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The Time Series Properties of the Real Exchange Rates Between the Member States of the European Monetary Union

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Abstract

The article "The Time Series Properties of the Real Exchange Rates Between the Member States of the European Monetary Union" analyses the time series behavior of the components of the real exchange rates between the founding member states of the EMU before and after the start of the EMU. Various panel and univariate country-specific tests show that the levels of these components are typically random walks. The resulting real exchange rates are also random walks and their components are not cointegrated. It is argued that these results question the operability of the EMU under the current policy regime in the long-run. One possibility to deal with this problem could be the suspension of the principle of a "single monetary policy".

Die Zeitreiheneigenschaften der realen Wechselkurse der Mitgliedsländer der Europäischen Währungsunion

Zusammenfassung

Der Artikel „Die Zeitreiheneigenschaften der realen Wechselkurse der Mitgliedsländer der Europäischen Währungsunion" untersucht das Zeitreihenverhalten der Komponenten der realen Wechselkurse zwischen den Gründerstaaten der EWU vor und nach dem Beginn der EWU. Verschiedene Panel- und univariate länderspezifische Test zeigen, dass die Niveaus dieser Komponenten typischerweise Zufallspfaden folgen. Die resultierenden realen Wechselkurse folgen ebenfalls Zufallspfaden und ihre Komponenten sind nicht kointegriert. Diese Ergebnisse, so schließt der Artikel, stellen die langfristige Funktionsfähigkeit der EWU unter dem gegenwärtigen geldpolitischen Regime in Frage. Eine Möglichkeit, dieses Problem zu adressieren, könnte in der Preisgabe des Prinzips der einheitlichen Geldpolitik bestehen.

Keywords: Real Exchange Rate, Purchasing Power Parity, Monetary Policy, European Monetary Union, European Central Bank, Single Monetary Policy, Unit Root Tests, Cointegration Tests, Dutch disease, Deindustrialization

JEL Classification: E50, E31, C12

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I. Introduction

Diverging inflation rates between the member states of a currency union can lead to diverging real interest rates, if monetary policy causes a convergence of nominal interest rates. According to "Walter's Critique[1]" high (low) inflation countries might experience growing (decreasing) inflation rates, since low (high) real interest rates favor spending (saving). This way, a self-reinforcing divergence process might result.[2] *Busetti* et al. (2007) find evidence for a convergence of inflation rates, in terms of stationary inflation differentials for the 12 founding member states of the European Monetary Union (EMU): during the European Exchange Rate Mechansim (ERM) period (1980–97) but the emergence of different "convergence clubs" afterwards (1997–2004). *Karanasos* et al. (2016) find evidence for three convergence clubs before the EMU period (1980Q1–1997Q4) and two convergence clubs and a relatively large group of countries with diverging inflation rates (Ireland, Italy, Greece, Netherlands, Portugal, Spain) for the period afterwards (1998Q1–2013Q4).

In this paper it is questioned whether stationary inflation differentials are sufficient for the long-run stability of a monetary union. Most central banks, practice inflation targeting and not price level targeting. Therefore price indices follow typically a random walk and not a linear time trend. As a result, even in the presence of stationary inflation differentials, a significant long-run divergence of price levels is possible (Figure 1). Such a divergence can cause severe economic imbalances. Price level divergence does not only affect the competitiveness of the prices for goods and services; it also affects the competitiveness of wages and salaries via its impact on unit labor costs. If real wages and the wage share in GDP shall stay constant, nominal wages must follow the price level. This however implies an increase of unit labor costs and reduces therefore wage competitiveness against countries with lower price levels.[3] As discussed in sec-

[1] For a discussion see *Miller/Sutherland* (1991).

[2] In a currency union with free trade, this is possible, if at least a part of the net credits received from the central bank or other member states is used to buy non-tradable domestic goods (*Maurer* 2010).

[3] Constancy of wage share demands:
$\frac{L_t * W_t}{Y_t * P_t} = \frac{L_{t+1} * W_{t+1}}{Y_{t+1} * P_{t+1}} \Leftrightarrow \frac{W_{t+1}}{W_t} = \frac{P_{t+1}}{P_t}\left(\frac{Y_{t+1}}{L_{t+1}} / \frac{Y_t}{L_t}\right)$, i.e. nominal wages must grow with the inflation rate times the real productivity growth rate (where t represents a time index and L the total working hours per year, W the average nominal hourly wage rate, P the GDP-deflator and Y the real GDP.). However constancy of unit labor costs allows nominal wages only to grow with the real productivity growth rate:
$\frac{L_t * W_t}{Y_t * P_T} = \frac{L_{t+1} * W_{t+1}}{Y_{t+1} * P_T} \Leftrightarrow \frac{W_{t+1}}{W_t} = \left(\frac{Y_{t+1}}{L_{t+1}} / \frac{Y_t}{L_t}\right)$, (where T represents a base year). Hence, if high inflation countries keep their wage shares constant, their unit labor cost will grow stronger than the unit labor costs of low inflation countries.

tion 5, this can lead to a shrinking tradable sector and cause a deindustrialization effect similar as in the "Dutch disease" model of *Corden/Neary* (1982).

Given these potential problems caused by price level divergence, this paper analyzes the stationarity of the real exchange rates between the 12 EMU founder states empirically. The analysis is based on panel and single country pairwise unit root and stationarity tests. Since the overall result indicates that real exchange rates are typically not stationary, cointegration tests are additionally applied in order to allow the coefficients of the real exchange rate components to deviate from unity.

The paper is organized in the following way: Section 2 shows, why inflation targeting of central banks can cause nonstationary price levels and why stationarity of inflation differences can come along with nonstationary of price level differences. Section 3 explains the data used for the empirical tests. Section 4 contains the specification of the unit root and cointegration tests and the results. Section 5 discusses policy conclusions.

II. Time Series Properties of Price Levels Under Inflation Targeting

Under normal assumptions, inflation targeting causes price levels to follow a random walk, while price level targeting leads to price levels that are stationary around a linear trend. If monetary policy can reach an inflation target, π^*, only with a certain error margin equal to $\varepsilon_t \sim N(0,\sigma)$, the price level of the following period will equal

$$p_{t+1} = p_t * (1+\pi^*) * (1+\varepsilon_{t+1})$$

$$<=>$$

(1) $$ln(p_{t+1}) = ln(p_t) + \pi^* + \varepsilon_{t+1}$$

under inflation targeting. As a result, for $p_0 = 1$ the price level follows a linear trend plus the sum of historic errors[4]

(2) $$ln(p_{t+1}) = (t+1)\ \pi^* + \sum_{j=0}^{t+1} \varepsilon_j$$

As equation (1) shows, in this case a regression of the type

(3) $$ln(p_{t+1}) = \alpha_1 + (t+1)\ \alpha_2 + \rho\ ln(p_t) + \varepsilon_{t+1}.$$

[4] Where $E_0[ln(p_t)] = t\ \pi^*$ and var($ln(p_t)$) = t σ^2.

will accept the existence of a unit root $\rho = 1$ with the target rate equal to an estimate of the drift parameter $\alpha_1 = \pi^*$ but reject a linear time trend, i.e. $\alpha_2 = 0$, since the previous price level $ln(p_t)$ but not the linear time trend $(t+1)\ \alpha_2$, is better able to account for the errors $\varepsilon_t \sim N(0,\sigma)$.

In case of price level targeting, monetary policy will target in every period the target inflation rate π^*minus the previous target deviation $\zeta_t^* = ln(p_t) - ln(p_t^*)$, with $ln(p_t^*) = ln(p_{t-1}^*) + \pi^*$. As a result, the price level of the next period will equal

$$ln(p_{t+1}) = ln(p_t) + (\pi^* - ln(p_t) + ln(p_t^*)) + \varepsilon_{t+1}$$

$$\Leftrightarrow$$

(4) $$ln(p_{t+1}) = ln(p_t^*) + \pi^* + \varepsilon_{t+1}$$

In this case, for $p_0 = 1$ the price level will then follow a simple linear trend

(5) $$ln(p_{t+1}) = (t+1)\ \pi^* + \varepsilon_{t+1}$$

and a regression of the type

(6) $$ln(p_{t+1}) = \alpha_1 + (t+1)\ \alpha_2 + \rho\ ln(p_t) + \varepsilon_{t+1}$$

will reject the existence of a unit root and instead accept the existence of a linear time trend with the estimate of the trend parameter equal to the target rate $\alpha_2 = \pi^*$ and $\alpha_1 = 0$.[5]

As the empirical analysis in section 4 will show, the null hypothesis of a unit root in the consumer price indices cannot be rejected against the alternative hypothesis of a linear time trend for all 12 EMU founder states over the "Bretton Woods" (1960:1–1972:12): "ERM" (1973:1–1998:12) and "EMU" (1999:1–2017:5) periods.

If price levels follow an I(1) unit root process, inflation difference between EMU member states will be stationary, even if their price levels are not cointegrated and can display unlimited divergence. In this case, the stationarity of inflation differences will be caused by mere differencing of the levels. This follows for a country pair (i, j) difference of the first difference of equation (1): if $\rho_i = \rho_j = 1$ and $|\pi_i^* - \pi_j^*| \geq 0$:

[5] For a similar result see *Svensson* (1999).

$$
\begin{aligned}
(7)\qquad & \Delta ln\left(p_{i,t+1}\right)-\Delta ln\left(p_{j,t+1}\right)\\
&=\pi_i^*-\pi_j^*+\left(\rho_i-1\right)ln\left(p_{i,t}\right)-\left(\rho_j-1\right)ln\left(p_{j,t}\right)+\varepsilon_{i,t+1}-\varepsilon_{j,t+1}\\
&=\pi_i^*-\pi_j^*+\varepsilon_{i,t+1}-\varepsilon_{j,t+1}
\end{aligned}
$$

These kind of pairwise inflation differentials can only be non-stationary, if price levels are I(2). This is however typically not the case as section 4 shows. Therefore, the inflation rate convergence results of *Busetti* et al. (2007) and *Karanasos* et al. (2016) may simply be the result of first-differencing but not of any meaningful economic mechanism (cp. the simulation in Figure 1).

To test for the existence of an economic mechanism, which keeps the price levels within a certain bandwidth, the stationarity of the real exchange rates or the degree of cointegration of the real exchange rate components have to be examined. Such an economic mechanism could be price arbitrage of goods and/or factor mobility.

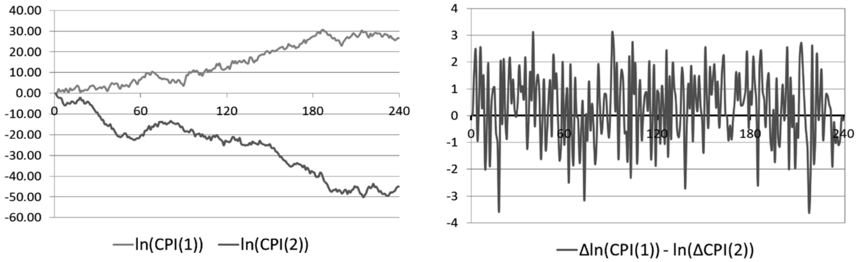

Figure 1: Price Level Divergence in the Presence of Stationary Inflation Differentials[6]

III. Data

The tests are based on monthly consumer price index (CPI) and exchange rate data from the OECD Main Economic Indicators Database (OECD 2010) as provided by the Federal Reserve Economic Database (FRED 2017). The base year for the CPI is 2010. The analysis focuses on the 12 founding states of the EMU. To compare the EMU period with pre-EMU periods national price indices are used as collected by national statistical offices reaching from 1960-01-01 to

[6] The following simulation model is used in Figure 2:

$ln\left(p_{i,t+1}\right)=\alpha_{i,1}+\rho_i\ ln\left(p_{i,t}\right)+\varepsilon_{i,t+1}$ with (i, j) = (1, 2): $\alpha_{i,1}=\frac{0.02}{12}$, $\rho_i=1$ and $\varepsilon_t\sim N(0,1)$.

The strongest deviation of price levels between country 1 and 2 over a series of 20 trials has been selected, to demonstrate that strong deviations of price levels are compatible with stationary inflation differentials between both countries. An Excel file of the simulation model is available on request.

2017-05-01. The sample is divided in the Bretton-Woods subsample (Bretton Woods period): 1960:1 to 1972:12, the European Exchange Rate Mechanism subsample (ERM period): 1973:1 to 1998:12 and the European Monetary Union subsample (EMU period): 1999:1 to 2017:5.

IV. Test Specification and Results

1. *Consumer Price Index Unit Root Tests*

To test the null hypothesis (H0) of a unit root possibly with drift in the CPI levels against the alternative hypothesis (H1) of stationary process around linear trend, panel tests as well as univariate "country-specific" tests are applied in the following. According to conventional wisdom, the power of panel test is higher than the power of univariate tests. However, as demonstrated by *Banerjee* et al. (2005): in the presence of cross-unit cointegrating relationships, the H0 of a unit root is rejected too often. Since price levels and exchange rates in a context of increasingly globalized financial markets are likely to suffer from such cross-unit correlations, it seems to be appropriate to apply also country-specific unit root tests.

As the time dimension T of the data used here is larger than the panel dimension N, panel tests are appropriate, which have asymptotics that rely on the assumption that first the time dimension goes to infinity and then the number of panels. Therefore the Breitung (*Breitung* 2000), Herwartz test (*Herwartz* et al. 2018) and the Hadri tests (*Hadri* 2000) are chosen. The Breitung test exhibits also higher power than the conventional Levin–Lin–Chu test (*Levin* et al. 2002) and the *Im-Pesaran-Shin* test (*Im* et al. 2003) in cases where the H1 relies on autoregressive parameters (the ρ in equation (1)) near unity and when panel-specific means and trends are included (*Breitung* 2000). It also has a good power in small datasets, as long as N is not growing faster than T. To deal with the problem of cross-sectional correlation the *Breitung/Das* (2005) test version is used here, which is robust to cross-sectoral correlation. One problem from which the *Levin* et al. (2002), the Im-Pesaran-Shin and the Breitung-Das tests suffer, is their dependence on the assumption of homoskedastic panels. This assumption is typically not fulfilled for most macroeconomic time series (*Westerlund* 2015). While the panel unit root test of *Westerlund* (2014) is heteroscedasticity-robust in the presence of an intercept term, it is not robust if the time series contains a linear trend. A linear time trend is however the alternative hypothesis to the null hypothesis of a unit root in the consumer price index, according to the theoretical considerations of section 2. Therefore a test proposed by *Herwartz* et al. (2017), which is also heteroscedasticity-robust for time series with linear trends, is additionally used in the following.

Furthermore, the Hadri Likelihood Multiplier test (*Hadri/Kurozumi* 2012) is applied, which relies on the opposite H0 that all panels are stationary, while the H1 states that at least one panel contains a unit root. This twist of the H0 is typically used to control for the problem that unit-root tests have a low power against alternative hypotheses of slightly persistent but nevertheless stationary processes (a ρ smaller but close to unity). The Hadri test relies however on the homoscedasticity assumption.

Table 1 presents the results of the panel tests. The Breitung-Herwartz tests do not reject the H0 of a unit root process with a possible drift parameter in all panels against the H1 of a stationary process around a linear trend in all three sub-sample periods. Correspondingly the Hadri test rejects the H0 of trend stationarity in all panels against the H1 of at least one panel with a unit root. Consequently, the panel test results are in line with the hypothesis derived in section 2 that inflation targeting, causes unit roots but not trend stationary processes in CPI levels. Appendix table 1 shows that applying the same tests to the first differences of the CPIs, indicates that the CPIs of some countries, especially in the ERM period might be integrated of a higher order than I(1). The country-specific tests provide further insights.

Table 1

Panel Unit Root Tests for of Consumer Price Index Levels

Test Specification					Results											
					1960:1–1972:12				1973:1–1998:12				1999:1–2017:5			
Test	H0	H1	Autoregression Parameter ρ	Lags / Selection	Panels	Periods	P-value	Accepted Hypothesis	Panels	Periods	P-value	Accepted Hypothesis	Panels	Periods	P-value	Accepted Hypothesis
Breitung	All Panels contain unit roots	All Panels are stationary	Uniform ρ	12	12	156	0,996	H0	12	311	0,560	H0	12	222	0,573	H0
Herwartz	Panels contain unit roots	Panels are stationary	Panel-specific ρ	AIC	12	156	0,715	H0	12	311	0,909	H0	12	222	0,498	H0
Hadri	All panels are stationary	Some panels contain unit roots	–	12	12	156	0,000	H1	12	311	0,000	H1	12	222	0,000	H1

The significance level for the rejection of the H0 is 5%. All Panels as strongly balanced. Panel-specific linear trends are allowed. In all tests a correction for cross-sectional dependence of the panels is applied. A *Bartlett* Kernel with 12 lags is used to estimate the long-run variance in the Hadri tests.

The country-specific unit root tests are based on the Augmented Dickey-Fuller test (ADF) (*Dickey/Fuller* 1979) and the Phillips-Perron test (PP) (*Phillips/Perron* 1988). The PP test is robust with respect to unspecified autocorrelation and heteroscedasticity in the disturbance process. However, according to a study of *Davidson/MacKinnon* (2004), a PP test performs worse than the ADF test in finite samples. Since both tests have low power to reject the H0 in the presence of structural breaks, a Zivot-Andrews test (ZA) (*Andrews/Zivot* 1992) with the H0 of a unit root and the H1 of a stationary process with a structural break in the intercept and the linear trend is also applied. As an analog to the Hadri panel test, the Kwiatkowski–Phillips–Schmidt–Shin tests (KPSS) (*Kwiatkowski* et al. 1992) of the H0 of a stationary process around a deterministic trend against the H1 of a unit root is used.

The results of these tests are displayed in appendix table 2 for the CPI levels. Since monthly CPI data are used and consumption behavior follows a seasonal pattern typically, the tests are applied to seasonally adjusted and unadjusted data. Seasonal adjustment is based on the Holt-Winters seasonal smoothing method. If seasonally adjusted data lead to the opposite H0-decision, this is indicated in the table with a "Yes" under "Seasonality".

As appendix table 2 shows, the tests results largely confirm to the hypothesis of a unit root possibly with a drift in the CPI levels against the alternative hypothesis of stationarity around a linear trend. The ADF test fails to reject the H0 in one case out of 36 (Austria, 1999:1–2017:5); the PP test fails to reject the H0 in another case (Netherlands, 1973:1–1998:12). Only in one case (Austria, 1999:1–2017:5) the KPSS test and the ADF test reject the unit root hypothesis simultaneously. The KPSS test rejects the unit root hypothesis in two other cases (Finland, 1960:1–1972:12 and 1999:1–2017:5).

As appendix table 3 shows, there is predominantly no rejection for trend stationarity in the first differences of CPIs. The PP test rejects the H0 always; the ADF test rejects the H0 of a unit root however in 16 out of 36 cases. In five cases the KPSS and the ADF test do not reject the unit root hypothesis simultaneously (Germany, 1960:1–1972:12; Germany, 1973:1–1998:12; Greece, 1973:1–1998:12; Greece, 1999:1–2017:5; Netherlands, 1973:1–1998:12). The KPSS test fails to reject the unit root hypothesis in one other case (France, 1973:1–1998:12).

Consequently, the time series behavior of the CPIs of the EMU founder states follows typically an I(1) process. This holds especially for the EMU period, where the PP and the KPSS test never reject the unit root hypothesis for the CPI levels, while for the first differences of CPIs the PP test always rejects the unit root hypothesis and the KPSS test in all but one case (Greece, 1999:1–2017:5). This means, as argued in section 2, that the convergence results of *Busetti* et al. (2007) and *Karanasos* et al. (2016) may indeed be simply the result of first-differencing but not of a meaningful economic mechanism. To test for the existence of an economic mechanism that keeps these non-stationary price levels together, the stationarity of the real exchange rate is analyzed in the following. Can non-stationary price levels of 12 countries result in 66 stationary real exchange rate pairs with waiver of 66 nominal exchange rates pairs as "slack variables" in the EMU period?[7]

[7] The appendix to this paper provides also unit root test for the levels and first differences of the nominal exchange rates over the BW and ERM period (Appendix Tables 4–7). These results do not reject the H0 of an I(1) unit root process with a possible drift

2. *Real Exchange Rate Unit Root Tests*

The natural log of a real exchange rate is calculated according to the formula

$$ln\left(re_t^{j,i}\right) = ln\left(p_t^i\right) - ln\left(p_t^j\right) - ln\left(e_t^{i,j}\right) \tag{8}$$

where $e_t^{i,j}$ equals the nominal exchange rate, i.e. the amount of currency units of country i that are paid on the foreign exchange market for one currency unit of country j such that $re_t^{j,i}$ informs about the natural log of number of the consumer goods baskets of country j that can be bought with one consumer goods basket of country i. In a world without transportation or other transaction costs in trade with goods, the arbitrage mechanism would cause $ln\left(re_t^{j,i}\right)$ to approach zero. In a world with all kind of transaction costs that prevent perfect arbitrage $ln\left(re_t^{j,i}\right)$ can be different from zero. However this difference should not accumulate over time, i.e. $ln\left(re_t^{j,i}\right)$ should not follow a random walk, because in case of too strong deviations from zero, arbitrage dealings in goods or production factors should become profitable despite transaction costs and cause a correction of the deviations. If this mechanism works sufficiently well, the nominal exchange rate $ln\left(e_t^{j,i}\right)$ should in the long-run adjust to compensate for the differences in the development of the price levels $ln\left(p_t^i\right)$ and $ln\left(p_t^j\right)$ such that purchasing power parity (PPP) of the currencies of both countries holds.

This hypothesis is contained in the equation

$$ln\left(re_t^{j,i}\right) = \alpha_1 + \rho\ ln\left(re_{t-1}^{j,i}\right) + \varepsilon_{t+1} \tag{9}$$

and the condition $0 < \rho < 1$. The intercept α_1 represents the size of transaction costs.[8] In this case the long-run intertemporal equilibrium value of $ln\left(re_t^{j,i}\right)$ equals[9] $ln\left(re_*^{j,i}\right) = \alpha_1 / (1-\rho)$. Solving this equation for α_1 and inserting the result in equation (9) and subtracting from both sides $ln\left(re_{t-1}^{j,i}\right) - ln\left(re_*^{j,i}\right)$ yields

$$ln\left(re_t^{j,i}\right) - ln\left(re_{t-1}^{j,i}\right) = (\rho - 1)\left(ln\left(re_{t-1}^{j,i}\right) - ln\left(re_*^{j,i}\right)\right) + \varepsilon_{t+1} \tag{10}$$

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parameter against the H1 of a stationary process around a linear trend. Consequently a combination of I(1) CPIs and I(1) nominal exchange rate has the potential to yield stationary real exchange rates before the EMU period.

[8] Since the CPI data are indexed on the base year 2010, the constant also allows for differences in the absolute price levels of the base year.

[9] Taking the expectation value of equation (9) yields $\mathrm{E}\left(ln\left(re_t^{j,i}\right)\right) = \alpha_1 + \rho\ \mathrm{E}\left(ln\left(re_{t-1}^{j,i}\right)\right)$ and inserting $ln\left(re_*^{j,i}\right)$ yields $ln\left(re_*^{j,i}\right) = \alpha_1 + \rho\ ln\left(re_*^{j,i}\right)$ what equals $ln\left(re_*^{j,i}\right) = \alpha_1 / (1-\rho)$.

This error-correction version of equation (9) shows if $0 < \rho < 1$ any positive [negative] deviation of $ln\left(re_{t-1}^{j,i}\right)$ from its long-run equilibrium value $ln\left(re_{*}^{j,i}\right)$ is followed by a decrease [increase] of $ln\left(re_{t}^{j,i}\right)$ compared to its previous value $ln\left(re_{t-1}^{j,i}\right)$. As a result, if $0 < \rho < 1$ the real exchange rate $ln\left(re_{t}^{j,i}\right)$ fluctuates around a stable long-run equilibrium value $ln\left(re_{*}^{j,i}\right) = \alpha_1 / (1-\rho)$, which depends positively on the measure for transaction costs α_1 and negatively on the speed of adjustment from deviations $(1-\rho)$. If however $\rho = 1$ no error correction takes place and $ln\left(re_{t}^{j,i}\right)$ follows a random walk. Hence equation (9) can be used to test for the long-run validity of purchasing power parity. If the unit root hypothesis $\rho = 1$ cannot be rejected, $ln\left(re_{t}^{j,i}\right)$ follows a random walk with drift α_1 and PPP does not hold, if $\rho = 1$ can be rejected $ln\left(re_{t}^{j,i}\right)$ will follow a stationary process fluctuating around $\alpha_1 / (1-\rho)$ and PPP holds.

Table 2 presents the results for the Breitung, Herwartz and Hadri panel unit root panel tests according to equation (9): i.e. with an intercept α_1 but without allowing for a linear trend. The Breitung test rejects the H0 for the ERM period and the EMU period at a significance level of 5%. At a significance level of 1%, the H0 is rejected for the ERM period only. The heteroscedasticity robust Herwarz test rejects the H0 only for the ERM period at conventional significance levels. The Hadri test always rejects the H0 that all panels are stationary in favor of the H1 that some panels contain a unit root. For the Bretton Woods period the unit root hypothesis is never rejected by any test.

Appendix table 8 shows that the H0 of a unit root is clearly rejected for the first differences of the real exchange rates in favor of the H1 of stationarity without linear trend. The only exception provides the Hadri test for the EMU peri-

Table 2

Panel Unit Root Tests of Real Exchange Rates Levels

Panel Unit Root Tests: Real Exchange Rate																	
Test Specification					Results												
					1960:1–1972:12				1973:1–1998:12				1999:1–2017:5				
Test	H0	H1	Autoregression Parameter ρ	Lags / Selection	Panels	Periods	P-value	Accepted Hypothesis	Panels	Periods	P-value	Accepted Hypothesis	Panels	Periods	P-value	Accepted Hypothesis
Breitung	Panels contain unit roots	Panels are stationary	Uniform ρ	12	66	156	0,982	H0	66	311	0,000	**H1**	66	222	0,038	**H1**
Herwartz et al.(2017)	Panels contain unit roots	Panels are stationary	Panel-specific ρ	AIC	66	156	0,689	H0	66	311	0,003	**H1**	66	222	0,188	H0
Hadri	All panels are stationary	Some panels contain unit roots	–	12	66	156	0,000	H1	66	311	0,000	H1	66	222	0,000	H1

The significance level for the rejection of the H0 is 5%. All Panels as strongly balanced. Panel-specific linear trends are not allowed. In all tests a correction for cross-sectional dependence of the panels is applied. A Bartlett Kernel with 12 lags is used to estimate the long-run variance in the Hadri tests.

od. This supports the view that real exchange rates typically follow a I(1) process.

To check the panel test results with univariate country-specific tests, implies to test for 66 possible country-specific real exchange rate combinations (66 = (11*12)/2). The results are displayed in appendix table 9. Table 3 summarizes the results.

Table 3

Summary of Country Specific Unit Root Tests of Real Exchange Rate Levels

Period	1960:1–1972:12				1973:1–1998:12				1999:1–2017:5			
Test	ADF	PP	KPSS	ZA	ADF	PP	KPSS	ZA	ADF	PP	KPSS	ZA
Rejection of H0	2	1	10	29	6	9	14	9	5	10	2	6

The significance level for the rejection of the H0 of a unit root is 5%.

In most cases the hypothesis of a unit root cannot be rejected at a significance level of 5%. The number of rejections according to the ADF tests and PP tests is 5–10 and roughly the same for the ERM and EMU period. For the Bretton-Woods period the ZA test indicates that structural breaks play an important role in the very low level of H0 rejections by ADF and PP tests. The ZA tests indicate that most of these breaks took place at the turbulent end of the Bretton-Woods system or at points in time when nominal exchange rate adjustments took place.

During the EMU period, the PP test rejects the H0 most often (10 times). This is the case for the following country pairs: Belgium/Austria, Luxembourg/Austria, Luxembourg/Belgium, Ireland/France, Ireland/Germany, Portugal/Germany, Spain/Germany, Spain/Greece, Luxembourg/Italy, Netherlands/Luxembourg. This could indicate that countries with high openness indices[10] like Luxembourg and Germany, tend to have relatively more often stationary real exchange rates with other countries. Interestingly, it is not possible to detect "cross stationarities", in the sense that Austria, Belgium, Italy and the Netherlands, which all have stationary real exchange rates with Luxembourg, have also stationary exchange rates across each other. The same holds for Ireland, Portugal and Spain, which have stationary exchange rates with Germany, but non-stationary exchange rates across each other. Table 4 reveals, that this does not even hold approximately – in terms of relatively lower test statistics. Consequently, it is not possible to detect "stationarity clusters" based on these country-specific unit root tests.[11]

[10] Openness index = (Exports + Imports)/GDP.

[11] As the summary at the end of Appendix Table 10 shows, the H1 of stationarity without a linear trend can typically not be rejected for the for the first differences of the real

Table 4

Check for Potential Cross Stationarity Clusters

	Austria		Belgium		Italy	
	5% sign. level	Test statistic	5% sign. level	Test statistic	5% sign. level	Test statistic
Belgium	**–2,88**	**–3,24**				
Italy	–2,88	–0,71	–2,88	–0,15		
Netherlands	–2,88	–2,00	–2,88	–1,40	–2,88	–2,18

	Ireland		Portugal	
	5% sign. level	Test statistic	5% sign. level	Test statistic
Portugal	–2,88	–1,52		
Spain	–2,88	–0,87	–2,88	–1,62

The significance level for the rejection of the H0 of a unit root is 5%.

In a sense, it is surprising that the results for the EMU period do not systematically deviate from the results of the Bretton-Woods and ERM period, since the transition to a monetary union accompanied by a tariff union and with a far reaching convergence of regulation requirements, such as offered by the single European market, should have facilitated the detection and realization of arbitrage possibilities profoundly. Even differences in labor productivity which according to the Balassa-Samuelson effect (*Kravis/Lipsey* (1991)) can cause non-stationary real exchange rates, does not explain these results in a satisfying way. The Balassa-Samuelson effect would imply that at least structurally alike countries with similar levels of labor productivity would display stationary real exchange rates. Hence according the Balassa-Samuelson effect at least two “stationarity clusters” for “northern” and “southern” EMU member states should exist. But this is not the case. Even such similar countries like France and Germany do not display stationary real exchange rates. To check for the robustness of these results, in the following weaker forms of stationarity are analyzed by testing for cointegration between the components of the real exchange rate.

exchange rates. The only exceptions are the PP and KPSS test for the EMU period, where in 24 (PP) and 20 (KPSS) cases the H0 of a random walk is not rejected. This indicates that real exchange rate follow typically I(1) processes, while for the EMU period in some cases higher integrated processes are possible.

3. *Real Exchange Rate Cointegration Tests*

The calculation of the real exchange rate by equation (8) imposes strong restrictions on the coefficients of the real exchange rate components. In fact this calculation sets the coefficients of all three components to unity $\beta_1 = \beta_2 = \beta_3 = 1$.

$$ln\left(re_t^{j,i}\right) = \beta_1\ ln\left(p_t^i\right) - \beta_2\ ln\left(p_t^j\right) - \beta_3\ ln\left(e_t^{i,j}\right) \tag{10}$$

These restrictions may be too strong in a world with transaction costs that impede arbitrage activities. The cointegration approach allows for less restrictive tests of the hypothesis that the three components of real exchange rates are "kept together" by arbitrage activities. The simplest form of a cointegration test can be based on the Augmented Engle-Granger cointegration test (AEG): which follows a two-step procedure. First, one component of the real exchange rate is regressed on the others using OLS.

$$ln\left(p_t^i\right) = \alpha_1 + \beta_1\ ln\left(p_t^j\right) + \beta_2\ ln\left(e_t^{i,j}\right) + \varepsilon_t \tag{11}$$

Then the residuals of this regression ε_t are tested for stationarity using an ADF test with the critical values derived by *MacKinnon* (1990). If the ADF test rejects the H0 of a unit root in the residuals, the components of the real exchange rate are cointegrated and the OLS estimation of equation (11) has the property of super consistency (*Stock* 1987). Since the Dickey-Fuller test requires that the residuals ε_t are serially uncorrelated, a heteroscedasticity robust version of the Cumby-Huizinga test is used in the following to test the hypothesis that all autocorrelations over a range of 1 to 12 lags are zero (*Cumby/Huizinga* 1992). Starting with an ADF test on the residuals ε_t with zero lags, additional lags are added, as long as the Cumby-Huizinga test does not reject the serial correlation. The AEG test is applied by allowing for a linear trend in the first step of the procedure as well as without a linear trend. The pure form of the PPP hypothesis does not allow for stationarity around a linear trend. But for exploratory reasons it is interesting to see, in how far an additional weakening of the PPP hypothesis has in impact on the results.[12] Appendix Table 11 presents the results for all 66 country pairs. Table 5 summarizes these results.

[12] Another approach to weaken the PPP hypothesis would be the time-varying cointegration coefficients, as used by *Beckmann* et al. (2011) in the analysis of the relationship between the nominal Euro-Dollar exchange rate and its fundamentals. It is very likely that by construction time-varying cointegration coefficients will allow for more cointegration relationships as the constant coefficient method. Especially the analysis of the resulting breakpoints could be a useful tool to better understand the reasons for the lack to price convergence found here. This type of analysis is left open here for future research.

Table 5

Summary of Country Specific Augmented Engle-Granger Cointegration Tests of Real Exchange Rate Components

Period	1960:1–1972:12		1973:1–1998:12		1999:1–2017:5	
Engle-Granger Test	With Trend	Without Trend	With Trend	Without Trend	With Trend	Without Trend
Rejection of No Cointegration Hypothesis	8	7	3	2	1	1

The significance level for the rejection of the H0 of a unit root is 5%.

Somewhat surprisingly, the relaxation of the restrictions of the pure PPP hypothesis does not lead to more rejections of the hypothesis of no PPP but to less rejections, with exception of the Bretton Woods period. Allowing for a trend does also not lead to more rejections. The one case, where no cointegration without a trend is rejected in the EMU period is Italy/Greece.

One explanation for this result could be the fact that the AEG test has some limitations compared to its alternative, the Johansen Cointegration (JC) test (*Johansen* 1995). Although, one can perhaps argue that the AEG test makes fewer distributional assumptions (*Schaffer* 2010): there are certainly a couple of disadvantages. For example, the preliminary unit root tests for all variables, which are not necessary for the JC test, have typically low statistical power, the choice of the dependent variable in equation (11) can influence the test results and at most one cointegrating relationship can be found. Even though the last problem is not relevant for the EMU period, where the real exchange rate has only two components, these problems nevertheless justify an additional application of the JC test.

The JC test uses the framework of a vector autoregression model (VAR) to test for cointegrating relations. In a VAR model all variables depend mutually on their lags:

(12) $$y_t = \Pi_1 \; y_{t-1} + \Pi_2 \; y_{t-2} + \ldots + \Pi_p \; y_{t-p} + \in_t.$$

To test for cointegration of the real exchange rate components, the vector y_t . equals $y_t = \left(p_t^i, \quad p_t^j, \quad e_t^{i,j}\right)'$, where from now on small letters mark the natural logarithm of a variable. In this case the Π_i are (3×3) matrices. This VAR can be equivalently reformulated as a vector error correction model (VEC):

(13) $$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \; \Delta y_{t-i} + \in_t$$

where $\Pi = \Pi_1 + \ldots + \Pi_p - \mathrm{I}$ and $\Gamma_1 = -\Pi_{i+1} - \ldots - \Pi_p$ with $i = 1, \ldots, p-1$. [13] If there are long-run equilibrium relationships, i.e. cointegrating relationships, between the variables of y_t, they must be contained in y_{t-1}. This follows from the long-run equilibrium solution of (13): which implies that all $\Delta y_{t-i} = 0$. for all lags i = 0,…,p. Taking the expectation value such that $E(\in_t) = 0$ shows, that the following equations hold in the long-run equilibrium for n = 3:

$$y_{t-1} = 0 \begin{pmatrix} \pi_{1,1} & \pi_{1,2} & \pi_{1,3} \\ \pi_{2,1} & \pi_{2,2} & \pi_{2,3} \\ \pi_{3,1} & \pi_{3,2} & \pi_{3,3} \end{pmatrix} \begin{pmatrix} p^i_{t-1} \\ p^j_{t-1} \\ e^{i,j}_{t-1} \end{pmatrix} = \begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}. \tag{14}$$

Equation (14) shows that three cases can emerge: *First*, matrix Π equals a zero matrix, i.e. the rank of the matrix is zero, $r(\Pi) = 0$. In this case no long-run equilibrium relation between the components of y_{t-1} exist, i.e. all three variables contain unit roots but are not cointegrated such that equation (13) transforms to a model in first differences. *Second*, the other extreme is that matrix Π contains three linearly independent vectors, i.e. the rank of the matrix is $r(\Pi) = 3$. In this case, three independent long-run equilibrium relationships exist, which is only possible, if all three variables are stationary in levels.[14] *Third*, the rank of the matrix is between these extremes, $0 < r(\Pi) < 3$. In this case one or two cointegrating relationships between the three variables exist. Consequently, while the AEG test can at best discover one cointegrating relationship via equation (11), the JC test can detect more than one cointegrating relationship by simply estimating the rank of Π. Since the rank of a matrix is equivalent to the number of non-zero eigenvalues (characteristic roots): every significantly non-zero eigenvalue indicates one cointegrating vector. Therefore, the Johansen trace test[15] ranks the eigenvalues according to their values, $\lambda_1 \geq \lambda_2 \geq \lambda_3$, and uses the following test statistic to test the H0 that the $r(\Pi) \leq$ r against the H1 that $r(\Pi) \geq$ r +1:

---

[13] To see the equivalence, take the case with three lags p = 3. Then the ECM reads $\Delta y_t = \Pi\, y_{t-1} + \Gamma_1\, \Delta y_{t-1} + \Gamma_2\, \Delta y_{t-2} + \in_t$ with $\Pi = \Pi_1 + \Pi_2 + \Pi_3 - \mathrm{I}$ and $\Gamma_1 = -\Pi_2 - \Pi_3$ and $\Gamma_2 = -\Pi_3$. Inserting these matrices in the VEC yields again the VAR: $y_t = \Pi_1\, y_{t-1} + \Pi_2\, y_{t-2} + \Pi_1\, y_{t-3} + \in_t$.

[14] If Π has full rank, its inverse Π^{-1} exists, such that equation (13) can be rewritten as follows: $y_{t-1} = \Pi^{-1} \Delta y_t - \Pi^{-1} \sum_{i=1}^{p-1} \Gamma_i\, \Delta y_{t-i} - \Pi^{-1} \in_t$. This implies that the levels of all three variables can be expressed as combinations of stationary variables. Consequently, the levels must be stationary too.

[15] *Johansen* (1995) provides also a maximum-eigenvalue statistic to test for cointegration. However, this method is not applied here, since this test implies a multiple-testing problem, for which so far not solution has been found.

(15) $$\lambda_{trace}(r) = -T\sum_{i=r+1}^{n} ln\left(1-\hat{\lambda}_i\right)$$

where T equals the number of observations, n equals the number of variables and r is the number of eigenvalues tested to be larger than zero. Consequently, the more the estimated eigenvalues $\hat{\lambda}_i$ deviate from zero, i.e. the larger the trace statistic. For the Bretton Woods a the ERM period, the number of variables equals n = 3 and the maximum number of cointegrating vectors equals 2, as the following table shows:

Table 6

The Johansen Trace Test for n = 3

Test	*H0*	*H1*	H0 rejected if $\lambda_{trace}(r)$ > 5% critical value
(1)	r = 0	$r \geq 1$	$\lambda_{trace}(0) = -T\left(ln\left(1-\hat{\lambda}_1\right)+ln\left(1-\hat{\lambda}_2\right)+ln\left(1-\hat{\lambda}_3\right)\right)$ > 29.68
(2)	$r \leq 1$	$r \geq 2$	$\lambda_{trace}(1) = -T\left(ln\left(1-\hat{\lambda}_2\right)+ln\left(1-\hat{\lambda}_3\right)\right)$ > 15.41
(3)	$r \leq 2$	r = 3	$\lambda_{trace}(2) = -T\left(ln\left(1-\hat{\lambda}_3\right)\right)$ > 3.76

The trace test starts with test (1). If the H0 of r = 0 is not rejected, all three variables contain unit roots but are not cointegrated and the trace test ends. If the H0 of r = 0 is rejected in favor of the H1 $r \geq 1$, test (2) is conducted. If the H0 of $r \leq 1$ is not rejected, one cointegrating relationship between the three variables exist, $[(r \geq 1)$ and $(r \leq 1)]$=> r = 1, and the trace test ends. If the H0 of $r \leq 1$ is rejected in favor of the H1 $r \geq 2$ then test (3) is conducted. If the H0 of $r \leq 2$ is not rejected, two cointegrating relationships between the three variables exist, $[(r \geq 2)$ and $(r \leq 2)]$ => r = 2, and the trace test ends. If the H0 of $r \leq 2$ is rejected in favor of the H1 $r = 3$ then the trace test ends and the hypothesis that Π has full rank is accepted, i.e. all three variables are stationary in levels. The critical values are tabulated and depend on the specification of deterministic trends (constants, dummies or trends) in equation (13). For the EMU period the number of variables equals n = 2, since the nominal exchange rates between the member states disappears and the maximum number of cointegrating vectors equals 1, as the following table shows:

Table 7

The Johansen Trace Test for n = 2

Test	*H0*	*H1*	H0 rejected if $\lambda_{trace}(r)$ > 5% critical value
(1)	r = 0	r ≥ 1	$\lambda_{trace}(1) = -T\left(ln\left(1-\hat{\lambda}_2\right) + ln\left(1-\hat{\lambda}_3\right)\right)$ > 15.41
(2)	r ≤ 1	r = 2	$\lambda_{trace}(2) = -T\left(ln\left(1-\hat{\lambda}_3\right)\right)$ > 3.76

If at least one cointegrating vector exists such that $0 < r(\Pi) < n$, the matrix Π can be decomposed as the product of two matrices α and β with the dimensions (n × r) and (n × r): respectively, where β contains the r cointegrating vectors and α contains the "amount" of each cointegrating vector that enters each equation of the VEC, also called the "adjustment parameter". Thus a corresponding estimation of the VEC offers the possibility, to check whether the signs of the cointegrating relationships are economically sensible and what causal structure underlies the adjustment process as the following representation of $\Pi\ y_{t-1}$ reveals for the case of $r(\Pi) = 2$:

$$
\begin{aligned}
&\Pi\ y_{t-1} \quad = \\
&\alpha\beta'\left(p^i_{t-1},\quad p^j_{t-1},\quad e^{i,j}_{t-1}\right)' = \\
(16)\quad &-\alpha_{1,1}\left(\beta_{1,1}\ p^i_{t-1} - \beta_{1,2}\ p^j_{t-1} - \beta_{1,3}\ e^{i,j}_{t-1}\right) - \alpha_{1,2}\left(\beta_{2,1}\ p^i_{t-1} - \beta_{2,2}\ p^j_{t-1} - \beta_{2,3}\ e^{i,j}_{t-1}\right) \\
&+\alpha_{2,1}\left(\beta_{1,1}\ p^i_{t-1} - \beta_{1,2}\ p^j_{t-1} - \beta_{1,3}\ e^{i,j}_{t-1}\right) + \alpha_{2,2}\left(\beta_{2,1}\ p^i_{t-1} - \beta_{2,2}\ p^j_{t-1} - \beta_{2,3}\ e^{i,j}_{t-1}\right) \\
&+\alpha_{3,1}\left(\beta_{1,1}\ p^i_{t-1} - \beta_{1,2}\ p^j_{t-1} - \beta_{1,3}\ e^{i,j}_{t-1}\right) + \alpha_{3,2}\left(\beta_{2,1}\ p^i_{t-1} - \beta_{2,2}\ p^j_{t-1} - \beta_{2,3}\ e^{i,j}_{t-1}\right).
\end{aligned}
$$

As indicated in equation (16): economic theory predicts for $\Delta y_t = \left(\Delta p^i_t,\ \ \Delta p^j_t,\ \ \Delta e^{i,j}_t\right)'$, see equation (13) that $\beta_{k,1}\ with\ k = 1,2$ and $\alpha_{1,k}\ with\ k = 1,2$ should have the opposite sign as $\beta_{k,l}\ with\ k = 1,2\ and\ l = 2,3$ and $\alpha_{k,l}\ with\ k = 2,3\ and\ l = 1,2$. Because a positive deviation of the cointegrating vectors $\left(\beta_{k,1}\ p^i_{t-1} - \beta_{k,2}\ p^j_{t-1} - \beta_{k,3}\ e^{i,j}_{t-1}\right)\ with\ k = 1,2$ from zero, should be followed by a negative Δp^i_t and a positive Δp^j_t and $\Delta e^{i,j}_t$ and vice versa.

To estimate equation (13) the lags p are determined by Akaike's information criterion (*Akaike* 1974). Following *Johansen* (1995) a constant[16] and orthogonalized seasonal indicators for 12 month are allowed. Appendix table 12 displays the results for the country specific JC tests as well as a couple of corresponding

[16] Since the CPI data are indexed on the base year 2010, the constant allows for differences in the absolute price levels of the base year.

estimation statistics.[17] Table 8 provides a summary for the estimated rank of Π at a significance level of 5%.

Table 8

Results of Country Specific Johansen Cointegration Tests

Period	1960:1–1972:12				1973:1–1998:12				1999:1–2017:5		
Rank of P	0	1	2	3	0	1	2	3	0	1	2
Count	32	21	2	11	10	15	10	31	54	4	8

The significance level for the rejection of the H0 is 5%.

According to table 8, the EMU period displays the highest count of no-cointegration of the real exchange rate components: in 54 out of 66 cases the H0 of a zero rank of Π cannot be rejected, i.e. the price levels are non-stationary and not cointegrated. Only in 4 cases a cointegration vector is found in Π, while in 8 cases Π has full rank, i.e. the hypothesis that the price levels are stationary cannot be rejected. The largest number of at least one cointegration vectors is found for the ERM period (25) and the Bretton Woods period (23). Taken together, these results indicate that the abolition of nominal exchange rates with the foundation of the EMU made the real exchange rates "less stationary".

From a more detailed perspective, the 4 cases where the no-cointegration of price levels can be rejected during the EMU period are: Netherlands/Austria, Germany/Belgium, Spain/Finland, Portugal/France. In all 4 cases the signs in the cointegration vector β are economically sensible;[18] for the case of Spain/Fin-

[17] Under "Cointegration rank at significance level 5%" a hyphen "–" indicates full rank of matrix Π , i.e. stationarity of all variable levels. A VEC lag selection according to Akaike's information criterion over a range of 24 month. A constant and orthogonalized seasonal indicators following *Johansen* (1995) are allowed. The 4 largest moduli of the eigenvalues of the VEC companion matrix are displayed. The modulus of a real eigenvalue is its absolute value. The modulus of a complex eigenvalue, a+b*i, is calculated according to $(a^2 + b^2)^{0.5}$. The companion matrix of a VEC with n endogenous variables and r cointegrating equations has n – r unit eigenvalues. If the process is stable, the moduli of the remaining r eigenvalues are strictly less than unity. If there are moduli larger than unity, the dynamic process is unstable and the assumptions of the JC test are not fulfilled. The Jarque-Bera test (*Jarque/Bera* 1987) is used to test for the H0 of a joint normal distribution of the VEC residuals. A Wald tests is used to test for the joint significance of Arch and Garch parameters.

[18] The identification of the elements of the β vector in equation (16) requires a normalization of one component to unity. Therefore, one coefficient of the estimations in appendix table 11 equals always exactly unity.

land the α-coefficient of Spain has however the wrong sign.[19] Thus the test results are not so different from the AEG test, where only in one case during the EMU period no-cointegration is rejected (Italy/Greece). Compared to the cointegration tests, the PP unit root tests rejected the existence of nonstationary real exchange rates more often (10 cases: Belgium/Austria, Luxembourg/Austria, Luxembourg/Belgium, Ireland/France, Ireland/Germany, Portugal/Germany, Spain/Germany, Spain/Greece, Luxembourg/Italy, Netherlands/Luxembourg).

The additional estimation statistics provided by appendix table 12 indicate that the residuals of the VEC estimation are generally serially correlated, i.e. the Jarque-Bera test of the H0 of a joint normal distribution of the VEC residuals is typically rejected, while the results of the heteroscedasticity tests for the residuals show, that most of the time the hypothesis of significant ARCH(1) and GARCH(1) processes in the residuals cannot be rejected. According to a simulation studies as *Cheung/Lai* (1993) the JC test is typically robust to excess kurtosis. Simulation studies of *Lee/Tse* (1996) and *Silvapulle/Podivinsky* (2000) show however that heteroskedastic residuals cause an "overrejection" of the H0 of no cointegration, i.e. the empirical probability of a rejection the H0 if the H0 is true, is larger than 5%. This implies the results of the country specific JC tests are probably biased towards indicating too many cases of cointegration. This explains perhaps the difference in comparison with the AEG tests. It does however not question the basic conclusion that the real exchange rate components over the EMU period are typically not cointegrated.

V. Conclusions

The empirical results of this paper indicate that for the twelve founding member states of the EMU, *firstly* consumer price indices as provided by the OECD (2010) are not stationary around linear trends, but follow a random walk with a drift and *secondly* that the resulting real exchange rates especially during the EMU period are typically random walks too. The various test statistics applied here, find some exceptions, but these exceptions are always related to different country pairs and do not allow the isolation of one or more country pairs with typically stationary real exchange rates.

The results certainly need to be complemented by similar tests applied to the Eurostat "harmonized" CPI and to subsets of the CPI, which contain tradable

[19] The four largest moduli of the companion matrix are in all four cases close to unity, whereas in the case of n variables an r cointegration vectors only n-r = 2–1 = 1 modulus should equal unity and one modulus should be strictly less than unity. So the four cointegration relationships, detected by the Johansen tests, are not confirmed by the eigenvalues of the companion matrix.

goods only. Another possibility on the search for stationary real exchange rates, is the disaggregation of price indices along the geographical dimension. In both cases the question is, at what disaggregation level do real exchange rates become stationary?

Taken the results of this study one can hardly argue that the European Currency Union is an optimal currency area. Market forces like arbitrage in goods and production factors are apparently not strong enough to keep prices at the CPI level together. As mentioned, the Balassa-Samuelson effect is not a really satisfying explanation for this result, since according to the Balassa-Samuelson effect at least "stationarity clusters" for structurally similar countries like "northern" or "southern" EMU member states should exist.

The result that there is no systematic stationarity of real exchange rates on the level of the CPIs questions the European Central Bank's principle of a "single monetary policy" (*Issing* 2001). Diverging CPIs are likely to give rise to diverging unit labor costs and thus diverging competitiveness of the member states, as mentioned in section 1. As a consequence, the competitiveness in the tradable goods sector of high-inflation (low-inflation) countries erodes (improves). The result can be deindustrialization (reindustrialization) in high-inflation (low-inflation) countries. There is a certain analogy to the mechanism in the "Dutch disease" model of *Corden/Neary* (1982). The difference is, however, the driving force working in the background: In the "Dutch disease" model this driving force is the growing demand for non-tradables financed with the external revenues from a booming resource sector. Here the driving force is a growing demand for non-tradables in high-inflation member states financed with credits at low real interest rates coming from other low-inflation member states with correspondingly high real interest rates.[20] As the experience with the Eurozone debt crisis shows, such disequilibria can built up over a time span of more than a decade until they lead to a sudden stop of capital inflows causing a crisis which finally triggers an internal devaluation (*Alcidi* et al. 2016; *Belke* et al. 2017).

If market forces are not sufficient to keep the CPI levels together on a steady base in a currency union, it might be justified to consider more differentiated country-specific monetary policies than the principle of a "single monetary policy" allows. Such policies could for example include country-specific minimum reserve requirements (*Palley* 2000; *Holz* 2007). By Article 19.1. of the ECB-Stat-

[20] Of course, under the assumption of "efficient" credit markets, such a divergence of real interest rates would not be possible in the long-run, since the shrinking creditworthiness of high inflation debtors would cause crowing risk premiums on real interest rates for high inflation debtors. However in a context, where commercial banks act as intermediaries between credit supply and credit demand, the assumption of "efficient" credit markets might not be justified, as empirical experience indicates (*Maurer* 2010).

ute, the European Central Bank has the full legal entitlement to set the minimum reserve rates. Another possibility to implement country-specific monetary policies could be the implementation of country-specific main refinancing rates. Such a regime had already been practiced by the United States' Federal Reserve System from 1914 to 1941, when discount rates were set district by district (Fraser Archive 1943).

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Appendix Tables

Online-Only-Appendix
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