Monitoring Euro Area Real Exchange Rates

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Abstract We apply the stationarity and cointegration monitoring procedure of Wagner and Wied (2014) to the real exchange rate indices, vis–*à*–vis Germany, of the first round Euro area member states. For all countries except Portugal structural breaks are detected prior to the onset of the Euro area crisis triggered in turn by the global financial crisis. The results indicate that a more detailed investigation of RER behavior in the Euro area may be useful for understanding the unfolding of the deep crisis currently plaguing many countries in the Euro area.

1 Introduction

An issue that has received a lot of attention in particular since the onset of the Euro crisis, or to be more precise, the deep economic crisis in many – often peripheral – Euro area countries, is the question whether persistent real exchange rate (RER) misalignment is one of the factors responsible for Euro area disequilibria. Clearly, given that nominal exchange rates across Euro area member states are by construction fixed at one, the nominal exchange rate is not available anymore as an instrument for readjusting RERs. Sizeable nominal exchange rate adjustments, often devaluations with respect to the "hard currencies" like the DM, have occurred almost regularly for some European countries in the decades before monetary uni-

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fication. Shutting down the currency devaluation channel may have contributed to persistent RER misalignments when institutional and price rigidities prevent smooth re-establishment of external equilibrium.

A simple, empirical cornerstone in the analysis of international equilibrium is the concept of purchasing power parity (PPP). Loosely speaking PPP states that when expressed in common currency similar baskets of goods should have similar prices across countries. To fix concepts, denote with E_t the nominal exchange rate of the country considered vis-*à*-vis Germany at time t in DM per unit of local currency, with P_t the consumer price index (CPI) of the considered country at time t and with P_t^* the CPI of Germany at time t, i.e., we consider Germany as the base country in our empirical analysis. The real exchange rate *index* Q_t with respect to the base country Germany is then defined as

$$Q_t := \frac{E_t P_t}{P_t^*}. \tag{1}$$

Taking logarithms, indicated by lower cases letters, leads to

$$q_t = e_t + p_t - p_t^*. \tag{2}$$

Given that (logarithms of) price indices and (flexible) nominal exchange rates are often modelled as integrated processes, the empirical analysis of *weak* PPP is often phrased as a unit root or cointegration testing problem, see, e.g., Wagner (2008). In this setting PPP is said to hold in its weak form, if q_t is stationary. Clearly, the value of a RER *index* that corresponds to strong purchasing power parity is undetermined when price indices rather than actual price data are used.[1]

When considering countries with different levels of development, often in addition a linear trend is included to proxy for trend RER appreciation in catching-up economies, via, e.g., the Balassa-Samuelson effect. The Balassa-Samuelson effect has been shown to be sizeably present also in Europe, see, e.g., Wagner (2005). Typically, trend stationary log RERs are interpreted as being consistent with structural ("catching-up" or convergence) processes towards PPP.

The definition (2) of the log RER is often used as a basis for a cointegrating relationship between the prices indices and the nominal exchange rate of the form

$$p_t = c + \delta t + \beta_1 e_t + \beta_2 p_t^* + u_t, \tag{3}$$

where $\{u_t\}$ is a stationary process and where – obviously – a trend stationary log RER corresponds to the restrictions $\beta_1 = -1,\ \beta_2 = 1$.[2]

[1] Strong PPP is typically defined as a RER equal to one and thus its logarithm equal to zero. Typically, however, due to the lack of actual price data empirical analysis is confined to work with RER indices where the "level information" is lost.

[2] It is natural to normalize the potential cointegrating relationship on the price index p_t – or p_t^* – rather than on the nominal exchange rate, which may have been fixed or almost fixed even before monetary unification (e.g., the exchange rate between the Austrian Schilling and the DM) and which is in that case almost by construction not an integrated process.

Given that RER misalignments are, as mentioned at the beginning, seen by many observers as a contributing factor to the crisis in some Euro area economies it is a natural question to ask whether a monitoring procedure detects structural changes in countries' RERs away from trend stationarity respectively cointegration in the looser discussed formulation (3). This is the question we analyze for the "first round" Euro area member states.[3] In the present situation there is, of course, some extra information concerning important dates available, most notably the fixing of the final nominal exchange rates between the 12 first round member states on December 31, 1998 and the introduction of the Euro as a virtual currency on January 1, 1999.[4] Consequently, we consider as the so-called calibration period for the monitoring procedure (see the description in the following section) a period until December 1998 for which trend stationarity of the log RER prevails. For all countries except Austria, where the calibration period is set to begin in January 1991 this means that the calibration period is set to begin in either February 1994 or February 1996. Subsequently, we monitor the behavior of the log RER from the beginning of the Euro area in January 1999 until July 2014 with the procedure described in the following section. The resulting *detection times*, if a break is detected, are the estimated break points at which a structural break in the considered RER indices has occurred.

2 A Brief Description of the Monitoring Procedure

The monitoring procedure used in this contribution has been developed in Wagner and Wied (2014), where detailed descriptions as well as a detailed analysis of the asymptotic and finite sample properties of the procedure and tables with critical values are contained.

We consider, under the null hypothesis, a cointegrated system in triangular form:

$$y_t = D_t'\theta + X_t'\beta + u_t \tag{4}$$

$$X_t = X_{t-1} + v_t, \tag{5}$$

with observations available for $t = 1, \ldots, T$. Here D_t is a deterministic trend function, in our application given by constant and linear trend. The joint vector process $\{[u_t, v_t']'\}$ is under the null hypothesis of a cointegrating relationship an I(0) vector

[3] These are Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain. Here it has to be noted that Luxembourg used the Belgian Franc prior to using to Euro and Greece was scheduled to join the Euro area in 2001 only. Thus, effectively we use the data for 11 countries as Luxembourg has to be excluded. As indicated in the main text we consider RERs with base country Germany. This is a natural choice given that Germany is the largest Euro area economy and the DM was the European anchor currency at the time. Alternatively, it is possible to calculate for each country its RER index with respect to the Euro area.

[4] The first quotation of the Euro was on January 4, 1999 and the physical introduction occurred at the beginning of 2002.

process, i.e.,

$$\frac{1}{\sqrt{T}}\sum_{t=1}^{\lfloor sT \rfloor}\begin{pmatrix} u_t \\ v_t \end{pmatrix} \Rightarrow_d \Omega^{1/2}W(s), \quad 0 \leq s \leq 1, \tag{6}$$

where $W(s)$ is a vector of standard Wiener processes, Ω is the so-called long-run variance matrix of $\{[u_t, v_t']'\}$ and $\lfloor z \rfloor$ denotes the integer part of a real number z. Throughout we assume that the long-run variance matrix of $\{v_t\}$ is positive definite, to exclude cointegration amongst the components of $\{X_t\}$. This implies that the relationship (4) is the only cointegrating relationship for the vector process $\{[y_t, X_t']'\}$. Under the alternative, the cointegrating relationship between $\{y_t\}$ and $\{X_t\}$ breaks down from some break fraction $\lfloor rT \rfloor$ onwards, for some $0 < m \leq r < 1$. Thus, the cointegrating relationship (4) turns into a spurious relationship at time point $\lfloor rT \rfloor$. For reasons discussed below, an initial period up to $\lfloor mT \rfloor$, in which the cointegrating relationship can safely be assumed to be present, is required. Clearly, in the absence of integrated regressors, with consequently $\dim(X_t) = 0$, the framework simplifies to monitoring trend stationarity of $\{y_t\}$.

To explain the idea of the monitoring procedure, assume for the moment that $\{u_t\}$ is observed and its long-run variance ω^2 known. Then the detector is given by

$$H^m(s) := \frac{1}{\omega^2}\left(\frac{1}{T}\sum_{i=\lfloor mT \rfloor+1}^{\lfloor sT \rfloor}\left(\frac{1}{\sqrt{T}}S_i\right)^2 - \frac{1}{T}\sum_{i=1}^{\lfloor mT \rfloor}\left(\frac{1}{\sqrt{T}}S_i\right)^2\right) \tag{7}$$

for $m \leq s \leq 1$ and with $S_i = \sum_{t=1}^{i} u_t$ denoting the partial sums of u_t. The detector is given by combining the well-known KPSS-statistic of Kwiatkowski et al. (1992) for the null hypothesis of stationarity with the monitoring approach of Chu et al. (1996). Under the null hypothesis, it holds that

$$H^m(s) \Rightarrow_d \mathscr{H}^m(s) := \left(\int_m^s W(z)^2 dz - \int_0^m W(z)^2 dz\right). \tag{8}$$

Since under the alternative, the partial sum process as scaled in (7) diverges for $s > r$, the null hypothesis is rejected if an appropriately scaled version of the detector exceeds a critical value, i.e., when

$$\left|\frac{H^m(s)}{w(s)}\right| > c(m, w, \alpha), \tag{9}$$

for some weighting function $w(s)$ and appropriate critical values $c(m, w, \alpha)$ that are chosen such that under the null hypothesis the detection time τ_m, i.e., the first time the critical value is exceeded, is asymptotically finite with probability equal to α. We use the same weighting function as Wagner and Wied (2014), i.e., $w(s) = s^5$ in our specification with intercept and linear trend.

In practice, rather than the errors $\{u_t\}$ one only observes residuals, $\hat{u}_{t,m}$ say (indicating the dependence upon the calibration fraction m), and also the long-run variances are unknown and have to be estimated. It is well-known in the cointegration literature that due to the endogeneity of the regressors, as $\{u_t\}$ and $\{v_t\}$ are allowed to be dynamically correlated, the OLS parameter estimators are consistent with their limiting distribution dependent upon nuisance parameters relating to regressor endogeneity and error serial correlation. This, of course, implies that the limiting distribution of the properly scaled partial sum process of the OLS residuals also depends upon nuisance parameters. For this reason, residuals based on an estimator that corrects for endogeneity and that takes into account serial correlation have to be used. The literature offers several possibilities in this respect, with the most prominent being Fully Modified OLS (FM-OLS) of Phillips and Hansen (1990), Dynamic OLS (D-OLS) of Saikkonen (1991), and Integrated Modified OLS (IM-OLS) of Vogelsang and Wagner (2014). In our empirical analysis we use all three estimators.

Obtaining a nuisance parameter free limiting distribution of the properly scaled partial sum residual process rests upon appropriate estimation under the null hypothesis. This can be done in several ways, e.g., by using a moving window or by using a calibration period $1,\dots,\lfloor mT \rfloor$ at the beginning of the sample. Wagner and Wied (2014) opt, following Chu et al. (1996), for the second route.

The null limiting distribution of $H^m(s)$, when calculated using residuals $\hat{u}_{t,m}$ as input, depends upon the deterministic components included, the number of integrated regressors, the calibration sample fraction m and the estimator chosen. Thus, critical values, for chosen weighting function, can be obtained by simulation.

Looking at equation (4) it is clear that the monitoring procedure based on the partial sum residual process using parameter estimates based on the calibration period is not only consistent against the spurious regression alternative, but also against breaks in the parameters θ or β that occur at some time point $\lfloor rT \rfloor \geq \lfloor mT \rfloor$. In case of such a structural break the scaled partial sum residual process contains for observations later than $\lfloor rT \rfloor$ a divergent component, as $\hat{\theta}_m$ and $\hat{\beta}_m$, indicating in the notation here the dependence of the estimates on the calibration sample fraction, converge by construction to the pre-break values. Thus, detection of a structural change need not necessarily indicate a spurious relationship but can also indicate a structural change towards a cointegrating relationship with different slope and/or trend parameters.

3 Empirical Analysis

In Figure 1 we display the log RER indices over the period January 1991 to July 2014. The red boxes included in the figure for each of the countries display the calibration period required for the monitoring procedure. The calibration period is set to end in December 1998, i.e., it ends just before the Euro introduction at the beginning of 1999, for all countries. For Austria the calibration period starts in January 1991, for Belgium, Finland, the Netherlands and Portugal in February 1994 and for France, Greece, Ireland, Italy and Spain in February 1996. This choice of the calibration period has been made to avoid the period of high exchange rate volatility and instability in the aftermath of the 1992 UK currency crisis ("Black Wednesday"). Clearly, such a trimming is necessary as we need a period of stationarity, respectively cointegration, for calibration. The trimming also means that the calibration period is very short, which implies that parameter estimation can be expected to be imprecise.

Graphical inspection of the series shows that for most countries the RER appreciates with respect to Germany over the largest part of the period.[5] Also note that Finland's RER is the most stable one with respect to Germany after the period of Finland's severe crisis following the collapse of the Soviet Union.

Table 1 Detected break points

Country	*Calibration Period*	*Trend Stationarity*	*Cointegration* FM-OLS	*Cointegration* D-OLS	*Cointegration* IM-OLS
Austria	91(1)–98(12)	2005-10			
Belgium	94(2)–98(12)	2001-12	2003-04	2003-12	2009-01
Finland	94(2)–98(12)		2003-04	2003-04	2013-07
France	96(2)–98(12)		2001-08	2001-09	2002-12
Greece	96(2)–98(12)		2004-08	2003-03	2014-06
Ireland	96(2)–98(12)		2001-10	2001-11	2002-10
Italy	96(2)–98(12)	2007-11	2004-12	2004-06	
Netherlands	94(2)–98(12)	2000-10	2002-03	2001-06	2002-10
Portugal	94(2)–98(12)				
Spain	96(2)–98(12)	2000-08	2001-05	2001-07	2002-01

In Table 1 we display the detected break points when trend stationarity of q_t as given in (2) or a cointegrating relationship of the form given in (3) is monitored. The main findings are: First, for all countries except Portugal a break point is detected. Second, all detected breaks occur well before the Euro area crisis, triggered in turn by the global financial crisis, has spread. Third, to a certain extent surprising, for four countries (Finland, France, Greece and Ireland) the more restrictive hypothesis of trend stationarity of the real exchange rate is not rejected, but the looser cointegration null hypothesis is. Fourth, by and large and as expected, the cointegration break dates are later than the stationarity break dates.

The detected break points are also indicated by vertical lines in Figure 1. For some countries, e.g., Austria and France, the detected break points correspond to the first clearly "visible" changes in the behavior of the log RER indices. For other countries no such clear mapping between visual inspection and statistical analysis is present. A detailed investigation of the detected break points as well as an interpretation of the findings is beyond the scope of this paper that is merely meant to illustrate the cointegration monitoring procedure.

[5] Only at the end of the period one observes RER depreciation relative to Germany in the peripheral crisis countries like Greece, Ireland, Portugal and Spain in line with the deep recession and structural transformation process ongoing in these countries.

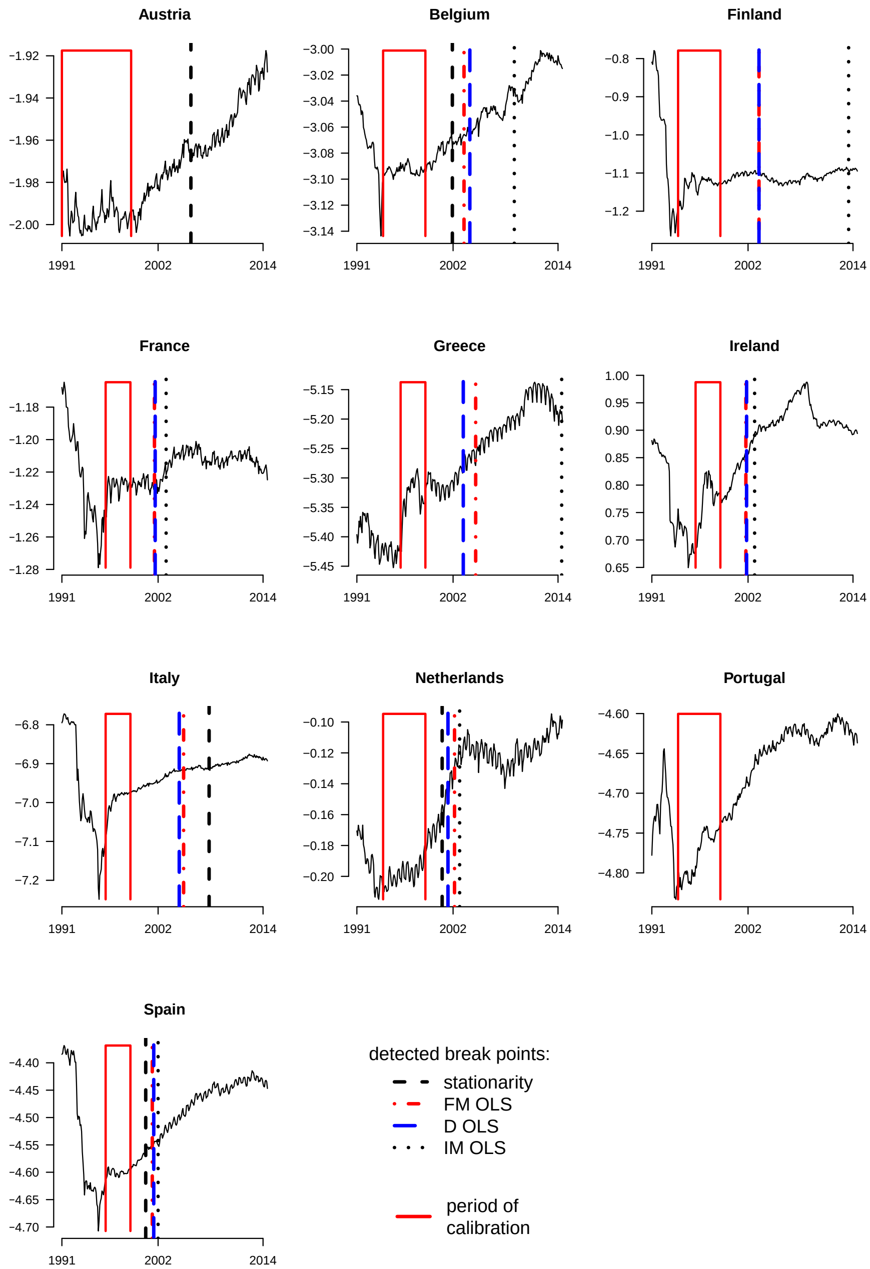

Fig. 1 Log RER indices, periods of calibration, and detected break points (intercept and linear trend)

4 Summary and Conclusions

We have applied the stationarity and cointegration monitoring procedure of Wagner and Wied (2014) to the RER indices, vis–*à*–vis Germany, of the first round Euro area member states. Clearly, our results are merely meant as an illustration and can only serve as one of many inputs into a thorough economic analysis of Euro area RER behavior. Nevertheless, the findings do indicate that methods for investigating the structural stability of stationary respectively cointegrating series may provide useful input for economic analysis. This in turn implies that several questions need to be addressed from an econometric theory perspective to provide potentially more useful tools: First, it may be relevant to flip null and alternative hypothesis, i.e., to monitor changes from I(1) or spurious to I(0) or cointegrating behavior to monitor entry into a period of "equilibrium". Second, multivariate monitoring procedures may be important for applied research in order to exploit the fact that often multiple series are affected at or at least around the same time. Exploiting in such cases also the cross-sectional dimension may lead to more powerful monitoring procedures. Third, especially important for monitoring data collected at higher frequencies, the effects of non-constant variances need to be investigated in detail. Robust, or correspondingly modified, procedures need to be developed for such situations.

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