with the issue of ERPT within Cointegrated VAR (CVAR) Framework. For five largest euro area (EA) countries, HÜFNER and SCHRÖDER (2002) found that the endogenous variables in their VAR system are cointegrated using the Johansen procedure.[1]. The authors estimated a Vector Error Correction Models (VECM), incorporating the long-run relationships among the variables, and derived impulse responses functions in order to examine how external shocks are propagated to domestic prices. The main drawback of HÜFNER and SCHRÖDER's (2002) study is that the information contained in "levels" variables was not analyzed. In other words, they did not measure the long-run ERPT in the "equilibrium" relationship. As a matter of fact, most of previous studies has focused on techniques and tools of VAR models, such as impulse response functions, variance decompositions and historical decompositions, to explore the impact of exchange rate shocks, and ignored the information contained in the cointegrating vector.

[1] The five largest EA countries are: Germany, France, Italy, Netherlands and Spain.

Therefore, in our paper we propose to use cointegration analysis to study the extent of pass-through, by focusing on the long-run equilibrium relationship contained in the cointegrating space. We propose a CVAR model as it allows us to take proper account of the non-stationarity of the data, i.e. look for cointegration properties in the data, and at the same time disentangle short- and long-run dynamics. This exercise is conducted for 12 EA Member States. As a major problem for an analyzing pass-through in the EA is the lack of sufficiently long time series (see e.g. HÜFNER and SCHRÖDER, 2002; HAHN, 2003), our study propose a larger sample period covering the pre- and post-euro episodes.

The rest of the paper is organized as follows: the next section outlines the CVAR model used for the empirical analysis. In section 3, the data set and their properties are discussed. Section 4 contains the main results from the cointegration analysis. Section 5 concludes.

2 Empirical Methodology

Our analysis aims at capturing the effects of changes in exchange rates on consumer prices which is the key variable for the policy issues. Our model relates consumer prices (cpi_t) to the the trade weighted effective exchange rate (e_t), oil price(oil_t), aggregate income (y_t) and interest rates (r_t) in cointegrated VAR (CVAR) framework. Using Johansen procedure, CVAR analysis could be useful in this context as it allows us to take proper account of the non-stationarity of the data, looking for cointegration properties in the data, and at the same time disentangle short- and long-run dynamics. Thus, it enables retention of the important information contained in levels variables. This levels information is lost in more traditional first-difference VAR models.

As a starting point of the analysis, we consider the following vector of endogenous variables:

$$x_t = (cpi_t, e_t, oil_t, y_t, r_t)' \tag{1}$$

Having firstly tested the stationarity of the variables, we apply cointegration tests for each country to check whether long-term relationships exist between the variables. The Johansen test is used to assess whether or not cointegration exists in the system of variables. In order to describe this, we begin firstly by considering the following system of five-equation VAR(k) model:

$$x_t = A_1 x_{t-1} + \ldots + A_k x_{t-k} + \mu + \psi D_t + \varepsilon_t, \qquad t = 1, 2, \ldots, T, \tag{2}$$

Equation (2) can be expressed as an error or vector equilibrium correction model (VECM), i.e. a CVAR, which is formulated in terms of differences as follows:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu + \psi D_t + \varepsilon_t \tag{3}$$

Where x_t is a (5×1)vector of $I(1)$ endogenous variables as given in Equation (1); k is lag lentgh;μ is a constant term; D_t is a vector including deterministic variables (centered seasonal dummies and intervention dummies) and weakly exogenous variables; and ε_t is a $(k \times 1)$ vector of errors which are assumed identically and independently distributed and follow a Gaussian distribution $\varepsilon_t \sim$ iid $N_p(0, \Omega)$, with Ω denotes the variance-covariance matrix of the disturbances.

The VECM representation, i.e. the CVAR model, encompasses both short- and long-run information of the data. The matrix Π assembles the long-run information and the Γ_is contain the short-run properties. $\Pi = \alpha\beta^{'}$ has reduced rank r. The matrices α and β are of dimension $(5 \times r)$, α depicts the speed of adjustment, and β represents the cointegrating vectors. The Johansen procedure estimates equation (3) subject to the hypothesis that Π has a reduced rank $r < 5$. This hypothesis can be written as:

$$\mathrm{H}(r) = \alpha\beta^{'} \tag{4}$$

JOHANSEN and JUSELIUS (1990) show that, under certain circumstances, the reduced rank condition implies that the processes Δx_t, and $\beta^{'} x_t$, are stationary even though x_t, itself is non-stationary. The stationary relations $\beta^{'} x_t$, are referred to as cointegrating relations. To determine the number of cointegrating vectors (r) in the system, i.e. the cointegration rank, we employ the widely used trace test statistics (see Table 7 in Appendix A):

$$\text{Trace} = -N \sum_{i=r+1}^{5} \ln(1 - \hat{\lambda}_i) \tag{5}$$

where N is the number of observations and $\hat{\lambda}_i$ is the estimated eigenvalue. When the appropriate model has been identified for the system in terms of lag length and cointegration rank, the coefficients on the α matrix reveal the long-run dynamic while the coefficients on the β matrix reveal the drivers towards the long-run equilibrium.

In order to determine the responsiveness of consumer prices to exchange rate changes, the coefficient estimates of the cointegrating vectors are normalized on consumer prices. Thus, the coefficients on exchange rate indicate the degree of ERPT. Also, in the cointegration analysis, we focus on the first cointegrating vector.

As discussed in JOHANSEN and JUSELIUS (1992), the first cointegrating vector is the most associated with the stationary part of the model since it has the highest eigenvalue. After estimating the ERPT coefficient in the long-run, we proceed by testing a number of restrictions on the long-run parameters in order to examine specific hypotheses on pass-through:

$\mathbf{H}_1$: Full ERPT to consumer prices with other long-run parameters unrestricted, i.e. test of whether the first cointegrating is as follows

$$\begin{array}{lcccccl} & cpi & e & oil & y & r & \\ \beta_1' = & (1 & 1 & \gamma & \eta & \varphi) & \sim I(0) \end{array}$$

$\mathbf{H}_2$: Full ERPT to consumer prices with zero constraints on other long-run parameters, i.e. test of whether the first cointegrating is as follows

$$\begin{array}{lcccccl} & cpi & e & oil & y & r & \\ \beta_1' = & (1 & 1 & 0 & 0 & 0) & \sim I(0) \end{array}$$

$\mathbf{H}_3$: Zero ERPT to consumer prices with other long-run parameters unrestricted, i.e. test of whether the first cointegrating is as follows

$$\begin{array}{lcccccl} & cpi & e & oil & y & r & \\ \beta_1' = & (1 & 0 & \gamma & \eta & \varphi) & \sim I(0) \end{array}$$

$\mathbf{H}_4$: Zero ERPT to consumer prices with zero constraints on other long-run parameters, i.e. test of whether the first cointegrating is as follows

$$\begin{array}{lcccccl} & cpi & e & oil & y & r & \\ \beta_1' = & (1 & 0 & 0 & 0 & 0) & \sim I(0) \end{array}$$

If $\mathbf{H}_1$ or $\mathbf{H}_2$ holds, this would imply that exchange rate changes are fully transmitted to consumer prices, while if $\mathbf{H}_3$ or $\mathbf{H}_4$ holds, there is a null pass-through, i.e. consumer prices do not respond to currency movements.

3 Data selection and their properties

In order to measure the effects of exchange rate changes on consumer prices, our CVAR model contains five endogenous variables. In addition to our two key variables - exchange rate and consumer prices - we have included three macroeconomic variables affect the inflation of consumer prices directly. The choice of the variables is based on the following considerations: first, oil prices enter the VECM to controls for the impact of supply shocks; second, to balance the model with respect to the demand side, a measure of national income is added in the system; and finally, a short-run interest rate is included to allow for the effects of monetary policy.[2] As discussed by PARSLEY and POPPER (1998), taking into account monetary policy significantly improves the estimation results of ERPT. In fact, central banks are concerned with keeping domestic inflation within its target range which may insulate prices from exchange rate movements. Thus, neglecting the effects of monetary policy results in the common omitted variables problem. In a subsequent step, our basic model will be augmented to include the whole pricing, i.e. import prices, producer prices and consumer prices.

In this study, we focus our analysis on 12 EA countries, namely Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Luxembourg, Netherlands and Portugal). For each country a set of quarterly data was collected covering the time period 1980:1 to 2010:4. The consumer price (cpi_t) is the overall consumer price index to provide the broadest measure of inflation at the consumer level. We did not use the Harmonized Index of Consumer Prices (HICP) due the short data availability of this variable. Exchange rate data are effective nominal exchange rates of the national currencies which use the trade weights of each country.[3] The oil price (oil_t) is represented by a crude oil price index denominated in US dollar in order to avoid multicollinearity issues with the exchange rate.[4] The national income (y_t) is proxied by the real GDP. The 3-month interest rate is used to model monetary policy. To collect data, we have followed a cascade order, choosing when possible only one institutional source, i.e. IMF's *International Financial Statistics* and OECD's *Main Economic Indicators* and *Economic Outlook*, in that order.

Next, we check the non-stationarity of the data. In order to test this, each of the variables are tested for unit roots using the traditional ADF-test which tests the null hypothesis of non-stationarity. To ensure robustness the order of integration of the variables, ADF test is supplemented by two stationarity test. First, the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test which is structured under the opposite null hypothesis that of stationarity against a unit root alternative. Second, the DF-GLS test, proposed by ELLIOTT, ROTHENBERG, and STOCK (1996), which is an augmented Dickey-Fuller test, similar to the test performed Dickey-Fuller tests, except that the time series is transformed via a generalized least squares (GLS) regression before performing the test. ELLIOTT, ROTHENBERG, and STOCK (1996) have shown that this test has significantly greater power than the previous versions of the augmented Dickey-Fuller test. In constructing the unit root tests, the variables in levels were tested in the presence of both an intercept

[2] With the exception of interest rates, all variables are in logs.

[3] The nominal effective exchange rate is defined as domestic currency units per unit of foreign currencies, which implies that an increase represents a depreciation for domestic country.

[4] McCarthy uses local price of oil to identify supply shocks, but this will include the exchange rate effect. Thus, much of the exchange rate effect may be mixed into the supply shock.

and trend. The subsequent tests of first differences included only an intercept given the lack of trending behaviour in the first-differences series.

Results of the unit root tests of the variables reveal that the majority of the variables to have been generated via an integrated of order one $I(1)$ process (see Table 5 in the Appendix A). First-differences variables are found to be stationarity in at least two of the three tests undertaken for most cases. We can summarize the results of the three unit root tests as follows: According to ADF tests all variables are stationary in first differences with exception of consumer prices in Ireland; for the KPSS test, the null hypothesis of stationarity is accepted for most of the variables in first differences except for consumer prices in Netherlands and Portugal, while import prices are stationary in level for Luxembourg. Finally, we find that all variables $I(1)$ within DF-GLS test, with the exception of consumer prices for Luxembourg and Portugal, nominal effective exchange rate for Portugal and producer prices in Ireland.

Building on these results, the Johansen cointegration tests were undertaken to assess the existence of long-run equilibrium relationships among the variables. Given that the choice of the rank of Π should be made on the basis of a well-specified model, it is important to include the appropriate number of lags before rank tests are undertaken. Thus, the lag structure for each VECM was based on assessment of the AIC compatible with well-behaved residuals. Results from trace test, reported in Table 7 in Appendix A, indicates the presence of one cointegrating vectors at least for each EA country (as in Austria and Netherlands). The null hypothesis of no cointegration was rejected for all our EA countries, with a cointegration rank identified of between one and three. A summary of the number of cointegrating equations (CE) identified across each country as well as the optimal lag length is reported in Table 1.

4 Cointegration Analysis

In the section, we focus on the long-term part of our VECM Framework, i.e. the long-run relationships present in the cointegrating space. Our primary concern is to assess the relative signs and magnitudes of the long-run ERPT coefficients across EA countries. To this end, there are some issues that must be considered here. First, ERPT equation must contain a proxy for foreign costs as recommended by the bulk of empirical literature (see GOLDBERG and KNETTER, 1997). Given that foreign costs are an exogenously determined variables regarding our EA countries, we propose to include a proxy for costs of a country's trading partners as an exogenous variables in our basic VECM. Therefore, to capture changes in foreign costs, we construct a typical export partners cost proxy (W_t^*) that used throughout the ERPT literature (see *inter alia* BAILLIU and FUJII, 2004; CAMPA and GOLDBERG, 2005):

$W_t^* = Q_t \times W_t/E_t$, where Q_t is the unit labor cost based real effective exchange rate, W_{it} is the domestic unit labor cost and E_t is the nominal effective exchange rate. Taking the logarithm we obtain the following expression: $w_{it}^* = q_t + w_t - e_t$. Since the nominal and real effective exchange rate series are trade weighted, we obtain a measure of foreign firms' costs with each partner weighted by its importance in the domestic country's trade.[5]

Table 1: Summary of basic CVAR Models

Country	**VAR lags**	**Number of CE**	**Model specification**
Austria	2	1	Restricted trend
Belgium	2	2	Restricted trend
Germany	2	3	Restricted trend
Spain	3	2	Unrestricted intercept
Finland	1	2	Restricted trend
France	2	3	Restricted trend
Greece	5	2	Restricted trend
Ireland	3	3	Unrestricted intercept
Italy	2	3	Restricted trend
Luxembourg	3	2	Unrestricted intercept
Netherlands	2	1	Restricted trend
Portugal	2	3	Restricted trend

Note: The optimal number of lags in the VECM was determined using the AIC criterion. The number of cointegrating equations is equal to the number of cointegration equations found by the Johansen trace test.

Second, besides the seasonal dummy variables a shift, we introduce dummy in 1990:07 (D_{90}) and kicks in until the end of the sample. Chow tests for multivariate models, as introduced by CANDELON and LUTKEPOHL (2001), denote the presence of structural break in vicinity of 1990 (see Table 8 in Appendix B).[6] Including D_{90} helps to restore the stability of the cointegrating vectors. Figure 2 in Appendix C provides an indication of the stability of the cointegrating vectors by means of recursive estimates of the eigenvalues. Plots reveals that recursive estimates of the eigenvalues, over a 40-month window, are broadly constant for most of EA countries which is an indication of the stability of the cointegrating vectors identified. It is worth noting that centered seasonal dummies, shift dummy and exogenous foreign costs enter the vector D_t in equation (3). Final issue concerns the specification of VECM of each of our 12 EA countries. In most of the cases

[5] To measure the extent of pass-through in the non-US G-7 countries, CHOUDHRI, FARUQUEE, and HAKURA (2005) enter two foreign exogenous variables - foreign interest rate and the foreign consumer price index - in their first-difference VAR model.

[6] We can select May 1998, the month on which the parities among European currencies replaced by the euro were announced, as the date for the break. However, as in most of empirical literature (see CAMPA and GOLDBERG, 2002, 2005, among others), the date of creation of the euro has not been found as a regime shifts in the monetary union countries.

the most appropriate model appears to be that which includes a trend in the cointegrating equation and permits the intercept to enter both the cointegration space and the VAR, i.e. unrestricted intercept and restricted trend. The only exceptions are Spain, Ireland and Luxembourg where we include only a constant in the cointegrating equations and in the short-term part of the VECM, i.e. unrestricted intercept.[7] Summary of our 12 CVAR models are reported in Table 1.

4.1 Long-run ERPT to consumer prices

As we mentioned above, we focus on the first (most statistically significant) cointegrating equation to measure the extent of pass-through in the long-run. The long-run parameters for each unrestricted CVAR model are reported in Table 2. In unrestricted form, it is clear that the signs of the parameters appear in most cases to accord with priors. In most of case, positive coefficients are observed on exchange rate, oil prices, real GDP and the interest rate series.[8] Thereby, a rise in exchange rate (i.e. depreciation), in oil prices, in real GDP or in interest rate is associated with a higher domestic consumer prices. In some cases, there appears to be some inconsistency regarding the sign on GDP or interest rate, but roughly speaking, our results tend to agree with the expected signs.

Concerning the degree of ERPT, our results point out cross-country differences in the responsiveness of consumer prices in the long-run (see Figure 1).[9] Germany, Finland and France have the lowest coefficients in our sample of EA, with long-run ERPT not exceeding 0.20%. The degree of ERPT appears to be most prevalent in Portugal and Greece. For Portugal, a 1% depreciation of exchange rate increases domestic consumer prices by roughly 0.84%, while for Greece, consumer prices rise by 58% following one percent depreciation of exchange rate. In their study on 20 industrialized countries, GAGNON and IHRIG (2004) found that Portugal and Greece have the highest long-run response of consumer prices over the period 1972 to 2000. The pass-through elasticities are: 0.43 percent for Portugal and 0.52 percent for Greece. Nevertheless, these pass-through coefficients are still lower compared to our results. As a matter of fact, GAGNON and IHRIG (2004) did not find any evidence of cointegration between variables in levels, that's why they estimate their pass-through *single-equation* in first-differences. Thus, their definition of "long-run" effect stems from the feedback effects resulting from the

[7] The use of unrestricted intercepts and restricted trends is consistent with data that exhibit some form of trending behaviour. When we expect some of the data to be trend stationary, a good idea is to start with a restricted linear trend and then test the significance of the trends.

[8] The positive relationship between consumer prices and interest rate is consistent with the long-run Fisher effect.

[9] All long-run rates of pass-through are significantly different from zero in our sample of EA countries.

inclusion of the lagged dependent inflation terms (dynamic equation).[10] We see that taking into account the times series proprieties of the date (non-stationarity and cointegration relationship) may give a more reliable long-run ERPT estimates.

Table 2: Coefficients of first cointegrating vector

Country	cpi_t	e_t	oil_t	gdp_t	r_t	$trend$
Austria	1,000	0,248*	0,124**	0,712***	-0,038***	0,026*
		(1,799)	(2,494)	(10,542)	(-10,594)	(1,819)
Belgium	1,000	0,282***	0,468***	-0,213	0,019***	0,007***
		(3,800)	(3,373)	(-0,506)	(4,250)	(2,835)
Germany	1,000	0,169**	0,464	0,968***	0,073***	0,011***
		(2,305)	(1,523)	(2,606)	(8,608)	(6,302)
Spain	1,000	0,337**	0,535*	0,880	0,002	-
		(2,254)	(1,910)	(1,538)	(0,310)	
Finland	1,000	0,117*	0,413***	0,578***	-0,009***	-0,004*
		(1,897)	(2,915)	(7,345)	(-2,946)	(-1,753)
France	1,000	0,166***	0,279**	-0,290	0,013***	0,006***
		(2,693)	(2,260)	(-0,752)	(2,781)	(2,747)
Greece	1,000	0,576***	1,027***	0,371	-0,036***	0,031***
		(4,494)	(5,416)	(1,002)	(-3,957)	(6,192)
Ireland	1,000	0,397***	0,208**	0,485***	0,003	-
		(4,009)	(2,495)	(5,126)	(1,571)	
Italy	1,000	0,352***	0,486***	1,098***	0,012***	0,003*
		(5,231)	(3,394)	(3,129)	(2,720)	(1,813)
Luxembourg	1,000	0,339***	0,667***	0,468***	0,008**	-
		(5,472)	(4,801)	(13,338)	(2,089)	
Netherlands	1,000	0,298***	0,683***	-0,637***	-0,039***	0,044***
		(4,400)	(5,327)	(-13,048)	(-10,820)	(8,236)
Portugal	1,000	0,833***	0,056	-0,084	0,013*	0,018***
		(8,553)	(1,244)	(-0,206)	(1,739)	(2,732)

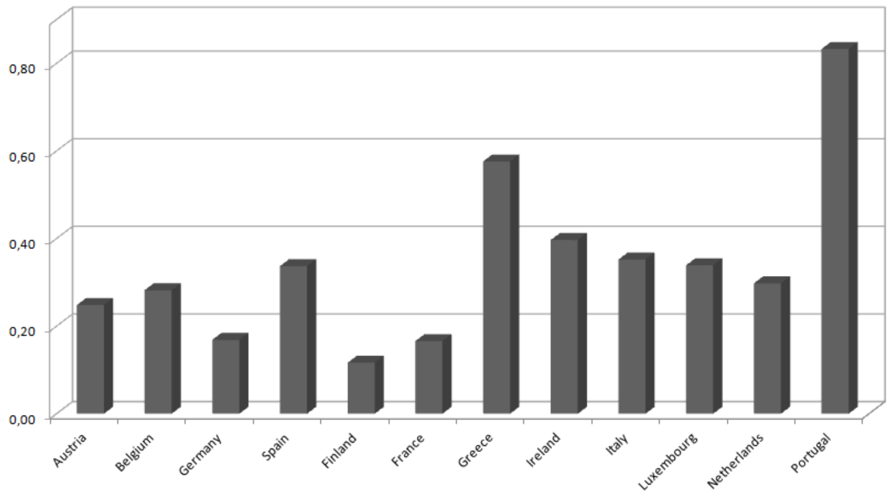
Note: *, ** and *** denote significance level at 10%, 5% and 1% respectively. t-stat are in parentheses.

Moreover, it should be noted that the response of consumer prices is still weak in comparison to import prices. Several explanations have been put forward by ERPT literature . In fact, imported goods have to go through distribution sector before they reach consumers in domestic country. Thus, local distribution costs (such as transportation costs, marketing, and services), may cause a wedge between import and consumer prices. Also, competitive pressure in distribution sectors may explain why consumer prices do not respond dramatically to exchange rate changes. As discussed in BACCHETTA and VAN WINCOOP (2002), the weakness of CPI

[10] The effects of an exchange rate change in period t will influence inflation over several periods subsequent to this as a result of these feedback effects.

inflation reaction to exchange rate changes is due, in part, to differences in the optimal pricing strategies of foreign producers and domestic wholesalers/retailers. Due to competitive pressure in the domestic market, domestic wholesalers import goods priced in foreign currency (PCP) and resell them in domestic currency (LCP). This would entail much lower ERPT to CPI inflation than expected. Finally, we can add that substitution effect can occur. If home currency is depreciating, domestic firms or wholesalers may reduce sourcing foreign products (since their price becomes higher), shifting towards substitute domestically produced goods.

Figure 1: Long-run ERPT in EA countries



Source: Personal calculation.

Thereafter, in Table 3, we set out the adjustment coefficients (or loading factors) revealing the speed with which the long-run equilibrium is achieved. It known that lack of significance on these parameters indicates the presence of weak exogeneity, meaning that the variable does not respond to or correct for deviations to the long-run equilibrium. For oil prices, we find non-significant adjustment coefficients in the half of EA countries, this could be a sign of weak exogeneity. We could impose weak exogeneity on oil prices but it does not alter the long-run coefficients of ERPT. We keep on only foreign as the only exogenous variables in our CVAR models.

Table 3: Adjustment coefficients in the basic VECM

	Austria	**Belgium**	**Germany**	**Spain**	**Finland**	**France**
cpi_t	0,018***	-0,056***	-0,026***	-0,053***	-0,013***	-0,042***
	(8,071)	(-8,626)	(-7,733)	(-6,782)	(-2,514)	(-8,395)
e_t	0,069***	-0,083	-0,075*	-0,153***	0,106***	-0,039*
	(3,050	(-1,022)	(-1,689)	(-2,648)	(4,162)	(-1,621)
oil_t	-0,096	0,018**	0,041*	0,024	0,006	0,069***
	(-1,451)	(2,447)	(1,786)	(0,783)	(0,171)	(3,473)
gdp_t	-0,003**	0,016***	0,049***	-0,021**	-0,009*	0,004
	(-1,998)	(2,573)	(3,386)	(-2,291)	(-1,641)	(0,531)
r_t	-0,453*	1,822*	0,875**	0,552	1,666*	-0,047***
	(-1,630)	(1,727)	(2,370)	(0,382)	(1,873)	(-3,286)
	Greece	**Ireland**	**Italy**	**Luxembourg**	**Netherlands**	**Portugal**
cpi_t	0,033**	-0,119***	-0,044***	-0,073***	0,022***	-0,041***
	(2,342)	(-11,256)	(-6,860)	(-6,110)	(6,313)	(-7,424)
e_t	0,238***	-0,225***	-0,027	-0,070**	0,173***	-0,058**
	(6,179)	(-2,723)	(-0,304)	(-2,516)	(3,895)	(-2,156)
oil_t	-0,027*	0,007**	0,000	-0,019	-0,047**	0,010
	(-1,694)	(-2,127)	(-0,010)	(-0,289)	(-2,047)	(0,755)
gdp_t	-0,016	0,058***	-0,002	-0,116**	-0,020**	-0,009**
	(-0,498)	(-2,709)	(-0,178)	(-2,295)	(-2,550)	(-2,448)
r_t	0,041**	-7,239*	1,306	-10,805***	-0,531	1,576***
	(2,049)	(-1,849)	(0,857)	(-5,393)	(-1,043)	(3,303)

Note: *, ** and *** denote significance level at 10%, 5% and 1% respectively. t-stat are in parentheses.

Also, it is important to assess the dynamics of adjustment to the long-run equilibrium on consumer prices equations. For example, in France, the error correction mechanism containing cpi_t enters its own equation with a coefficient of adjustment equal to -0,042 and a highly significant t-statistic of -8,395. This means that when consumer prices exceed their long-run equilibrium level, they adjust downwards at a rate of 4.2% per quarter until equilibrium is restored. This implies a long period of half-life adjustment which is equal to four years.[11] For information, the half-life measures are calculated as follows: for consumer prices equation in France, we see that adjustment coefficient is 0.042. We know that $(1-0.042)^n = 0.5$, where n is the number of periods in the half-life of deviations of cpi_t from equilibrium. Taking natural logs of both sides of the equation and rearranging gives $n = (\ln 0.5/\ln 0.96) \simeq 16$ (quarters). According to Table 3, a similar slow adjustment of consumer prices towards equilibrium is found across our sample of EA countries.[12] In fact, this slow adjustment would explain why ERPT coefficients are very weak in the short-run, as reported in the literature.

[11] The so-called half-life is defined as the expected time to revert half of its deviation from the long-run equilibrium.

[12] The faster adjustment is found in Ireland with a half-life measure of one and a half years.

Final step in our cointegration analysis, we turn to the number of restrictions on the long-run parameters postulated in section 2. Thus, we explore the hypotheses of full ERPT ($\mathbf{H}_1$ and $\mathbf{H}_2$) and Zero ERPT ($\mathbf{H}_3$ and $\mathbf{H}_4$). Regarding the tests of restrictions on the long-run parameters in Table 4, it is clear that $\mathbf{H}_1$ is rejected for 9 out of 12 EA countries, implying that EPRT is not complete for this sub-sample. However, We cannot reject the hypothesis of full pass-through for Portugal and Greece. These findings provide corroboration for our earlier empirical results that these two countries have the highest degree of ERPT in our sample of 12 EA.

Table 4: Restrictions on long-run parameters

Country	Full ERPT		Zero ERPT	
	H_1	H_2	H_3	H_4
Austria	8,937	33,623	1,230	27,519
	[0,030]	[0,000]	[0,267]	[0,000]
Belgium	8,396	8,478	3,290	11,780
	[0,004]	[0,076]	[0,070]	[0,019]
Germany	3,684	42,011	2,284	44,954
	[0,055]	[0,000]	[0,103]	[0,000]
Spain	3,793	18,468	2,568	19,335
	[0,051]	[0,001]	[0,109]	[0,017]
Finland	7,708	51,644	2,284	15,558
	[0,005]	[0,000]	[0,131]	[0,004]
France	13,753	26,356	1,475	20,917
	[0,000]	[0,000]	[0,225]	[0,001]
Greece	1,157	12,883	4,210	15,960
	[0,282]	[0,012]	[0,040]	[0,003]
Ireland	3,342	28,287	2,891	12,041
	[0,068]	[0,000]	[0,089]	[0,017]
Italy	10,882	20,073	5,005	9,750
	[0,001]	[0,000]	[0,025]	[0,045]
Luxembourg	2,744	17,919	3,768	17,696
	[0,098]	[0,001]	[0,052]	[0,001]
Netherlands	7,422	39,723	4,918	36,272
	[0,006]	[0,000]	[0,027]	[0,000]
Portugal	1,524	8,829	2,593	20,463
	[0,294]	[0,066]	[0,094]	[0,001]

Note: Restrictions based on Likelihood Ratio tests with a chi-squared distribution, with the number of degrees of freedom equal to the number of restrictions imposed;*p*-values in square brackets.

$\mathbf{H}_2$ is also rejected but for all countries, indicating that complete ERPT is rejected when other variables in the system (oil prices, real GDP and interest rate) are constrained to have no effect on domestic consumer prices. Concerning $\mathbf{H}_3$, the hypothesis of null ERPT is rejected for all EA countries except for Austria, Finland and France. For the latter countries, the weakness of degree of pass-through was confirmed throughout the empirical literature. For instance, GAGNON and IHRIG (2004) found the lowest ERPT elasticity in Finland with a coefficient equal to 0.01%. Finally, the hypothesis of zero ERPT when other variables in the system are constrained to have no effect on consumer prices, namely $\mathbf{H}_4$, is rejected for the whole of our EA sample.

5 Conclusion

In this paper we analyze the pass-through of exchange rate to consumer prices for 12 EA countries within a CVAR framework. Using quarterly data ranging from 1980:1 to 2010:4, our study provides new up-to-date estimates of ERPT with paying attention to either the time-series properties of data and variables endogeneity. Using the Johansen cointegration procedure, results indicate the existence of one cointegrating vectors at least for each EA country of our sample. When measuring the long-run effect of exchange rate changes on consumer prices, we found a wide dispersion of ERPT rates across countries. The degree of ERPT appears to be most prevalent in Portugal and Greece. For Portugal, a 1% depreciation of exchange rate increases domestic consumer prices by roughly 0.84%, while for Greece, consumer prices rise by 58% following one percent depreciation of exchange rate. While the lowest coefficients of long-run ERPT were found in Germany, Finland and France (not exceeding 0.20%). It is important to note that the higher pass-through coefficients in Greece and Portugal were confirmed in the empirical literature (see GAGNON and IHRIG (2004)). Besides, when assessing the adjustment coefficients, we point out a very slow adjustment of consumer prices towards their long-run equilibrium is found across EA countries. This would explain the weakness of ERPT estimates in the short-run in the presence of price stickiness.

References

BACCHETTA, P. and E. VAN WINCOOP (2002), “Why do consumer prices react less than import prices to exchange rates?”, Working Paper No. 9352, NBER.

BAILLIU, J. and E. FUJII (2004), “Exchange rate pass-through and the inflation environment in industrialized countries: An empirical investigation”, Working Paper No. 2004-21, Bank of Canada.

CAMPA, J. and L. GOLDBERG (2002), “Exchange rate pass-through into import prices: A macro or micro phenomenon?”, *NBER Working Paper No. 8934*.

CAMPA, J. and L. GOLDBERG (2005), “Exchange rate pass-through into import prices”, *The Review of Economics and Statistics*, 87 (4), 679–690.

CANDELON, B. and H. LUTKEPOHL (2001), “On the reliability of chow-type tests for parameter constancy in multivariate dynamic models”, *Economics Letters*, 73, 155–160.

CA’ZORZI, M., E. KAHN, and M. SÁNCHEZ (2007), “Exchange rate pass-through in emerging markets”, *ECB Working Papers Series 739*.

CHOUDHRI, E., H. FARUQUEE, and D. HAKURA (2005), “Explaining the exchange rate pass-through in different prices”, *Journal of International Economics*, Vol 65, 349–374.

ELLIOTT, G., T. ROTHENBERG, and J. STOCK (1996), “Efficient tests for an autoregressive unit root”, *Econometrica*, Vol 64 (4), 813–836.

FARUQEE, HAMID (2006), “Exchange rate pass-through in the euro area”, *IMF Staff Papers*, 53, 63–88.

GAGNON, J. and J. IHRIG (2004), “Monetary policy and exchange rate pass-through”, *International Journal of Finance and Economics*, 9 (4), 315–38.

GOLDBERG, P.K. and M. KNETTER (1997), “Goods prices and exchange rates: What have we learned?”, *Journal of Economic Literature*, 35, 1243–72.

HAHN, E. (2003), “Pass-through of external shocks to euro area inflation”, Working paper no. 243, European Central Bank.

HANSEN, H. and S. JOHANSEN (1999), “Some tests for parameter constancy in the cointegrated var”, *Econometrics Journal*, Vol 2, 306–333.

HÜFNER, F.P. and M. SCHRÖDER (2002), "Exchange rate pass-through to consumer prices: A european perspective", Discussion paper no. 02-20, ZEW Centre for European Economic Research.

ITO, T. and K. SATO (2008), "Exchange rate changes and inflation in post-crisis asian economies: Vector autoregression analysis of the exchange rate pass-through", *Journal of Money, Credit and Banking*, 40 (7), 1407–1438.

JOHANSEN, S. and K JUSELIUS (1990), "Maximum likelihood estimation and inference on cointegration - with applications to the demand for money", *Oxford Bulletin of Economics and Statistics*, Vol 52, 169–210.

JOHANSEN, S. and K. JUSELIUS (1992), "Testing structural hypotheses in a multivariate cointegration analysis of the ppp and the uip for uk", *Journal of Econometrics*, Vol 53, 211–244.

MCCARTHY, J. (2007), "Pass-through of exchange rates and import prices to domestic inflation in some industrialized economies", *Eastern Economic Journal*, 33 (4), 511–537.

PARSLEY, D.C. and H.A. POPPER (1998), "Exchange rates, domestic prices, and central bank actions: recent u.s. experience", *Southern Economic Journal*, Vol 64 (4), 957–972.

Appendix A. Specification Tests

Table 5: Results of the Unit Root Tests

Country	ADF		KPSS		DF-GLS	
	Level	1^{st} diff.	Level	1^{st} diff.	Level	1^{st} diff.
CPI						
Austria	-2,512	-3,0937*	0,518243**	1,039	-1,154	-3,011*
Belgium	-0,972	-4,1361**	0,414952**	0,120	-1,246	-2,820*
Germany	-1,372	-3,1581*	0,267249**	0,085	-1,843	-3,616**
Spain	-1,667	-4,5183**	0,255063**	0,104	-0,996	-2,757*
Finland	-0,923	-3,4638**	0,556912**	0,069	-1,165	-3,239*
France	-2,084	-5,0115**	0,474786**	0,298	-1,619	-2,038*
Greece	-1,515	-3,051*	0,631763**	0,124	-1,445	-2,894*
Ireland	-1,084	-1,115	0,306229**	0,166	-1,070	-2,126*
Italy	-2,125	-3,1928*	0,547267**	0,086	-1,565	-2,864**
Luxembourg	-1,062	-4,9549**	0,329039**	0,160	-1,693	-1,404
Netherlands	0,192	-4,5356**	0,251542**	0,491894*	-1,436	-2,966*
Portugal	-1,796	-4,6488 **	0,304176**	0,558627*	-1,505	-2,245
Nominal Effective Exchange Rate						
Austria	-0,951	-8,5404**	0,561378**	0,373	-1,244	-4,692**
Belgium	-2,931	-7,0648**	0,187107*	0,175	-1,700	-2,953**
Germany	-2,129	-8,4702 **	0,393234**	0,086	-2,535	-4,784**
Spain	-3,296	-7,4751 **	0,200878*	0,384	-1,121	-3,282**
Finland	-2,243	-7,7206**	0,326954**	0,066	-2,829	-4,158**
France	-1,953	-8,9225 **	0,210352*	0,337	-1,466	-3,296**
Greece	-0,771	-8,3779**	0,619331**	0,207	-0,529	-2,045*
Ireland	-2,027	-8,3887 **	0,214021*	0,203	-1,400	-2,801**
Italy	-1,904	-7,5047 **	0,349329**	0,102	-1,266	-4,457**
Luxembourg	-1,763	-7,2815**	0,289906**	0,113	-1,683	-3,355**
Netherlands	-1,916	-8,0551**	0,227835**	0,120	-2,459	-4,964**
Portugal	-1,942	-5,8794**	0,352049**	0,143	-1,383	-2,469
GDP						
Austria	-2,913	-7,9166**	1,810626**	0,121	-2,158	-2,816**
Belgium	-0,334	-5,0765 **	2,499921**	0,130	-1,930	-3,671**
Germany	-1,198	-7,6517 **	2,139009**	0,082	-2,602	-4,124**
Spain	-0,979	-2,9660*	2,401370**	0,154	-2,238	-2,844**
Finland	-0,418	-8,5876**	2,276504**	0,098	-2,557	-2,466*
France	-1,215	-4,8021**	2,410361**	0,163	-2,090	-2,495*
Greece	0,108	-3,9337 **	2,262626**	0,368	-0,792	-4,126**
Ireland	-0,717	-3,1905*	2,405650**	0,419	-1,583	-2,658**
Italy	-1,873	-6,9938**	2,355765**	0,305	-0,941	-3,794**
Luxembourg	-0,723	-10,6741**	2,409973**	0,166	-1,799	-3,460**
Netherlands	-0,199	-10,0595**	2,427777**	0,288	-1,507	-2,720**
Portugal	-1,194	-3,7643**	2,376592**	0,395	-1,911	-2,143*

Note: The tests were performed on the logs of the series (except interest rates) for levels including an intercept and trend. The critical values at 1% and 5% levels respectively are: ADF: -3.99, -3.43; KPSS: 0.216, 0.146; DFGLS: -3.48, -2.89. For the first-differences, the tests included only an intercept and were based on the following critical values at the 1%, 5%, and 10% levels respectively: ADF: -3.46, -2.88; KPSS: 0.739, 0.463; DFGLS: -2.58, -1.95. ** and * respectively refer to significance at the 1% and 5%.

Table 5: *Continued*

Country	ADF		KPSS		DF-GLS	
	Level	1^{st} diff.	Level	1^{st} diff.	Level	1^{st} diff.
Interest Rate						
Austria	-3,129	-6,3400**	1,574195**	0,084	-1,568	-2,144*
Belgium	-2,081	-6,2442 **	0,223654**	0,054	0,271	-3,309**
Germany	-1,957	-5,3625 **	1,476766**	0,096	-1,513	-4,384**
Spain	-1,405	-5,7984 **	2,354212**	0,092	-0,032	-4,463**
Finland	-0,988	-7,4761**	2,146838**	0,091	-0,770	-4,944**
France	-1,054	-6,3973 **	2,201712**	0,067	-0,100	-4,562**
Greece	0,321	-4,7748**	2,246433**	0,227	-0,166	-4,951**
Ireland	-1,857	-7,7125**	2,214257**	0,051	0,002	-4,256**
Italy	-0,957	-5,6865**	2,371547**	0,115	0,082	-2,107*
Luxembourg	-2,081	-6,2442**	2,156896**	0,054	0,271	-3,309**
Netherlands	-2,028	-6,1353**	1,681538**	0,054	-0,222	-5,270**
Portugal	-1,054	-5,6402**	2,262505**	0,109	-0,616	-4,737**
Foreign costs						
Austria	-1,902	-8,9765**	0,431481**	0,090	-1,871	-5,005**
Belgium	-3,247	-7,3116**	0,118	0,087	-1,582	-3,085**
Germany	-1,917	-9,2879 **	0,348081**	0,072	-2,476	-4,254**
Spain	-2,518	-7,1836**	0,140	0,253	-1,681	-3,368**
Finland	-1,950	-8,0770**	0,321281**	0,302	-2,398	-3,713**
France	-3,376	-8,0228 **	0,128	0,095	-1,466	-3,866**
Greece	-0,771	-9,1764**	0,632549**	1,177470**	-0,382	-2,045*
Ireland	-2,377	-9,2779**	0,383129**	0,306	-1,089	-4,729**
Italy	-1,873	-7,9844**	0,465492**	0,430	-0,922	-4,490**
Luxembourg	-3,060	-8,0464**	0,113	0,076	-2,093	-3,279**
Netherlands	-1,645	-8,9249**	1,443464**	0,090	-2,606	-4,025**
Portugal	-2,045	-6,4277**	0,540070**	0,251	-1,332	-2,501
Oil Price Index	-1,778	-9,4680**	0,545785**	0,200	-1,111	-3,199**

Note: The tests were performed on the logs of the series (except interest rates) for levels including an intercept and trend. The critical values at 1% and 5% levels respectively are: ADF: -3.99, -3.43; KPSS: 0.216, 0.146; DFGLS: -3.48, -2.89. For the first-differences, the tests included only an intercept and were based on the following critical values at the 1%, 5%, and 10% levels respectively: ADF: -3.46, -2.88; KPSS: 0.739, 0.463; DFGLS: -2.58, -1.95. ** and * respectively refer to significance at the 1% and 5%.

Table 6: Lag selection for CVAR model

Lag Order	**Austria**	**Belgium**	**Germany**	**Spain**	**Finland**	**France**
0	-707,649	-1719,496	-1802,515	-981,602	-91,776	-1116,134
1	-3116,696	-4064,421	-4674,068	-3685,959	**-2301,892**	-3608,798
2	**-3167,887**	**-4098,574**	**-4704,113**	-3721,812	-2297,081	**-3643,728**
3	-3140,334	-4050,572	-4697,863	**-3741,014**	-2294,245	-3627,965
4	-3119,215	-3937,324	-4654,755	-3684,688	-2294,572	-3603,182
5	-3159,858	-3818,671	-4575,063	-3637,260	-2288,652	-3587,996
6	-3142,221	-3660,384	-4455,447	-3543,071	-2237,942	-3532,887
7	-3095,840	-3463,321	-4295,768	-3401,457	-2179,878	-3483,165
8	-3047,109	-3224,676	-4112,951	-3182,100	-2126,362	-3414,625
Lag Order	**Greece**	**Ireland**	**Italy**	**Luxembourg**	**Netherlands**	**Portugal**
0	45,165	-74,602	-328,235	-338,479	-555,000	52,303
1	-1992,475	-2052,221	-2647,038	-2441,852	-2654,382	-2381,231
2	-2035,218	-2076,474	**-2683,614**	-2445,347	**-2681,073**	**-2424,304**
3	-2081,116	**-2085,094**	-2676,574	**-2457,898**	-2657,344	-2417,389
4	-2055,176	-2053,489	-2653,193	-2419,032	-2631,082	-2387,883
5	**-2088,453**	-1996,437	-2614,111	-2372,169	-2624,315	-2327,063
6	-2052,830	-1937,264	-2570,460	-2321,386	-2584,743	-2275,389
7	-2002,637	-1866,812	-2521,844	-2264,651	-2537,194	-2213,436
8	-1926,744	-1840,338	-2471,435	-2205,522	-2469,914	-2171,147

Note: The minimum of the AIC values are in bold.

Table 7: Johansen Trace Test

H_0: *rank* $= r$	**Austria**	**Belgium**	**Germany**	**Spain**	**Finland**	**France**
0	98,540	126,785	150,707	211,569	217,555	165,362
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1	41,959	59,014	76,514	108,246	50,993	96,584
	(0,160)	(0,003)	(0,002)	(0,000)	(0,023)	(0,000)
2	13,324	23,376	43,402	24,111	19,720	51,757
	(0,875)	(0,236)	(0,043)	(0,460)	(0,453)	(0,006)
3	3,048	6,922	13,679	12,779	6,436	21,500
	(0,957)	(0,593)	(0,688)	(0,390)	(0,649)	(0,173)
4	0,411	2,339	2,708	3,425	0,061	7,596
	(0,521)	(0,126)	(0,896)	(0,515)	(0,805)	(0,306)
H_0: *rank* $= r$	**Greece**	**Ireland**	**Italy**	**Luxembourg**	**Netherlands**	**Portugal**
0	135,823	114,898	125,457	81,370	113,458	164,218
	(0,000)	(0,000)	(0,000)	(0,004)	(0,001)	(0,000)
1	85,627	62,078	74,453	47,904	57,245	93,617
	(0,007)	(0,001)	(0,004)	(0,048)	(0,314)	(0,000)
2	47,398	28,905	43,141	21,175	26,916	50,983
	(0,125)	(0,064)	(0,046)	(0,357)	(0,794)	(0,005)
3	25,689	5,763	22,141	4,666	12,708	19,253
	(0,236)	(0,725)	(0,137)	(0,840)	(0,784)	(0,272)
4	12,220	0,073	4,760	1,932	0,060	2,990
	(0,172)	(0,787)	(0,636)	(0,165)	(1,000)	(0,867)

Note: *p*-value are in parentheses.

Appendix B. Chow tests for multiple time series systems

CANDELON and LUTKEPOHL (2001) consider two versions of Chow tests, sample-split (SS) tests and break-point (BP) tests. The BP Chow test for checking for a structural break in period T_B proceeds as follows. The model under consideration is estimated from the full sample of T observations and from the first T_1 and the last T_2 observations, where $T_1 < T_B$ and $T_2 = T - T_B$. The resulting residuals are denoted by $\hat{u}_t$, $\hat{u}_t^{(1)}$ and $\hat{u}_t^{(2)}$, respectively. Using the notation $\tilde{\Sigma}_u = T^{-1} \sum_{t=1}^{T} \hat{u}_t \hat{u}_t'$, $\tilde{\Sigma}_{1,2} = (T_1 + T_2)^{-1} (\sum_{t=1}^{T_1} \hat{u}_t \hat{u}_t' + \sum_{t=T-T_2+1}^{T} \hat{u}_t \hat{u}_t')$, $\tilde{\Sigma}_{(1)} = T_1^{-1} \sum_{t=1}^{T_1} \hat{u}_t^{(1)} \hat{u}_t^{(1)'}$ and $\tilde{\Sigma}_{(2)} = T_2^{-1} \sum_{t=T-T_2+1}^{T} \hat{u}_t^{(2)} \hat{u}_t^{(2)'}$, the BP test statistic is:

$$\lambda_{BP} = (T_1 + T_2) \log\det \tilde{\Sigma}_{1,2} - T_1 \log\det \tilde{\Sigma}_{(1)} - T_2 \log\det \tilde{\Sigma}_{(2)} \approx \chi^2(k), \qquad (6)$$

where k is the difference between the sum of the number of parameters estimated in the first and last subperiods and the number of parameters in the full sample model. Note that also the potentially different parameters in the white noise covariance matrix are counted. The null hypothesis of constant parameters is rejected if λ_{BP} is large.

$$\lambda_{SS} = (T_1 + T_2)[\log\det \tilde{\Sigma}_{1,2} - \log\det (T_1 + T_2)^{-1} (T_1 \tilde{\Sigma}_{(1)} + T_2 \tilde{\Sigma}_{(2)})] \approx \chi^2(k^-) \quad (7)$$

Here k^- is the difference between the sum of the number of coefficients estimated in the first and last subperiods and the number of coefficients in the full sample model, not counting the parameters in the white noise covariance matrix.

CANDELON and LUTKEPOHL (2001) pointed out that especially for multivariate time series models the asymptotic χ^2-distribution may be a poor guide for small sample inference. Even adjustments Based on F approximations can lead to distorted test sizes. Therefore, they have proposed using bootstrap versions of the Chow tests to improve their small sample properties. They are computed as follows. From the estimation residuals $\hat{u}_t$, centered residuals $\hat{u}_1 - \bar{\hat{u}}, \ldots, \hat{u}_T - \bar{\hat{u}}$ are computed. Bootstrap residuals $u_1^*, \ldots, u_T^*$ are generated by randomly drawing with replacement from the centered residuals.

Based on these quantities, bootstrap time series are calculated recursively starting from given pre-sample values $y_{p+1}, \ldots, y_0$. Then the model is reestimated with and without allowing for a break and bootstrap versions of the statistics of interest, say λ_{BP}^* and λ_{SS}^* are computed. The p-values of the tests are estimated as the proportions of values of the bootstrap statistics exceeding the corresponding test statistic based on the original sample.

Table 8: Chow test for VECM

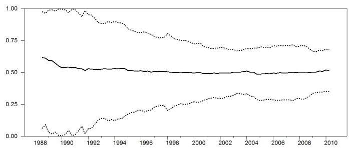
Chow Test	**Austria**	**Belgium**	**Germany**	**Spain**	**Finland**	**France**
Break Point test	741,118	1038,114	590,512	783,370	550,273	928,758
bootstrapped p-value	(0,000)	(0,020)	(0,000)	(0,000)	(0,000)	(0,070)
Sample Split test	481,848	460,226	419,039	495,692	402,790	639,689
bootstrapped p-value	(0,000)	(0,010)	(0,020)	(0,000)	(0,010)	(0,000)
Chow Test	**Greece**	**Ireland**	**Italy**	**Luxembourg**	**Netherlands**	**Portugal**
Break Point test	896,897	663,445	414,849	333.706	755,095	759,694
bootstrapped p-value	(0,100)	(0,000)	(0,000)	(0,030)	(0,000)	(0,000)
Sample Split test	576,060	367,118	410.522	214.397	467,483	448,913
bootstrapped p-value	(0,000)	(0.090)	(0.000)	(0,040)	(0,000)	(0,000)

Note: Bootstrapped p-values are obtained from 1000 bootstrap replication.

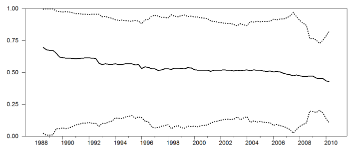
Appendix C. Recursive Analysis of Eigenvalues

A variety of diagnostic tools can be used to investigate parameter constancy by means of recursive estimation as proposed by HANSEN and JOHANSEN (1999). Starting from a base sample $X_{-k+1},\dots,X_{T_0}$, the eigenvalues are calculated recursively for increasing samples $X_{-k+1},\dots,X_t$ for $t = T_0+1,\dots,T$ based upon which the diagnostic tests are calculated. In Figure 2, We report the plots of time paths the largest eigenvalue of the unrestricted VAR model for each country. Non-constancy of β_i or α_i will be reflected in the eigenvalue λ_i.

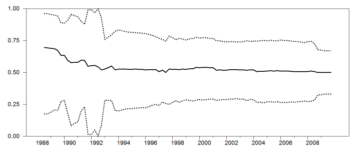
Figure 2: Time paths of eigen values with 95% confidence bands



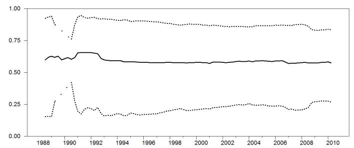
Austria



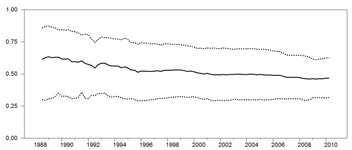
Belgium



Deutschland



Spain



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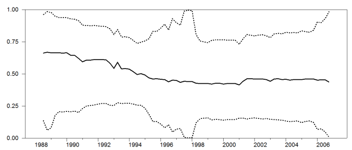
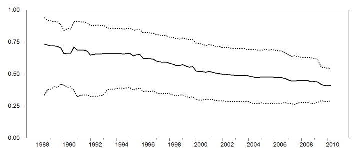
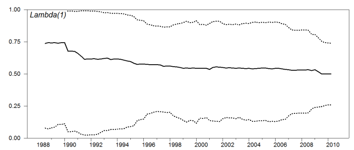


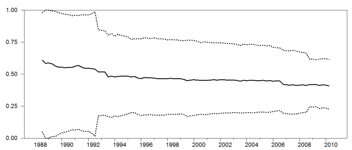
France



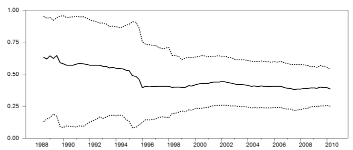
Greece



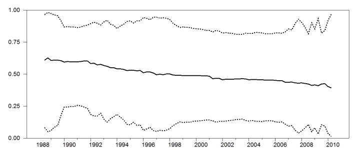
Ireland



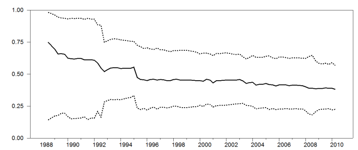
Italy



Luxembourg



Netherlands



Portugal