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On European monetary integration and the persistence of
real effective exchange rates

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Abstract

This paper deals with the possibility of changing persistence in European real effective exchange rates as initially analyzed by Gadea and Gracia (2009). By applying a CUSUM of squares-based test for constant versus changing persistence with desirable statistical properties, an OECD data set is reconsidered. The empirical results suggest that persistence remains constant over time for nearly all time series. Thus, European monetary integration has not affected the persistence of external competitiveness significantly. Moreover, strong evidence for non-stationarity is found. Explanations for the sharp contrast of new results towards the ones by Gadea and Gracia (2009) are provided.

JEL classification: C22, E61, F31, F42

Key Words: Changing persistence, unit roots, structural breaks, European monetary integration.

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1 Introduction

European monetary integration has gained much attention during the last decades. The European Monetary System (EMS) has been the centrepiece of integration process prior to the launch of Economic and Monetary Union (EMU). The evolution of real effective exchange rates is of particular interest as it is an overall measure of the country's external competitiveness. It is defined as a weighted average of a basket of foreign currencies adjusted by foreign and domestic price levels.

For most European countries, and especially for peripheral, the European monetary integration process has been accompanied by an increasing monetary stability. The European economics may have benefited from the requirements of the nominal convergence process. Thus, as a result of the Euro introduction or previously increased macroeconomic stability, persistence of real effective exchange rates may have been reduced. In order to study this possibility, Gadea and Gracia (2009) apply a test proposed by Busetti and Taylor (2004) for stationarity against changing persistence. The empirical usefulness of this test is questionable as it may generate spurious rejections. These may occur if the true data generating process is a random walk without any change in persistence. Leybourne et al. (2007) demonstrate by means of a Monte Carlo study that the test by Busetti and Talyor (2004) exhibits spurious rejection probabilities up to 90%. As real effective exchange rates are potentially strongly autocorrelated, see Sarantis (1999) for early evidence, this problem may be of great relevance here. Moreover, Leybourne et al. (2007) propose a CUSUM of squares-based test which is robust with respect to stationary and non-stationary models.

The main contribution of this paper is to reconsider the OECD data set originally analyzed in Gadea and Gracia (2009) by applying the CUSUM of squares-based test which has desirable statistical properties. This test remedies the pitfall of rejecting the null hypothesis in favor of a change in persistence when data is highly persistent. The empirical results suggest that persistence remains constant over time for nearly all countries in the sample. In addition, the evidence for non-stationarity in real effective exchange rates is particularly strong. This gives rise to the conclusion that the results in Gadea and Gracia (2009) may be driven by non-stationary data instead of time series with changing persistence. This implies that the process of European monetary integration and the increased monetary stability has not affected the persistence properties of real effective exchange rates in Europe significantly.

This paper is organized as follows: Section 2 reviews econometric procedures for changing persistence. Empirical results are presented and discussed in section 3, while section 4 concludes.

2 Tests for changing persistence

Consider the following data generating process (DGP) which consists of a deterministic term d_t and an autoregressive error process v_t :

$$y_t = d_t + v_t \tag{1}$$

$$v_t = \rho_t v_{t-1} + \varepsilon_t \tag{2}$$

with $t = 1, 2, \dots, T$. The deterministic term is either a constant μ or a constant plus a linear trend $\mu + \beta t$. The innovations (ε_t) are assumed to be linear and stationary and to have short memory. A change in persistence means a change in the degree of inte-

gration. If, for example, $\rho_t = 1$ for all t , then v_t is a unit root process over the entire sample period. On the contrary, if $\rho_t = \rho$ with $|\rho| < 1$, then v_t is a stationary AR(1) process for $t = 1, 2, \dots, T$. In these two cases, persistence of y_t is constant over time and not subject to a structural change. In the case where $\rho_t = 1$ for $t = 1, 2, \dots, T_B$ and $|\rho_t| < 1$ for $t = T_B + 1, T_B + 2, \dots, T$, persistence is said to be decreasing. The case of increasing persistence is defined analogously. The unknown breakpoint is given by $T_B = \lceil \tau T \rceil$ where $\lceil x \rceil$ denotes the biggest integer smaller than x and $\tau \in (0, 1)$.

Tests against changing persistence has been proposed (among others) by Kim (2000), Kim et al. (2002), Busetti and Taylor (2004) and Leybourne et al. (2007). In general, the direction of change needs not be specified a priori for all these tests. In this application, however, a decline in persistence is of main interest: Gadea and Gracia (2009) discuss a decline in persistence of real effective exchange rates due to greater stability of prices and nominal exchange rates. Nonetheless, the possibility of increasing persistence is not excluded in this work.

Gadea and Gracia (2009) apply the stationarity test of Busetti and Taylor (2004) against a decline in persistence. It can be formulated as a test of H_0 against H_{10} :

$$\begin{aligned}
 H_0 & : \quad |\rho_t| < 1 \quad \text{for all } t \\
 H_{10} & : \quad \begin{cases} \rho_t = 1 & \text{for } t = 1, \dots, T_B \\ |\rho_t| < 1 & \text{for } t = T_B + 1, \dots, T. \end{cases}
 \end{aligned}$$

In the context of stationarity, unit roots and changing persistence, the fourth possibility of a constant $I(1)$ process plays an important role, i.e.,

$$H_1 : \rho_t = 1 \quad \text{for all } t.$$

The tests proposed by Kim (2000), Kim et al. (2002) and Busetti and Taylor (2004) for H_0 against H_{10} (or H_{01}) share the major drawback that they reject the null hypothesis spuriously if actually H_1 is true. In this situation, the null hypothesis is wrong, but neither are the alternatives true. In other words, if the time series under consideration does not display any change in persistence, but is $I(1)$ throughout the entire sample, then the mentioned tests reject their null hypothesis of stationarity (H_0) spuriously.

It is of great importance to note that the test statistics are not diverging under the validity of H_1 as they are of order $O_p(1)$. Nonetheless, as clearly demonstrated in Leybourne et al. (2007), these tests are dramatically over-sized with simulated rejection probabilities up to 90%. Therefore, these tests lack of empirical usefulness in the sense that they are not able to discriminate between time series with constant and changing persistence.

The recently proposed test by Leybourne et al. (2007) overcomes this problem by suggesting a CUSUM of squares-based test for H_1 against H_{10} (or H_{01}). As shown in Leybourne et al. (2007), the CUSUM of squares-based test statistic behaves conservatively under the validity of H_0 . It converges to a constant (one) in probability. This means that the asymptotic size of this statistic equals zero and that no spurious rejections may occur asymptotically. Even in samples of small and moderate size, this test does not reject the unit root hypothesis spuriously. The test statistic (R) is given by

$$R = \frac{\inf_{\tau \in \Lambda} K^f(\tau)}{\inf_{\tau \in \Lambda} K^r(\tau)},$$

where $K^f(\tau)$ and $K^r(\tau)$ are CUSUM of squares-based statistics. They are based on

the forward and reversed residuals of the data generating process as given below. The relative breakpoint $\tau \in \Lambda = [\underline{\tau}, \bar{\tau}]$ is assumed to be unknown so an estimator for τ is given below. In more detail, $K^f(\tau)$ and $K^r(\tau)$ are given by

$$K^f(\tau) = \frac{1}{[\tau T]^2 \hat{\gamma}_0^f(\tau)} \sum_{t=1}^{[\tau T]} \hat{v}_{t,\tau}^2$$

and

$$K^r(\tau) = \frac{1}{(T - [\tau T])^2 \hat{\gamma}_0^r(\tau)} \sum_{t=1}^{T - [\tau T]} \tilde{v}_{t,\tau}^2.$$

Here, $\hat{v}_{t,\tau}$ are the residuals from the OLS regression of y_t on a constant based on the observations up to $[\tau T]$. This is

$$\hat{v}_{t,\tau} = y_t - \bar{y}(\tau)$$

with $\bar{y}(\tau) = [\tau T]^{-1} \sum_{t=1}^{[\tau T]} y_t$. Similarly $\tilde{v}_{t,\tau}$ is defined for the reversed series $z_t \equiv y_{T-t+1}$. In addition, $\hat{\gamma}_0^f(\tau)$ and $\hat{\gamma}_0^r(\tau)$ are OLS variance estimators for $\Delta \hat{v}_{t,\tau}$ and $\Delta \tilde{v}_{t,\tau}$, respectively. Analogous expressions for the case of de-trending can be found in Leybourne et al. (2007). The null hypothesis of a constant unit root process is rejected for large values of R in favor of the alternative. Regarding the unknown breakpoint, Leybourne et al. (2007) prove the consistency of a breakpoint estimator under H_{10} which is given by

$$\hat{\tau}^r = \arg \inf_{\tau \in \Lambda} \frac{1}{(T - [\tau T])^2} \sum_{t=1}^{T - [\tau T]} \tilde{v}_{t,\tau}^2.$$

Note, that $\frac{1}{(T - [\tau T])^2} \sum_{t=1}^{T - [\tau T]} \tilde{v}_{t,\tau}^2$ is equal to the unstandardized backward statistic $K^r(\tau)$ (without the long-run variance estimator). A similar consistent breakpoint estimator can be constructed under H_{01} , see Leybourne et al. (2007).

The simulation evidence in Leybourne et al. (2007) suggest that this test is correctly sized and has satisfying power properties. In addition, it allows valid inference under

constant persistence, irrespectively if y_t is constantly $I(1)$ or constantly $I(0)$.

3 Empirical results

The used data set is exactly the same as in Gadea and Gracia (2009). The CUSUM of squares-based test is applied to log real effective exchange rates $y_t = s_t - p_t + p_t^*$. For further details and the construction of the data set, the interested reader is referred to Appendix A in Gadea and Gracia (2009). The sample spans from 1975Q1 to 2003Q2 so that the sample size equals 114. The interval of potential breakpoints is set equal to [1977Q3,2000Q3] which corresponds to the common choice of $\underline{\tau} = 0.1$ and $\bar{\tau} = 0.9$ in the structural break literature. Alternatively, a sub-sample is considered as well. It ranges from 1975Q1 to 1996Q4 ($T = 88$) and excludes the euro period (late nineties and early 2000s) from the sample. In this case, the interval of potential breakpoints is set equal to [1976Q4,1994Q4]. In the following, results for the full sample are reported, as the results for the sub-sample analysis are very similar and do not change any main conclusion. Both types of deterministic terms are considered: (i) constant and (ii) constant plus linear trend.

Results for the CUSUM of squares-based test are reported in Table 1. The most striking feature is that only one rejection occurs at the nominal 10% level of significance if $d_t = \mu$ (de-meaning) and only three if $d_t = \mu + \beta t$ (de-trending). For Belgium, Finland and Portugal evidence in favor of a change in persistence is found. Moreover, rejections occur only for time series where real effective exchange rates are constructed on the basis of unit labour costs (lc). These clear results are in a sharp contrast to the results found in Gadea and Gracia (2009).

Table 1: Empirical test results for changing persistence

Country	lc	xp	cp	lc	xp	cp
d_t	De-meaning			De-trending		
Austria	0.549	1.383	0.614	0.668	0.395	0.685
Belgium	1.475	0.240	2.037	2.579	1.108	1.040
Denmark	0.592	1.215	1.962	1.375	1.776	1.272
Finland	1.715	1.609	1.502	1.945	1.141	1.453
France	0.825	0.523	0.728	0.623	0.776	0.831
Germany	1.308	1.395	0.720	0.998	0.882	0.724
Greece	2.466	0.606	0.656	0.737	0.476	0.747
Ireland	0.539	1.071	0.592	0.862	1.057	0.733
Italy	0.569	0.861	1.368	0.700	0.819	0.687
Netherlands	1.638	1.220	1.826	0.795	1.363	1.137
Portugal	4.343	0.355	0.741	5.401	0.776	0.817
Spain	0.717	0.512	0.853	1.112	0.865	0.826
Sweden	0.700	0.836	0.850	1.255	1.027	1.088
United Kingdom	2.484	2.380	1.524	1.529	1.707	1.612

Notes: Reported values are CUSUM of squares-based test statistics. Bold faced values indicate a rejection at the nominal 10% level of significance. lc , xp and cp denote real effective exchange rates based on unit labour costs, export prices and consumer prices, respectively.

Due to the statistical properties of the CUSUM of squares-based test statistic, it is concluded that changes in persistence are rarely the case, but it would be premature to conclude that the considered time series are $I(1)$. The reason is that the test is conservative if y_t is $I(0)$ and therefore, non-rejections may be caused by both type of processes, namely $I(1)$ and $I(0)$ without a change in persistence. In order to further investigate the time series properties of the data set, the Dickey-Fuller GLS test proposed by Elliot et al. (1996) is applied to all time series where the CUSUM of squares-based test does not reject the null hypothesis in favor of a change in persistence. Test results for 38 remaining time series are reported in Table 2. As can be seen, five rejections of the unit root hypothesis are found if a constant is specified as the deterministic term and four rejections if a linear trend is additionally included. The evidence for non-stationarity is particularly strong.¹

If these results are taken together with the simulation evidence in Leybourne et al. (2007), one may explain why Gadea and Gracia (2009) find evidence for changes in persistence in this data set: The non-stationarity of time series can lead to severe over-rejections of the test proposed by Busetti and Taylor (2004). Therefore, it is concluded that the test by Busetti and Taylor (2004) produces spurious evidence for changes in persistence in this empirical application as predicted by the simulation results in Leybourne et al. (2007). It is recommended to use tests like the one suggested by Leybourne et al. (2007) for empirical work. It allows valid inference under both, constant $I(0)$ and constant $I(1)$ processes.

¹This general finding appears to be same when applying other unit root tests.

Table 2: Empirical test results for unit roots

Country	<i>lc</i>	<i>xp</i>	<i>cp</i>	<i>lc</i>	<i>xp</i>	<i>cp</i>
<i>d_t</i>	Constant			Constant plus trend		
Austria	-1.467	0.380	-0.859	-1.741	-1.257	-1.529
Belgium	-0.479	-0.231	-1.123	—	-0.903	-1.833
Denmark	0.332	-1.554	-1.530	-1.520	-1.847	-2.285
Finland	-0.422	-2.085	-1.283	—	-3.455	-1.791
France	0.115	0.060	-1.298	-2.831	-2.662	-2.563
Germany	-0.361	-0.781	-0.852	-1.517	-2.482	-1.738
Greece	-0.499	0.452	-1.119	-1.654	-1.840	-1.237
Ireland	0.649	-0.749	-1.713	-3.147	-3.628	-1.867
Italy	-1.963	-0.629	-2.042	-1.931	-2.596	-2.075
Netherlands	-0.501	-0.702	-1.771	-1.153	-1.946	-1.967
Portugal	—	0.779	-0.924	—	-2.389	-1.098
Spain	0.196	-1.408	-1.338	-2.348	-1.635	-2.038
Sweden	-0.480	-0.592	-0.753	-2.447	-2.413	-2.255
United Kingdom	-0.658	-1.110	-1.425	-2.090	-2.417	-2.008

Notes: Reported values are DF-GLS test statistics. Bold faced values indicate a rejection at the nominal 10% level of significance. *lc*, *xp* and *cp* denote real effective exchange rates based on unit labour costs, export prices and consumer prices, respectively. No test statistic is reported if a time series exhibits a change in persistence, see Table 1.

4 Conclusions

This paper re-considers the persistence properties of European real effective exchange rates, with a focus on potentially changing persistence. Such a structural change from non-stationarity to stationarity may be the result of increased macroeconomic stability in terms of more stable nominal exchange rates and inflation rates during the period of European monetary integration. The application of a CUSUM of squares-based test for changing persistence leads to the conclusion that persistence has not changed during the considered period from 1975 to 2003 with some minor exceptions. Weak evidence for a decline in persistence is found only for Belgium, Finland and Portugal. These results are conflicting with the ones reported in Gadea and Gracia (2009), where the authors apply a test with less desirable statistical properties. The discrepancy between the results can be explained by the non-stationarity of most time series in this data set. Indeed, an application of a unit root test suggests that more than two-third of the 42 analyzed time series are $I(1)$. These conclusions still hold if a sub-sample excluding the Euro period is considered.

From an economic viewpoint, one may conclude that the increased macroeconomic and monetary stability in the peripheral countries has not affected the persistence properties of its real effective exchange rates. Moreover, most of them show a constantly high degree of persistence, meaning that shocks do not die out. The empirical results also suggest that neither the introduction of the Euro nor any other important event in the history of European monetary integration has influenced the persistence of external competitiveness.

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