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On the relation between exchange rates and tourism demand: A nonlinear and asymmetric analysis

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ABSTRACT

Previous studies of demand for tourism assumed that the effect of exchange rate fluctuations on the demand for tourism were symmetric for appreciations and depreciations, in both sign and size. In this paper we separate appreciations from depreciations and find that the assumption is not justified. Thus, this paper examines the extent to which exchange rate changes have an asymmetric effect on tourism demand in ten European countries by using hidden cointegration analysis within a likelihood-based panel framework. The findings indicate that there is a long-run relationship between tourism demand and exchange rate fluctuations which give support for the view that tourism demand have responded asymmetrically to the exchange rate changes. The economic implication of our results is that depreciations and appreciations affect tourism demand differently both in sign and size. It is worth mentioning that the methodological framework utilized in this paper can be a useful tool for designing and predicting the behavior of exchange rates.

1. Introduction

This paper examines the impact of exchange rate changes on tourism demand in ten European countries (Denmark, Norway, Sweden, Switzerland, Czech Rep., Russian federation, Croatia, Hungary, Poland, and Romania). Tourism and travel help to create trade and generate prosperity in addition to building bridges between people from different places and cultures. Sustainable tourism development creates income, jobs, and tax revenues for the community and significantly contributes to the development of better and more attractive infrastructure. Tourism has greatly contributed to countries in terms of GDP and employment. For example in Sweden, tourism's proportion of GDP is higher than for agriculture, forestry, commercial fishery and the food industry combined (Swedish Agency for Economic and Regional Growth, 2015). Table A1, Appendix A, shows the importance of tourism in the countries under review.

Much of the development taking place in the area of tourism depends on changes in consumer behavior which is largely determined by exchange rate fluctuations. The link between foreign exchange rate changes and tourism has captured the interest of both policy makers and travelers as it plays an important role in the development of an economy. The relationship between tourism and exchange rates stems from the fact that a depreciation (an appreciation) of the home currency makes travelling for foreign visitors cheaper (expensive).

Most of the previous empirical studies use the real exchange rate in modelling tourism demand (e.g., Lim, 1999; Dritsakis, 2004; Li, Song, & Witt, 2005; Vogt, 2008; Seo, Park, & Yu, 2009; Seetaram, 2010; Thompson & Thompson, 2010; Chang & McAleer, 2012; Cheng, Kim, & Thompson, 2013; Culiuc, 2014; Falk, 2015; Tang, Sriboonchitta, Ramos, & Wong, 2016; Yazdi & Khanalizadeh, 2017; Ongan, Isik, & Özdemir, 2017; Stettler, 2017). They apply different methods such as gravity frameworks, dynamic heterogeneous panel data co-integration technique, multivariate generalized autoregressive conditional heteroskedasticity (MGARCH), vector autoregression

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error correction model, the Copula-Based GARCH Model, and the generalized Poisson regression model for different countries to identify the effect of exchange rate on tourism demand. Although most studies find that exchange rate or its volatility is an important determinant of tourism inflows or industry trade in general, the degree of significant varies from study to study, i.e., from lack of cointegration or weak cointegration to strong cointegration relation.[1]

No matter which method is used to estimate the demand for tourism, a common feature and the implicit assumption is that exchange rate fluctuations have symmetric effects. The symmetry assumption implies that if a home currency's depreciation causes foreign residents who holiday abroad increase, appreciation should decrease it by the same amount. This may not be the case due to changes in expectations and speculations. In reality, consumers may react to unanticipated and anticipated changes in the real exchange rate differently (Baum, Caglayan, & Barkoulas, 2001; Ho & Iyke, 2017). A negative shock to exchange rates could create higher uncertainty because of the increased expectations of a speculative attack linked to it (Byrne & Davis, 2005). Since the effect of exchange rate changes is based on expectations and speculations, it is expected that its effects on the demand for tourism to be asymmetric.

In this paper, we open a new path in the literature by arguing that the effects of exchange rate changes on tourism demand could be asymmetric. Thus, this paper examines the impact of exchange rate changes on tourism demand by applying a combination of multivariate hidden cointegration (Granger & Yoon, 2002) and likelihood-based panel cointegration (Larsson, Lyhagen, & Lothgren, 2001) to ten European countries over the period 1995-2016.

The hidden cointegration technique allows both for asymmetry in the long-run relationship between data components and for distinct cointegrating relationships between subcomponents of time series even when cointegration between two time series is not identified. Another advantage of using this technique is that it allows a straightforward delimitation of the data in an economically sensible way. The estimation process is an asymptotic theory of likelihood-based panel test of the cointegrating rank for inference in panel-vector autoregressive (VAR) models which allows for multiple cointegrating vectors.

The paper is organized as follows. Section 2 outlines the model and methods, Section 3 reports the results supporting asymmetry effects, while Section 4 presents summary and conclusion.

2. Data and methodology

The sample consists of ten European countries (Denmark, Norway, Sweden, Switzerland, Poland, Czech Republic, Croatia, Hungary, Romania, and Russian Federation) and covers the period 1995-2016. The data used in this study are real effective exchange rates (defined as domestic currency per unit of a weighted average of several foreign currencies) and number of non-resident arrivals which transformed to natural logarithms. The series were obtained from the World Bank. The reason for choosing the time period was dictated by the availability and the reason for choosing the sample countries is justified by the fact that none of the sample countries has Euro as a currency and most of the currencies have been subject to severe fluctuations.[2] The sample includes both Western and Eastern European countries. The descriptive statistics and time plot of the variables are shown in Table A2 and Figs. 1–2, Appendix A, respectively.

As it is shown in the introduction, all previous studies use linear relationships in which the adjustment of tourism demand to exchange rate shocks is assumed to be independent of whether shocks are positive or negative. However, if tourism demand responds larger in magnitude to appreciation than depreciation or vice versa, then this relationship is asymmetric.

There is no any common approach as to the existence, or nature, of the asymmetric relationship between tourism demand and exchange rate changes. Error correction models (ECMs) or VAR models used in the literature suffer from problems of low power in test statistics and bias issues stemming from mis-specified exogenous thresholds for determining statistical regimes. Furthermore, ECM or VAR models assume asymmetry as a short-run relationship between the series, that is, asymmetry presents only in the adjustments process to the equilibrium and not in the cointegration relationship and there exists no long-run asymmetry.

To overcome the problems related to threshold autoregressive (TAR) ECMs, Granger and Yoon (2002) approach is adopted here that deals effectively with these issues.[3] They investigate the presence of a cointegrating relationship not between the aggregate series but between their components, which they call 'hidden cointegration'. That is, they allow for the possibility that, even if no linear cointegration exists, there may be a long-run relation between the positive and negative non-stationary components of some series.

This technique allows us to study not only if tourism demand respond to exchange rate shocks but also if this response depends on the sign of the shocks. Granger and Yoon also show that the non-linear adjustment mechanism to long-run equilibrium can be easily reduced to a linear one without any loss of information. Their hidden cointegration technique identifies the dynamics between data components. The data components include both cumulative positive and negative changes of time series. If the components of two data series (negative or positive) are cointegrated, then the data have a hidden cointegration. This is an example of non-linear cointegration that ordinary linear cointegration fails to identify. Suppose X_t and Y_t are two random walk time series:

[1] The effect of exchange rate on industry trade has also been analyzed by Bahmani-Oskooee, Harvey, and Hegerty (2018). Using non-linear ARDL, they conclude that the effect of exchange rate on industry trade is asymmetric.

[2] Real effective exchange rate was missing in Croatia during 1995-1997 and number of arrivals was missing in Switzerland in 2004. These were estimated by regression imputations.

[3] Another approach to investigate the possibility of asymmetric adjustment is threshold cointegration tests (e.g., Balke & Fomby, 1997; Enders & Siklos, 2001). In the threshold autoregressive (TAR) model, asymmetry is a combination of short-run and long-run processes. The hidden cointegration is more flexible than threshold cointegration or the standard ECMs since it is not limited to two regimes and it is possible to explore all different combinations of cointegration between data components.

$$X_t = X_{t-1} + \varepsilon_t = X_0 + \sum_{i=1}^{t} \varepsilon_i, \tag{1}$$

$$Y_t = Y_{t-1} + \eta_t = Y_0 + \sum_{i=1}^{t} \eta_i, \tag{2}$$

where $t = 1,2, \ldots T$ and X_0, Y_0 are initial values, ε_i and η_i denote mean zero white noise disturbance terms. A standard cointegration exists if $\{X_t, Y_t\}$ are cointegrated by one cointegrating vector. When movements of X_t and Y_t are asymmetric, it is possible to detect hidden cointegrations between them. Granger and Yoon (2002) define positive and negative shocks as follows:

$$\varepsilon_i^+ = \max(\varepsilon_i, 0), \varepsilon_i^- = \min(\varepsilon_i, 0), \eta_i^+ = \max(\eta_i, 0), \eta_i^- = \min(\eta_i, 0),$$

$$\varepsilon_i = \varepsilon_i^+ + \varepsilon_i^- and \eta_i = \eta_i^+ + \eta_i^-.$$

Thus:

$$X_t = X_{t-1} + \varepsilon_t = X_0 + \sum_i^t \varepsilon_i^+ + \sum_i^t \varepsilon_i^- \quad \text{and} \quad Y_t = Y_{t-1} + \eta_t = Y_0 + \sum_i^t \eta_i^+ + \sum_i^t \eta_{i.}^-. \tag{4}$$

Then the notation can be simplified with:

$$X_t^+ = \sum_i^t \varepsilon_i^+, X_t^- = \sum_i^t \varepsilon_i^-, Y_t^+ = \sum_i^t \eta_i^+, and Y_t^- = \sum_i^t \eta_i^-. \tag{5}$$

Thus:

$$X_t = X_0 + X_t^+ + X_t^- and Y_t = Y_0 + Y_t^+ + Y_t^-. \tag{6}$$

Consequently:

$$\Delta X_t^+ = \varepsilon_t^+, \Delta X_t^- = \varepsilon_t^-, \Delta Y_t^+ = \eta_t^+, \Delta Y_t^- = \eta_t^-. \tag{7}$$

The first difference, $(\Delta X_t = X_t - X_{t-1})$ is calculated for both of the time series, which sort observations to positive and negative movements ($\Delta X_t^+ and \Delta X_t^-$). Then, it is also calculated the cumulative sum of positive (negative) changes, at a given time, for all variables $X_t^+ = \sum \Delta X_t^+ and X_t^- = \sum \Delta X_t^-$ via equations that are presented above. Similar calculations are implemented for Y. The variables X and Y are said to have hidden cointegration if their components are cointegrated.

The process is estimated by implementing likelihood-based panel framework developed by Larsson and Lyhagen (1999) and Larsson et al. (2001).[4] By using this method, the assumption of a unique cointegrating vector and the problem of normalization is relaxed which is not the case with the usual residual-based tests of cointegration approach.[5] Let LR denote the cross-section-specific likelihood-ratio (trace) statistic of the hypothesis that there are at most r cointegrating vectors in the system. The standardized LR-bar statistic is given by:

$$Y_{\bar{LR}} = \frac{\sqrt{N(\bar{LR} - \mu\,)}}{\sqrt{V}}, \tag{8}$$

where $\bar{LR}$ is the average of the N cross-section LR statistics, μ is the mean, and ν is the variance of the asymptotic trace statistic. Asymptotic values of μ and ν (with and without constant and trend) can be obtained from stochastic simulations as described in Johansen (1995).[6] In the presence of both stochastic and deterministic trends, Hatemi-J and El-Khatib (2016) suggest an alternative approach for transforming variables into partial cumulative positive and negative components.

Two steps should be followed before applying any cointegration tests: testing the panel for cross-sectional dependence and testing for cross-country heterogeneity. The first issue implies the transmission of shocks from one variable to others. In other words, all countries in the sample are influenced by globalization and have common economic characteristics. The second issue indicates that a significant economic connection in one country is not necessarily replicated by the others. A set of three tests is constructed in order to check the cross-sectional dependence assumption: the Breusch and Pagan (1980) cross-sectional dependence (CD_{BP}) test, the Pesaran (2004) cross-sectional dependence (CD_P) test, and the Pesaran, Ullah, and Yamagata (2008) bias-adjusted LM test (LM_{adj}). Regarding the country-specific heterogeneity assumption, the slope homogeneity tests ($\bar{\Delta}$ and $\bar{\Delta}_{adj}$) of Pesaran and Yamagata (2008) are used (Appendix B provides more information about these tests).

[4] It is based on Johansen's (1995) maximum likelihood approach.
[5] The residual-based panel cointegration was developed by Kao (1999) and Pedroni (1999).
[6] For other applications of hidden cointegration and likelihood-based panel cointegration see Irandoust and Ericsson (2005) and Irandoust (2017).

The panel unit root tests used previously do not take into account cross-sectional dependence of the contemporaneous error terms. Failing to consider cross-sectional dependence may lead to misleading results. Thus, to remove this problem, we use the cross-sectionally augmented panel unit root test (CIPS) which allows for parameter heterogeneity and serial correlation between the cross-sections (Pesaran, 2007).[7]

Finally, we check if the underlying assumptions are satisfied, i.e., if the residuals are normal distributed and there is no autocorrelation. The normality test is a multivariate extension of the Bowman–Shenton test developed by Doornik and Hansen (1994) and the test for autocorrelation is the Ljung–Box test statistics.

3. Estimation results

Positive and negative values for the number of arrivals and the real effective exchange rate variables, as described by Equations (1)–(7), are generated for each country. The hidden cointegration technique naturally allows a straightforward delimitation of the data in an economically sensible way: the positive component of real effective exchange rate carries the concept of appreciations (i.e., a domestic currency appreciates against foreign currencies), while the negative component carries the concept of depreciations. In the same way, the negative component of number of arrivals carries the concept of a decrease in the number of arrivals and the positive component of number of arrivals carries the concept of an increase in the number arrivals.

As a pre-test for the cointegration analysis, we first investigate cross-sectional dependence and slope homogeneity assumptions. Table 1 reports the results of cross-sectional dependence tests (CD_{BP}, CD_p, and LM_{adj}) and slope homogeneity tests ($\bar{\Delta}$ and $\bar{\Delta}_{adj}$). The first set of tests, for cross-sectional dependence, clearly reveals that the null hypothesis of no cross-sectional dependence is rejected for all significance levels. More precisely, this implies that there is a cross-sectional dependence in the case of our sample countries for both specifications. Any shock in one country is transmitted to others. The second part of the Table shows that the null hypothesis of slope homogeneity is rejected for both tests and for all significance levels. In this case, the economic relationship in one country is not replicated by the others. As there are both cross-sectional dependence and slope heterogeneity for both specifications, the cointegration tests can be applied.

We test for panel non-stationarity among the variables before applying cointegration test. The results of the cross-sectionally augmented IPS test are listed in Table 2. After inspection of the data, we only include a constant term (it could also be due to measurement errors). When applying the Schwartz criterion to decide the optimal lag length, the common lag length was set to three. The table shows that all variables seem to support the null hypothesis of panel non-stationarity. Furthermore, note that our approach does not exclude the possibility of including stationary variables. The effect of one stationary variable in the system is that the rank order increases with one.

Turning to the cointegration analysis, we consider both positive and negative components. The likelihood ratio tests for positive and negative components are given in Table 3. The Bartlett corrected critical values are gained by using the estimated model as data generating process when calculating the sample mean. Using the Bartlett corrected critical values, the test rejects the null of 0 cointegrating rank but accepts the null of 1 cointegrating vector for both components. Note that if we use the asymptotic critical values, the estimated rank is 2.

Hence, the panel cointegration tests reveal that the common cointegrating rank is one and it is therefore interesting to estimate the cointegrated vectors between the positive and negative components. The estimated cointegrating vectors, normalized with respect to *lnARR*+ and *lnARR*-, are presented in Table 4. The fact that *lnREX*+ carries negative coefficient implies that, as domestic currency appreciates against foreign currencies the number of arrivals decreases in the sample countries and positive coefficients of *lnREX*-implies that, as domestic currency depreciates against foreign currencies the number of arrivals increases in all countries under review. Furthermore, when testing for a common cointegrating vector by using the likelihood ratio test, we obtain (−1.000, 0.638) with a *p* value of 5 percent for appreciation (*lnARR*-, *lnREX*+) and (−1.000, 0.773) with a *p* value of 5 percent for depreciation (*ln ARR*+*, lnREX*-). An alternative approach suggested by Hatemi-J and El-Khatib (2016) to transform data and include a deterministic trend was also tested but it did not alter the general outcome. In the case of Sweden, Norway, Denmark, Poland, Switzerland, and Russian the results remain almost the same but the procedure reduces the absolute values of appreciation and depreciation coefficients in Czech Repub. (*lnREX*-, 4. 257*, lnREX*+, −3.942), Hungarian (*lnREX*-, 3.439, *lnREX*+, −2.806), Croatia (*lnREX*-, 4.328, *lnREX*+, −4.824), and Romania (*lnREX*-, 2.037, *lnREX*+ −1.565). This may stem from the fact that real exchange rates are nonstationary stochastic process which do not revert to a deterministic path and they just contain stochastic trends.

However, the magnitude of parameters varies from country to country and they appear to be smaller in the Eastern European countries compared to the Nordic countries and Switzerland. The higher coefficient elasticity may be due to the larger share of business travel in the Nordic countries and Switzerland compared to the Eastern European countries. Table 4 also shows that the sign and size of the effect are different for appreciations compared to depreciations, supporting the asymmetric hypothesis.

In Table 5, the results from the diagnostic tests are reported. It seems that we do not have any problem with autocorrelation since all p-values are very high but the null hypothesis of normality is rejected in the case of appreciations. The problem of normality could not be solved by using more lags.

[7] The CIPS panel unit root test is based on the Im, Pesaran, and Shin (2003) test (IPS), which controls for cross-sectional heterogeneity in the estimated coefficients. The CIPS is the average of the individual country cross-sectionally augmented ADF (CADF) statistics.

Table 1
Cross-sectional dependence and slope homogeneity tests.

Method	Test statistic ($lnARR+$, $lnREX-$)	Test statistic ($lnARR-$, $lnREX+$)
Cross-sectional dependence test		
CD_{BP}	243.632*** (0.000)	275.391*** (0.000)
CD_P	27.261*** (0.000)	32.372*** (0.000)
LM_{adj}	35.527*** (0.000)	43.244*** (0.000)
Slope homogeneity test		
$\bar{\Delta}$ test	9.546*** (0.000)	12.805*** (0.000)
$\bar{\Delta}_{adj}$ test	8.168*** (0.000)	10.726*** (0.000)

Notes.
1. *** indicate significance for 0.01 levels. The numbers within parentheses show *p*-values.
2. CD_{BP} test, CD_P test, and LM_{adj} test show the cross-sectional dependence tests of Breusch and Pagan (1980), Pesaran (2004), and Pesaran et al. (2008), respectively.
3. $\bar{\Delta}$ and Δ^{-}_{adj} tests show the slope homogeneity tests proposed by Pesaran and Yamagata (2008).

Table 2
Panel unit root test.

Variable	CIPS statistic
lnREX+	−1.936
lnREX-	−1.875
lnARR+	−2.013
lnARR-	−2.125

Note.
a.. Critical values for the CIPS test are −2.60 (1%), −2.34 (5%), and −2.21 (10%), Pesaran (2007).

Table 3
Test for cointegrating rank.

H_o	ACV [a]	BCV [b]	$-2logQ_T$
(*lnREX+*, *lnARR-*)			
$R = 0$	204.73	238.95	247.96
$R \leq 1$	86.55	138.64	126.33
(*lnREX-*, *lnARR+*)			
$R = 0$	236.51	275.74	289.22
$R \leq 1$	91.79	164.11	150.25

Notes.
[a] The asymptotic critical values at 5% significance level.
[b] Bartlett corrected critical values at 5% significance level.

Table 4
Cointegrating vectors normalized on *lnARR+* and *lnARR-*.

	Sweden	Denmark	Norway	Switzerland	Poland	Russian Fed.	Czech Rep.	Croatia	Hungary	Romania
lnARR-	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000
lnREX+	−0.657	−0.603	−0.645	−0.710	−0.313	−0.559	−0.537	−0.563	−0.472	−0.336
lnARR+	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000
lnREX-	0.745	0.648	0.620	0.846	0.421	0.601	0.519	0.528	0.413	0.353

4. Summary and conclusion

Several authors have considered the demand for tourism and have applied different methods and have addressed different issues. No matter which method is used to estimate the demand for tourism, a common feature and the implicit assumption is that exchange rate fluctuations have symmetric effects. In this paper, we open a new path in the literature by arguing that the effects of exchange rate changes on tourism demand could be asymmetric. The departure from earlier studies of the role of exchange rate on tourism demand is in the asymptotic theory of panel cointegration and nonlinear or asymmetric analysis. The countries under review are Sweden, Norway,

Table 5
Diagnostic tests.[a].

Model	Normality[b]	Autocorrelation[c]
lnARR+, lnREX-	0.0604	0.3538
LnARR-, lnREX+	0.2632	0.4719

Notes.
[a] The table reports the p-values.
[b] The test is a multivariate extension of the Bowman–Shenton test developed by Doornik and Hansen (1994).
[c] This is the Ljung–Box test statistics for autocorrelation.

Denmark, Czech Rep., Russian Fed., Croatia, Switzerland, Hungary, Romania, and Poland for the period 1995-2016.

The hidden cointegration approach within the framework of likelihood-based panel cointegration analysis is applied to reveal the long-run relationship between exchange rate changes and tourism demand. The findings show that appreciations and depreciations are negatively and positively associated with number of arrivals, respectively, in all countries in the sample. The magnitude of parameters varies from country to country and they appear to be smaller in the Eastern European countries compared to the Nordic countries and Switzerland.

Furthermore, the findings indicate that the effect of exchange rate changes on tourism demand is asymmetric. The economic implication of our results is that depreciations and appreciations affect tourism demand differently both in sign and size. The policy implication of the findings is that a country depending on its tourism industry, should impose some restrictions on the use of exchange rate policies to correct its international competitiveness, as these policies may lead to an exchange rate volatility that could, in turn, reduce its tourism inflows.

Acknowledgement

The comments made by an anonymous referee are greatly appreciated. The usual disclaimer applies.

Conflicts of interest

The material in the manuscript does not infringe upon any copyright and there is no any conflict of interest. Thank you for your time.

Appendix A

Table A1
The importance of tourism in the sample countries (2017)

Country/Total Contribution to	GDP (USD bn)	Employment (000 jobs)
Sweden	52.4 (9.5)	558 (11.1)
Denmark	25.1 (7.7)	214 (7.8)
Norway	36.3 (9)	337 (12.7)
Switzerland	61.9 (9.1)	602 (12)
Poland	23.9 (4.5)	738 (4.5)
Russian Fed.	76.1 (4.8)	3.256 (4.5)
Czech Rep.	16.8 (7.8)	477 (9.4)
Croatia	13,743.3 (25)	320 (23.5)
Hungary	10,991.8 (8)	324 (7.3)
Romania	11,185.7 (5.3)	529 6.3

Notes.
a. GDP is calculated in 2017 constant prices and exchange rates.
b. The numbers in parentheses present percentage of total GDP and total employment, respectively.
c. Source: World Travel and Tourism Council (2018).

Table A2
Descriptive statistics of the real effective exchange rates and tourism inflows in the sample countries

Country	Mean	S.D.	Skewness	Kurtosis
Sweden				
REXSWE	105.5000	7.326302	0.383569	3.125472
ARRSWE	4426909	1291669	−0.297912	2.349541
Denmark				
REXDEN	96.73302	3.233382	−0.267572	2.722571
ARRDEN	6513591	3465427	−0.189725	1.233190

(*continued on next page*)

Table A2 (*continued*)

Country	Mean	S.D.	Skewness	Kurtosis
Norway				
REXNOR	94.13655	4.681658	−0.247294	1.714066
ARRNOR	3962636	926255.8	0.349199	2.100821
Switzerland				
REXCHE	97.63493	7.783165	0.633154	2.122412
ARRCHE	7882636	911301.3	0.105047	1.552799
Russian Fed.				
REXRUS	80.67378	18.03402	−0.094010	1.969357
ARRRUS	22688227	5384397	0.101067	3.397836
Poland				
REXPOL	91.84398	9.575021	−0.440806	2.821647
ARRPOL	15755409	2301433	0.176896	1.980509
Czech Rep.				
REXCZE	81.81331	15.52022	−0.217235	1.667086
ARRCZE	6221318	1452898	0.331576	2.699303
Croatia				
REXCRO	93.71206	4.655973	0.199299	2.084211
ARRCRO	7801818	3158979	−0.117601	2.493297
Hungary				
REXHUN	86.33110	13.01841	−0.463630	1.750035
ARRHUN	3515136	722342.3	1.156244	3.221887
Romania				
REXROU	88.49365	15.96226	−0.649950	2.413463
ARRROU	6733682	1642634	0.465858	2.018612

Figs. 1–2: Time plots of real effective exchange rates and number of arrivals in the sample countries (1995-2016).

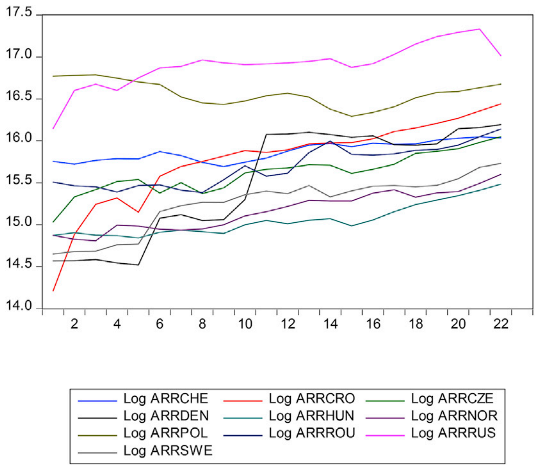

Fig. 1. Number of arrivals in the sample countries (1995-2016)

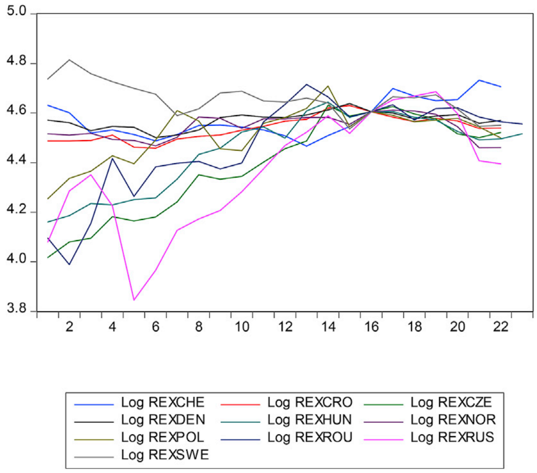

Fig. 2. Real effective exchange rates in the sample countries (1995-2016)

Appendix B

Cross-sectional dependence tests

Breusch and Pagan's (1980) LM test has been used in many empirical studies to test cross-sectional dependency. LM statistics can be calculated using the following panel model:

$$y_{it} = \alpha_i + \beta_{it}^{\circ} x_{it} + \mu_{it}, \qquad i = 1, 2 \quad , N \quad t = 1, 2, ..., T, \tag{1A}$$

where i is the cross-section dimension, t is the time dimension, x_{it} is $k \times 1$ vector of explanatory variables while α_i and β_i are the individual intercepts and slope coefficients, respectively, that are allowed to differ across states. In the LM test, the null hypothesis of no cross-sectional dependence H_0: $Cov(\mu_{it}, \mu_{jt}) = 0$ for all t and $i \neq j$ is tested against the alternative hypothesis of cross-sectional dependence H_1: $Cov(\mu_{it}, \mu_{jt}) \neq 0$ for at least one pair of $i \neq j$. For testing the null hypothesis, Breusch and Pagan (1980) developed the following test:

$$CD_{BP} = T \sum_{i-1}^{N-1} \sum_{j=i+1}^{N} \rho_{ij}^{\wedge 2}, \tag{2A}$$

where$\rho_{ij}^{\wedge 2}$ is the estimated correlation coefficient among the residuals obtained from individual OLS estimation of Eq. (1A). Under the null hypothesis, the LM statistic has an asymptotic chi-square distribution with $N(N\text{-}1)/2$ degrees of freedom. Pesaran (2004) proposes that the LM test is only valid when N is relatively small and T is sufficiently large. To overcoming this problem, Pesaran (2004) introduces the following LM statistic for the cross-section dependency test:

$$CD_p = \sqrt{\frac{1}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \left(T\rho_{ij}^{\wedge 2} - 1 \right). \tag{3A}$$

However, Pesaran et al. (2008) state that while the population average pair-wise correlations are zero, the CD test will have less power. Therefore, they proposed a bias-adjusted test that is a modified version of the LM test by using the exact mean and variance of the LM statistic. The bias-adjusted LM statistic is calculated as follows:

$$LM_{adj} = \sqrt{\frac{2T}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \rho_{ij}^{\wedge 2} \frac{(T-k)\rho_{ij}^{\wedge 2} - u_{Tij}}{\sqrt{v_{Tij}^2}}, \tag{4A}$$

where $u_{Tij}$and $v_{Tij}^2$are the exact mean and variance of $(T-k)\rho_{ij}^{\wedge 2}$, which are provided in Pesaran et al. (2008). Under the null hypothesis of no cross-sectional dependence with $T \to \infty$ first followed by $N \to \infty$, the results of this test follow an asymptotic standard normal distribution.

Slope homogeneity tests

In order to relax the assumption of homoscedasticity in the F-test, Swamy (1970) developed the slope homogeneity test that examines the dispersion of individual slope estimates from a suitable pooled estimator. Pesaran and Yamagata (2008) state that both the F-test and Swamy's test require panel data models where N is relatively small compared to T. To overcome this problem, they proposed a standardized version of Swamy's test (the so-called $\tilde{\Delta}$ test) for testing slope homogeneity in large panels. The $\tilde{\Delta}$ test is valid when *(N, T)* → ∞ without any restrictions on the relative expansion rates of N and T when the error terms are normally distributed. Pesaran and Yamagata (2008) then develop the following standardized dispersion statistic:

$$\bar{\Delta} = \sqrt{N}\left(\frac{N^{-1}S^{\approx} - k}{\sqrt{2k}}\right), \tag{5A}$$

where $S^{\approx}$ is Swamy's statistic. Under the null hypothesis with the condition of *(N, T)* → ∞ and when the error terms are normally distributed, the $\tilde{\Delta}$ test has an asymptotic standard normal distribution. The small sample properties of the $\tilde{\Delta}$ test can be improved when there are normally distributed errors by using the following mean and variance bias adjusted version:

$$\bar{\Delta}_{adj} = \sqrt{N}\left(\frac{N^{-1}S^{\approx} - E\left(z_{it}^{\approx}\right)}{\sqrt{\mathrm{var}\left(z_{it}^{\approx}\right)}}\right), \tag{6A}$$

where $E(z_{it}^{\approx}) = k, \mathrm{var}(z_{it}^{\approx}) = 2k(T - k - 1))/(T + 1)$.

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