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# PPP hypothesis and temporary structural breaks

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## Abstract

In this study our aim is to explore a better testing strategy for the PPP hypothesis under a temporary structural break. For this purpose we use the exponential smooth transition (EST) function in the unit root testing framework and compare this methodology with the one that uses a Fourier function. Although the Fourier function is extensively used in the literature to test the validity of the PPP hypothesis under temporary breaks, this investigation shows that it leads to misleading results.

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## 1. Introduction

One of the puzzling questions that still challenge many scholars in the open economy macroeconomics literature is how to reconcile the theory of PPP with the weak empirical evidence supporting its validity[1]. One potential answer is that a testing methodology powerful enough to reject the unit root behavior in the real exchange rates (RERs) has not yet been developed. To circumvent this problem one strand of the literature inspired by Perron (1989) have developed univariate or panel unit root tests that take structural breaks in the RERs into account. Along these lines Hegwood and Papell (1998) have emphasized that, if these deviations are permanent then this supports quasi-PPP, whereas the standard PPP still holds under the temporary break condition.

Following these studies, Christopoulos and Leon-Ledesma (2010) (CL) have attempted to test the standard version of the PPP hypothesis allowing for temporary structural breaks in the RERs using the Flexible Fourier Form (FFF)[2]. In their study CL claim that the FFF can be used to model temporary smooth breaks that are consistent with the standard long-run PPP hypothesis. However, FFF type of intercept evolves continuously over time with only the start and end points being the same[3]. Thus, with the FFF modeling structure it is not possible to find a constant mean that is required for the temporary break case of the standard PPP hypothesis to hold. This study shows that the appropriate way to test for this hypothesis is to employ an exponential smooth transition (EST) function, which allows for two constant and identical means at both ends of the temporary structural break. While the outer regimes of the EST function have symmetric structures that enable the constant mean values necessary for the long-run PPP to hold, the inner regime provides the temporary structural break type of behavior. Thus, testing for the standard PPP hypothesis under temporary structural breaks will best be accomplished using an EST type of function rather than a Fourier function.

EST type of detrending is first used in Omay and Emirmahmutoğlu (2017) (OE). These authors have utilized the EST function to model the structural break in the series when testing for a unit root. However, OE have only compared the optimization algorithms within the context of the smooth transition (ST) type of detrending and overlooked the economic intuition behind the usage of such an EST unit root test.

Section 2 of this paper explains the proposed test statistics and presents their critical values. It also provides the small sample performance of the proposed test in comparison with the power of the alternative tests. Section 3 applies the aforementioned test to the PPP hypothesis. Section 4 concludes.

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[1] See Carvalho and Julio (2012) for an extensive survey of the recent literature.

[2] CL has tested for the PPP hypothesis using both a Fourier ADF test and a test that includes both a Fourier and an ESTAR function. See CL for further details.

[3] CL considers a break to be temporary if the mean to which the RERs convert are the same at the start and end of the sample. For the RERs to display a temporary break and be compatible with the long-run PPP, the mean value to which they revert before and after the break should be the same during a considerably long period of time and not at just two points. Moreover, the start and the end values of the FFF intercept is the same, but this is not due to the fact that the means are the same at both ends of the sample. This equality emerges naturally because of the behavior of the FFF function. Furthermore, rejecting unit root behaviour with this type of an FFF intercept, in fact, provides ambiguous results with respect to the validity of the PPP hypothesis. Rejection of the null neither means that the standard PPP holds nor that the quasi-PPP version of it is valid. It rather provides a RER that is mean reverting around an occasionally changing equilibrium RER.

## 2. The model and testing framework

Let $y_t$ be a changing trend function with smooth transition:

Model A: $y_t = \alpha_1 + \alpha_2 F_t(\gamma,\tau) + \varepsilon_t$ (2.1)

Model B: $y_t = \alpha_1 + \beta_1 t + \alpha_2 F_t(\gamma,\tau) + \varepsilon_t$ (2.2)

Model C: $y_t = \alpha_1 + \beta_1 t + \alpha_2 F_t(\gamma,\tau) + \beta_2 F_t(\gamma,\tau) t + \varepsilon_t$ (2.3)

where $\varepsilon_t$ is a zero mean $I(0)$ process and $F_t(\gamma,\tau)$ is the following EST function, based on a sample of size *T*:

$$F_t(\gamma,\tau) = 1 - \exp\left[-\gamma(t-\tau)^2\right], \quad \gamma > 0 \tag{2.4}$$

In this modeling strategy the structural change is modeled as a smooth transition between different regimes rather than an instantaneous structural break[4]. In these specifications no change and one instantaneous structural change are limiting cases[5].

We establish the hypotheses for unit root testing based on models A, B and C as follows:

$$\begin{aligned} &H_0 : \textit{Unit Root}\,(\text{Linear Nonstationary}) \\ &H_1 : \textit{Nonlinear Stationary}\,(\text{Stationary around smoothly changing trend and intercept}) \end{aligned} \tag{2.5}$$

Following Leybourne *et al.* (1998) the test statistics proposed here are calculated with a two-step procedure:

**Step 1.** Using constrained nonlinear optimization algorithm via SQP[6], we estimate only the deterministic component of the preferred model and compute its residuals.

Model A: $\hat{\varepsilon}_t = y_t - \hat{\alpha}_1 - \hat{\alpha}_2 F_t(\hat{\gamma},\hat{\tau})$ (2.6)

Model B: $\hat{\varepsilon}_t = y_t - \hat{\alpha}_1 + \hat{\beta}_1 t - \hat{\alpha}_2 F_t(\hat{\gamma},\hat{\tau})$ (2.7)

Model C: $\hat{\varepsilon}_t = y_t - \hat{\alpha}_1 - \hat{\beta}_1 t - \hat{\alpha}_2 F_t(\hat{\gamma},\hat{\tau}) - \hat{\beta}_2 F_t(\hat{\gamma},\hat{\tau}) t$ (2.8)

**Step 2.** Compute the ADF statistic, which is the *t* ratio associated with $\hat{\rho}$ in the ordinary least squares (OLS) regression

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[4] For details see Omay and Emirmahmutoglu (2017) and for the logistic function you can find further details in Omay and Yıldırım (2014) and Omay et. al (2014).

[5] For further discussion and possible extensions see Leybourne *et al.* (1998).

[6] We have used the Sequential Quadratic Programming (SQP) algorithm in our estimation process since it is shown to be the best performing algorithm in estimating EST type of equations. For details see Omay and Emirmahmutoglu (2017).

$$\Delta\hat{\varepsilon}_t = \rho\hat{\varepsilon}_{t-1} + \sum_{j=1}^{p}\delta_j\Delta\hat{\varepsilon}_{t-j} + u_t \tag{2.9}$$

For models A, B and C we denote the t statistics for $\hat{\rho}$ as $\tilde{s}_{\alpha}$, $\tilde{s}_{\alpha(\beta)}$, and $\tilde{s}_{\alpha\beta}$, respectively.

$H_0 : \rho = 0,$ (Linear Nonstationary)
$H_1 : \rho < 0,$ (Stationary around nonlinear trend and/or intercept )

By applying the above described estimation process, we have obtained the critical values for the EST test statistics $\tilde{s}_{\alpha}$, $\tilde{s}_{\alpha(\beta)}$ and $\tilde{s}_{\alpha\beta}$ as follows:

**Table 1** Critical values for EST models

| | $\tilde{s}_{\alpha}$ | | | $\tilde{s}_{\alpha(\beta)}$ | | | $\tilde{s}_{\alpha\beta}$ | | |
|---|---|---|---|---|---|---|---|---|---|
| *T* | 1% | 5% | 10% | 1% | 5% | 10% | 1% | 5% | 10% |
| 25 | -5.989 | -4.954 | -4.497 | -6.783 | -5.732 | -5.231 | -7.299 | -6.188 | -5.656 |
| 50 | -5.230 | -4.537 | -4.185 | -5.910 | -5.147 | -4.782 | -6.228 | -5.468 | -5.081 |
| 100 | -5.017 | -4.374 | -4.051 | -5.544 | -4.900 | -4.572 | -5.797 | -5.166 | -4.844 |
| 200 | -4.810 | -4.271 | -3.976 | -5.328 | -4.766 | -4.494 | -5.592 | -5.036 | -4.738 |

Note: The critical values are obtained by 20000 replication.

### 2.1 Finite sample performance

We have investigated the empirical size of the EST unit root test by using the following data generating process[7]:

$$y_t = y_{t-1} + \varepsilon_t, \quad y_0 = 0 \quad \varepsilon_t \sim iidN(0,1)$$

**Table 2** Empirical sizes of $\tilde{s}_{\alpha}$, $\tilde{s}_{\alpha(\beta)}$ and $\tilde{s}_{\alpha\beta}$ for pure random walk DGP

| *T* | $\tilde{s}_{\alpha}$ | $\tilde{s}_{\alpha(\beta)}$ | $\tilde{s}_{\alpha\beta}$ |
|---|---|---|---|
| 25 | 5.45 | 5.75 | 5.95 |
| 50 | 5.25 | 5.15 | 5.15 |
| 100 | 4.60 | 5.45 | 5.30 |
| 200 | 5.45 | 5.60 | 5.30 |

**Note:** Sample size is taken as $T = 100$.

Table 2 indicates that the EST test has good size properties. We have also investigated the empirical power of the EST test by using the following data generating process where the process is a stationary linear adjustment around a smooth transition. Thus, the following EST-AR(1) was employed as the DGP:

$$y_t = 1.0 + \alpha_2 F_t(\gamma, \tau) + \varepsilon_t, \quad \mu_0 = 0$$
$$\varepsilon_t = 0.8\varepsilon_{t-1} + u_t \quad \varepsilon_0 = 0 \quad u_t \sim iidN(0,1)$$

[7] We have conducted the size analysis for also the following DGP: $\Delta y_t = \varphi\Delta y_{t-1} + \varepsilon_t$, $y_0 = 0$ $\varepsilon_t \sim iidN(0,1)$ ARIMA (1,1,0). The results are available upon request.

where $F_t(\cdot)$ is defined as before, and all combinations of the following parameter values were used: $\gamma = -0.01, -0.05, -0.10, -0.30, -0.50$, $\tau = 0.2, 0.5$, and $\alpha_2 = 5.0, 10.0$. The results from these power experiments for a sample size of $T = 100$ are given in Table 4[8].

**Table 3** The power analysis of EST test with Model A

| $\alpha_2$ | $\gamma$ | $\tau$ | $\tilde{s}_\alpha$ | $s_\alpha$ | $\tau_{DF,c}$ | $\tau_{FDF,c}$ |
|---|---|---|---|---|---|---|
| 5 | 0.01 | 0.2 | **30.95** | 12.25 | 22.80 | 22.00 |
| 5 | 0.05 | 0.2 | 22.50 | 8.95 | **27.00** | 20.85 |
| 5 | 0.10 | 0.2 | 22.50 | 13.00 | **37.75** | 30.00 |
| 5 | 0.30 | 0.2 | 25.10 | 26.55 | **58.85** | 51.80 |
| 5 | 0.50 | 0.2 | 29.55 | 35.10 | **68.40** | 62.30 |
| 5 | 0.01 | 0.5 | **32.15** | 2.80 | 6.40 | 20.60 |
| 5 | 0.05 | 0.5 | **23.85** | 6.25 | 22.85 | 21.60 |
| 5 | 0.10 | 0.5 | 22.55 | 9.85 | **32.90** | 30.15 |
| 5 | 0.30 | 0.5 | 25.75 | 24.45 | **57.60** | 52.80 |
| 5 | 0.50 | 0.5 | 29.85 | 32.55 | **67.50** | 62.60 |
| 10 | 0.01 | 0.2 | **30.60** | 1.35 | 0.35 | 0.30 |
| 10 | 0.05 | 0.2 | **19.65** | 0.05 | 0.30 | 0.20 |
| 10 | 0.10 | 0.2 | **17.95** | 0.15 | 2.40 | 0.45 |
| 10 | 0.30 | 0.2 | 15.35 | 6.80 | **39.60** | 22.85 |
| 10 | 0.50 | 0.2 | 15.50 | 27.55 | **71.75** | 54.90 |
| 10 | 0.01 | 0.5 | **30.45** | 0.00 | 0.00 | 0.20 |
| 10 | 0.05 | 0.5 | **20.30** | 0.00 | 0.20 | 0.20 |
| 10 | 0.10 | 0.5 | **17.30** | 0.15 | 1.40 | 0.65 |
| 10 | 0.30 | 0.5 | 15.25 | 4.20 | **34.25** | 23.15 |
| 10 | 0.50 | 0.5 | 15.15 | 23.15 | **67.30** | 56.55 |

**Notes :** $\tilde{s}_\alpha$, $s_\alpha$, $\tau_{DF,t}$ and $\tau_{FDF,t}$ denote the EST, LNV, ADF and Fourier ADF tests, respectively. The sample size is $T = 100$.

In most of the regions the power of the ADF test exceeds that of the newly proposed test as it is shown in Omay and Emirmahmutoğlu (2017). However, OE have compared the power of their test with that of ADF using the intercept and trend model because as stated in Leybourne *et al.* (1998) Model A's natural competitor is the ADF test that includes both an intercept and trend term. On the other hand, testing the PPP hypothesis necessitates the inclusion of only an intercept term in the test regressions. Thus, the power analysis conducted by including only an intercept term shows that in both the small and large break cases with low transition speed the newly proposed test has better power performance than that of the other tests[9]. Moreover, increasing the size of break further improves the power of the newly proposed test. Fortunately, as it is described in Perron (1989), RERs deviate persistently from their equilibrium value due to long lived events that also causes large breaks.

## 3. Empirical Example

In this section we empirically apply the EST, ADF and Fourier ADF tests to examine the validity of the standard purchasing power parity (PPP) hypothesis for 24 OECD countries over the period 1990:1-2013:11[10]. Monthly data on the bilateral exchange rate of the national

[8] The power analysis for the small values of the break parameter (i.e., $\alpha_2 = 0.5, 2.0$) is available upon request. The ADF test performs better in these regions.

[9] The power analyses for models B and C are available upon request.

[10] We exclude LNV test from empirical analysis due to the reason that LNV type of unit root structure indicates persistent structural break but we are dealing with temporary ones.

currency against the U.S. dollar and on consumer price indices (CPI) were taken from the IFS database of IMF. All variables were put into natural logarithms before the analysis.

**Table 4.** Model A Results[11]

| | Country | $\tilde{s}_{\alpha}$ | $\tau_{DF,t}$ | $\tau_{FDF,c}$ |
|---|---|---|---|---|
| 1 | Austria | **-5.415*** | -2.205 | -3.123 |
| 2 | Belgium | **-5.217*** | -2.116 | -3.126 |
| 3 | Canada | -3.696 | -1.668 | **-3.546***** |
| 4 | Denmark | **-4.907**** | -2.373 | -3.261 |
| 5 | Finland | **-4.481**** | -2.155 | -3.056 |
| 6 | France | **-4.519**** | -2.122 | -2.644 |
| 7 | Greece | -2.756 | -2.809 | **-3.558***** |
| 8 | Hungary | -3.604 | -2.370 | -3.423 |
| 9 | Iceland | -2.624 | -2.622 | **-3.459**** |
| 10 | Israel | **-4.473**** | -1.743 | -3.318 |
| 11 | Italy | **-4.246***** | -2.390 | -3.170 |
| 12 | Japan | -3.580 | -2.761 | -2.981 |
| 13 | Korea | -2.912 | -2.222 | -2.820 |
| 14 | Luxembourg | **-5.107*** | -2.064 | -3.096 |
| 15 | Mexico | **-4.787**** | **-3.500**** | **-3.881**** |
| 16 | Netherlands | **-5.152*** | -2.357 | -3.285 |
| 17 | Norway | **-5.116*** | -2.598 | **-3.888**** |
| 18 | Poland | -3.847 | **-4.359*** | **-4.410*** |
| 19 | Portugal | -3.572 | -2.497 | -3.351 |
| 20 | Spain | -3.348 | -2.348 | -3.266 |
| 21 | Sweden | -2.993 | -2.318 | -2.846 |
| 22 | Switzerland | -3.933 | -2.208 | **-3.518***** |
| 23 | Turkey | **-4.442**** | -2.945 | **-4.539*** |
| 24 | UK | -2.830 | **-3.193***** | **-3.716**** |

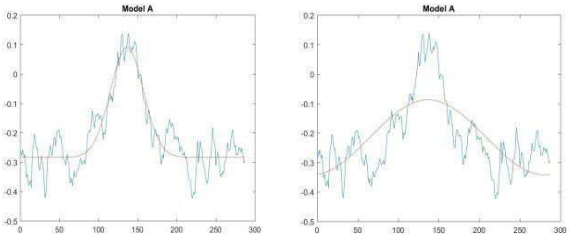
*, ** and *** denote significance at the 1%, 5% and 10%, respectively.

It can be seen from Table 4 that the EST test rejects the unit root null for 12 countries. However, this number falls to 9 and 3 when the Fourier ADF (FADF) and ADF tests are used, respectively.

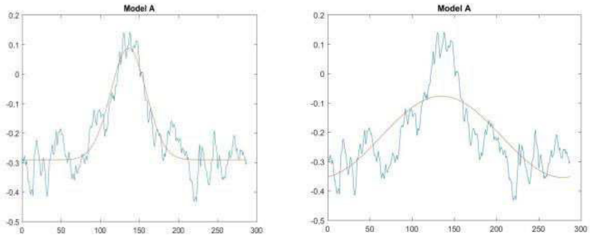
[11] Estimated values of the EST parameters are available upon request.

| EST | Fourier | EST | Fourier |
|---|---|---|---|

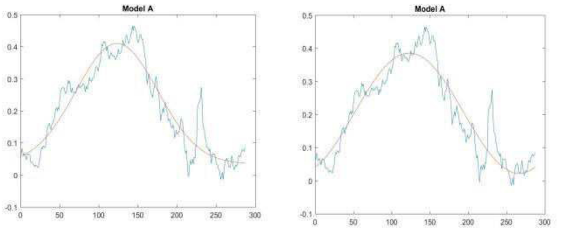
Austria



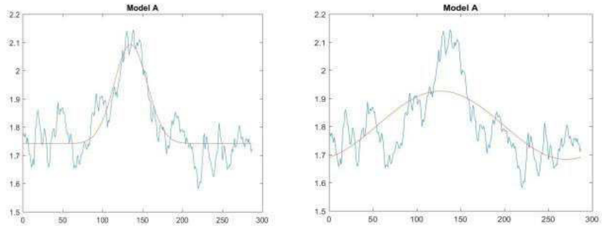

Belgium



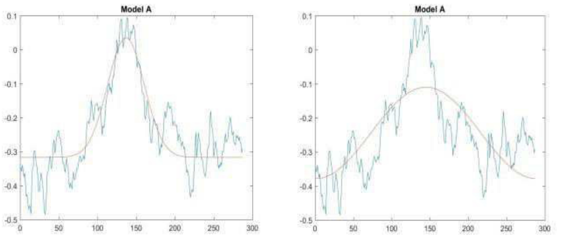

Canada



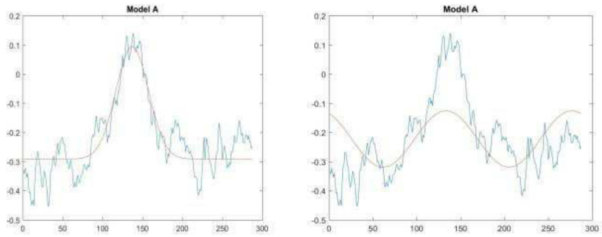

Denmark



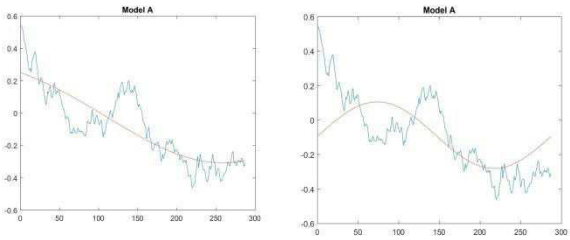

Finland



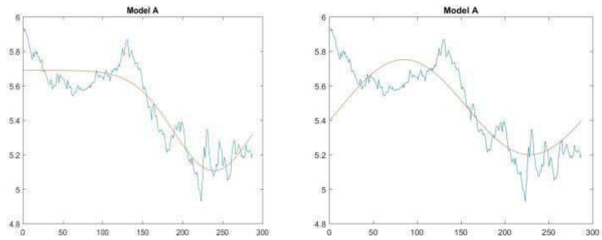

France



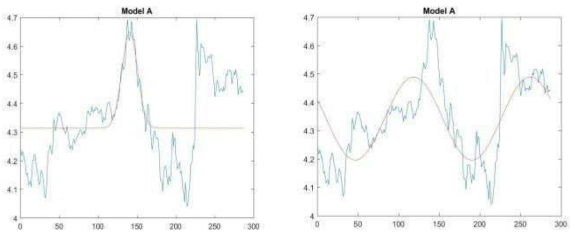

Greece



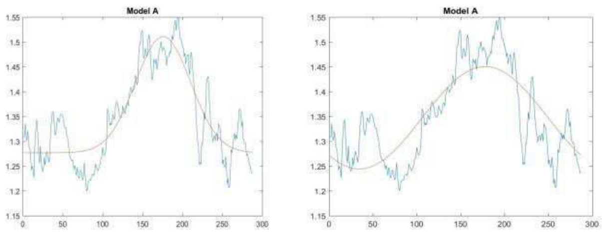

Hungary



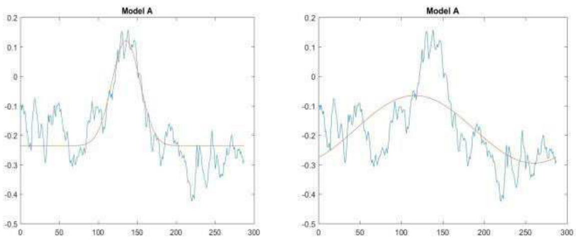

Iceland



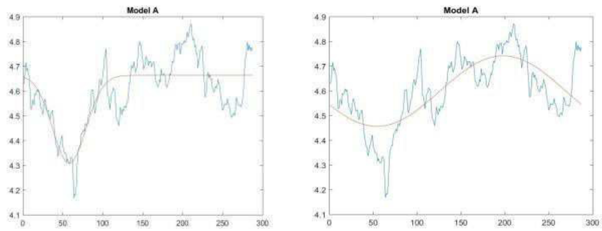

Israel




Italy




Japan




Korea

Luxembourg

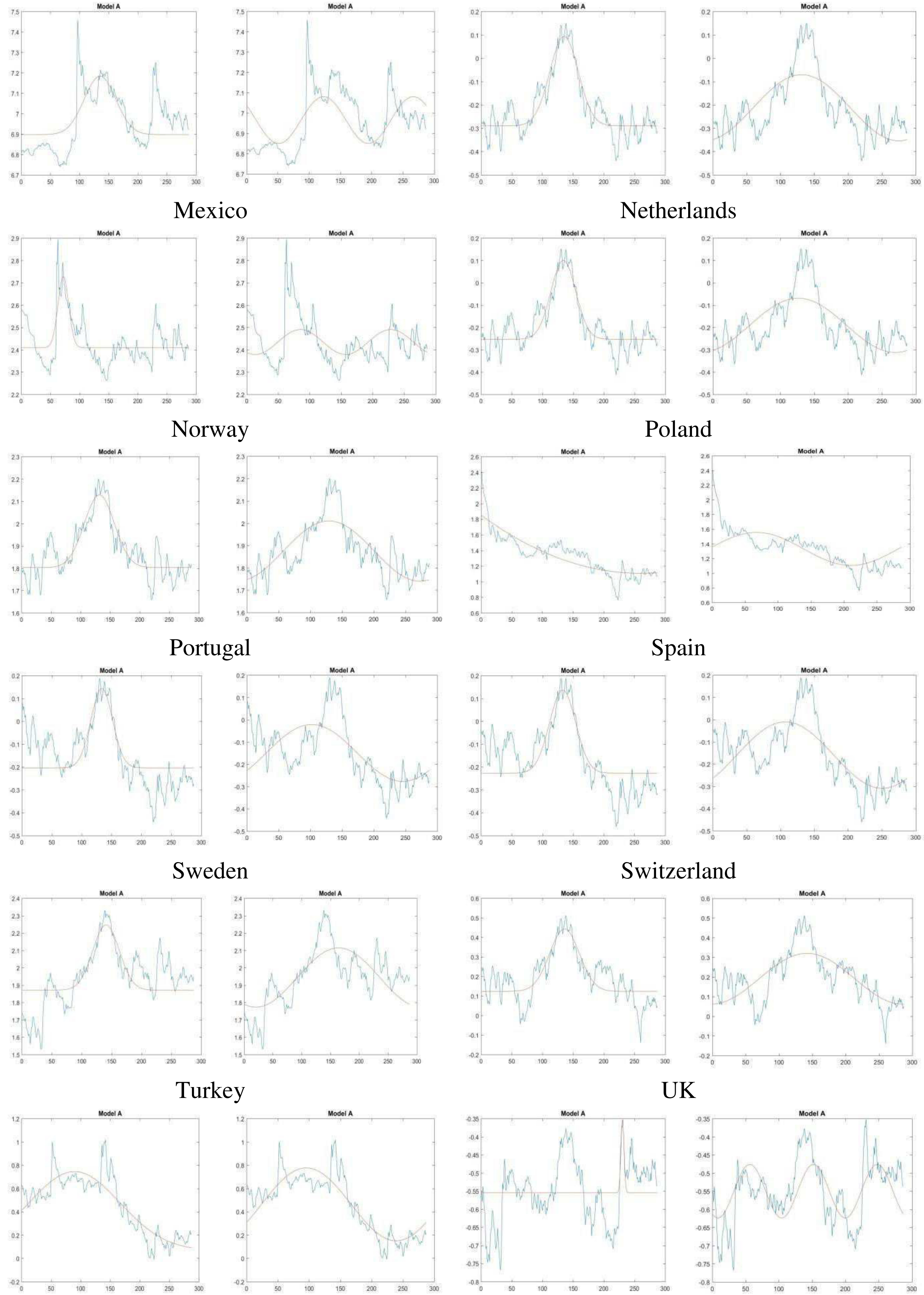



**Figure 1** Comparison of EST and FFF intercepts

The differences between modeling the breaks with EST and FFF are clearly highlighted in Figure 1. To discuss the results we will group the countries into three cases. In the first case we have the group of countries where the PPP hypothesis is found to hold only using the EST test (and not the FADF test). Austria is one of those countries that belongs to this group. From Figure 1 we can clearly see that for Austria the EST type of intercept produces the constant mean that is compatible with the long-run PPP at both ends of the temporary structural break. On the other hand, although the Fourier intercept displays a similar behavior, there is no such thing as a constant mean occurring at both ends of the temporary break. As we mentioned before only the start and end points are same. Same arguments follow for the RERs of Belgium, Denmark, Finland, France, Israel, Italy, Luxembourg, and Netherlands. As a second case we show the counter examples of Canada, Greece, Iceland, Poland, Switzerland, and UK; where this time the FADF test has concluded that the PPP holds. When we look at the Poland data RER has a clear negative trend that the EST model exactly catches. Again in this example the Fourier intercept produces the same start and end values, but the RER does not have a constant mean. Finally, we have the last case which includes those countries where both tests conclude that the PPP holds. This is true for Mexico, Norway and Turkey. Among these countries Mexico provides an interesting example. While the EST intercept consistently finds the temporary break, the Fourier function acts as if there are multiple structural breaks. Overall, if the EST unit root test cannot find a constant mean which is passing through the data, the standard PPP hypothesis with temporary break will not hold.

## 4. Conclusion

In this study we have showed that the true approach for testing the standard PPP hypothesis by allowing for temporary breaks is to use a unit root test that employs an EST function. This study has, therefore, raised important questions about the usage of the FFF in assessing the validity of the PPP hypothesis.

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