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# A Nonlinear View of Long-Run PPP Using Cross-Sectionally Dependent Heterogeneous Panels

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Empirical studies have flourished on long-run PPP using panel unit root tests, which supposedly have high power. In this paper, we adopt, in a panel data context, a nonlinear multiple-regime model, namely Threshold Autoregression (TAR), and perform a panel unit root test for each regime in the TAR. This new procedure takes advantage of two existing approaches: the inference for a TAR model with a unit root, and the panel unit root tests with the augmented panel Dickey-Fuller regression. The real exchange rate dynamics in a panel of 17 OECD countries over the recent floating exchange rate period are investigated. Three distinct regimes are identified. In particular, the support for long-run PPP is much stronger in the post-Maastricht-Treaty period than the period before; in the pre-Maastricht-Treaty period, there is some evidence for long-run PPP when the sterling-dollar appreciated, while the evidence is weak when the sterling-dollar depreciated. Our results are robust to expanding the data set to 2012, using the euro foreign exchange reference rates.

Keywords: asymmetric adjustment, cross-sectional dependence, long-run PPP, real exchange rate, threshold autoregression (TAR)
JEL classification: C33, F31, G15

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## 1 Introduction

A nonlinear model such as the threshold autoregression (TAR, Tong (1978))[1] has long been used to capture the *asymmetric* adjustment to some *threshold variable*, such as lagged GDP growth, or lagged exchange rate. Along this line of research, Lin and Teräsvirta (1994) suggest to use the time index as the threshold variable and devise a test for structural change with this type of models.

This paper uses a TAR model to capture (i) the structural change; and (ii) the asymmetric adjustment, of the long-run PPP. The PPP theory is an important as well as challenging topic in international macroeconomic analysis and open economy models, and is of great practical relevance to policy making (Rogoff, 1996; Sarno and Taylor, 2002; Taylor and Taylor, 2004). Empirically speaking, if the (relative) long-run PPP holds, the real exchange rate is mean-reverting, regardless of there being structural change and/or asymmetric adjustment.

Since the end of the Bretton Woods system in 1973, many countries have undergone a number of structural changes. In particular, the exchange rate mechanism of the European Monetary System (ERM) had explicit finite bands within which exchange rates were allowed to fluctuate. The bands within ERM were $\pm 2.25$ percent around a cental parity. Due to the changes in market conditions, or otherwise, realignments or shifts in the central parity occasionally took place in the late 1970s and in the 1980s. In particular, the degree of sterling devaluation against the dollar is much more marked during the 1980s (see the last diagram in Figure 1b). The strong appreciation of the dollar in the early 1980s was then reverted rapidly after the G5 Plaza Accord of 1985.

Lopez and Papell (2007) present strong evidence in favor of the PPP for the full panel of euro area currencies after the adoption of the Maastricht Treaty in 1992. The Treaty established the convergence criteria defining access to a single currency, and required the member countries achieve exchange rate and inflation stability. The continued process of economic integration in Europe accelerated convergence to PPP in the post-Maastricht Treaty period. Assuming a known structural change point (year of 1999), Koedijk, Tims, and van Dijk (2004) further point out, that such converge-

[1] An alternative is to use the smooth transition autoregression (STAR, Teräsvirta (1994)) and to take a Taylor's expansion. See Cerrato et al. (2011).

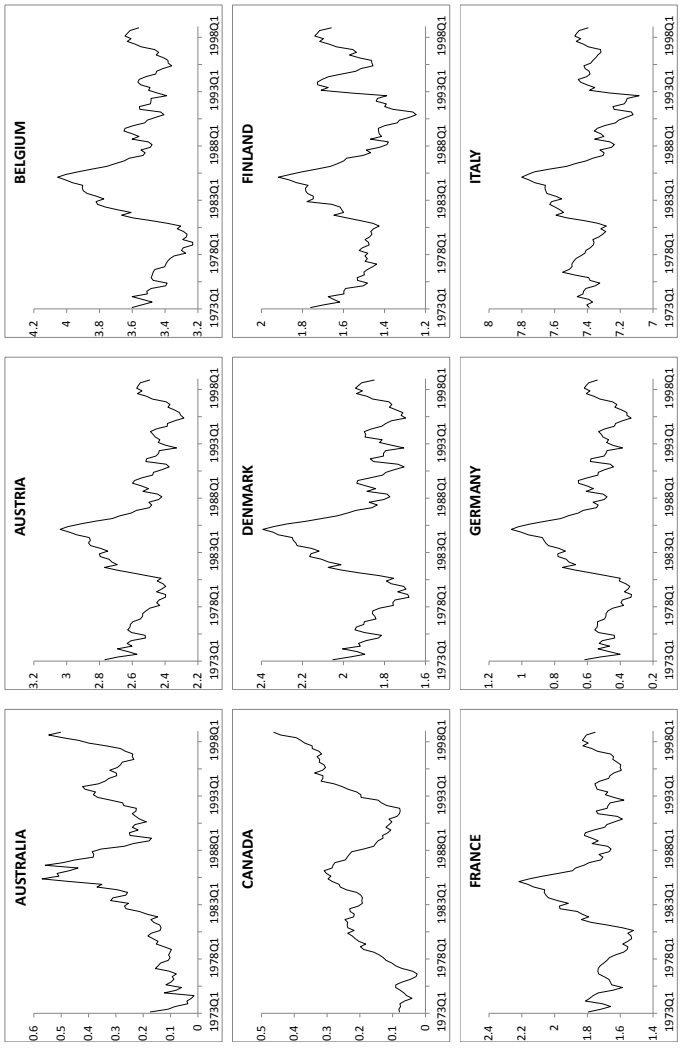



Figure 1a: Real exchange rates relative to the US dollar 1973Q1–1998Q4

Notes: Each plot depicts the real exchange rate of one of the 17 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, and the UK.

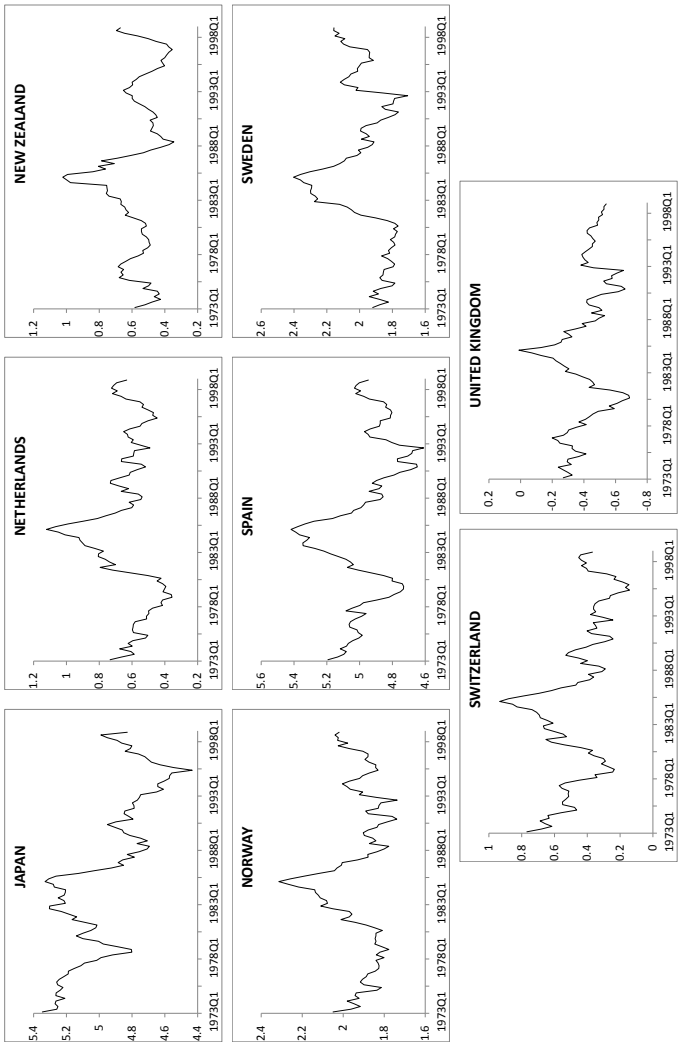



Figure 1b: Real exchange rates relative to the US dollar 1973Q1–1998Q4

Notes: Each plot depicts the real exchange rate of one of the 17 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, and the UK.

nce process is consistent with a transactions-costs goods-market arbitrage view of the mean reversion properties of the real exchange rate. As Eichengreen (2003) suggests, Denmark's rejection of the Maastricht Treaty in the mid-1991 and the tightening of German monetary policy after the reunification in 1990 contributed to a wave of speculative attacks on currency pegs in Europe. Eventually, two member currencies, the Italian lira and the British pound, were forced to leave the system, while the Spanish peseta, the Portuguese escudo and the Irish punt were devalued involuntarily.

Parallel to the above literature, there have been discussions on whether the real exchange rate adjusts asymmetrically. Theoretically, one can show in a general equilibrium model with transaction costs, that real exchange rate adjusts in an asymmetric fashion (Uppal, 1993). When the asymmetric adjustment is due to transaction costs, profitable commodity arbitrage opportunities can be exploited (Michael, Nobay, and Peel, 1997; Taylor and Peel, 2000). Other plausible explanations of asymmetry include exchange rate intervention (Taylor, 2004) and heterogeneous agents (Kilian and Taylor, 2003). Applications of TAR models (or STAR models) to exchange rates include Obstfeld and Taylor (1997), O'Connell (1998) and Sarno, Taylor, and Chowdhury (2004). They all find evidence supporting the asymmetric adjustment process.

While the literature on TAR models by and large assumes mean-reversion, Caner and Hansen (2001) develop a test for thresholds in a TAR model that may or may not contain a unit root. In other words, their test is robust to data which are *not* mean-reverting. Moreover, this threshold test is general enough to allow some structural change which takes place at *unknown* points.[2] This approach allows us to test statistically the structural change point(s) such as that suggested in Koedijk, Tims, and van Dijk (2004). In a pure time-series context, their threshold test is applied to some real exchange rates to see if nonlinearity and/or unit root exists (Basci and Caner, 2005).[3] Following the strand of the time series literature surrounding struc-

[2]As suggested in Bai (2010), common breaks are generally observed in various economic situations and provide computational advantages over unknown change points.

[3]In a pure time-series context, Bec, Salem, and Carrasco (2004) consider a model similar to that in Caner and Hansen (2001). However, their paper focuses on testing a unit root against a threshold alternative. In contrast, this paper first considers a threshold test, following the lines in Caner and Hansen (2001), which is robust to the existence of unit roots.

tural changes and unit-root tests, to name a few examples, Banerjee, Lumsdaine, and Stock (1992), Christiano (1992), Zivot and Andrews (1992) and Lumsdaine and Papell (1997), all agree that standard (purely time-series) unit root tests have low power against trend-stationary alternatives with structural changes, and the change points should be data-dependent.

Recently, there is a growing literature on linear panel unit root test which allows cross-sectional dependence and heterogeneity.[4] Noticeably, Pesaran (2007) (see also Pesaran (2006)) proposes a unit root test using an error factor structure, where the cross-sectional averages of dependent variable and regressors are added to the traditional augmented Dickey-Fuller (ADF) equations. In so doing, the cross-sectional averages are able to capture this common factor. This will filter out the common factor for the errors from the ADF equations, and thus adjust for the presence of cross-sectional dependence. This method is easy to implement and the literature documents that it works well in finite-samples. Wu, Cheng, and Hou (2011) adopt Pesaran (2007) test to re-examine the effects of country characteristics on the long-run PPP. They find the long-run PPP holds for a panel of 76 countries, and the panel of Africa and Latin America, when cross-sectional dependence and heterogeneity are taken into account.

Under this backdrop, this paper proposes a nonlinear view of long-run PPP by modifying the threshold test (Caner and Hansen, 2001) and the panel unit root test (Pesaran, 2007). The threshold test is first modified (see sub-section 2.4 below) such that cross-sectional dependence and heterogeneity are allowed. Given the number of regimes and the threshold points, the panel unit root test is modified (see sub-section 2.5 below) accordingly.

The next section exposits our proposed approach for panel unit root tests, with a small-scale Monte Carlo experiment which is discussed in the Appendix. Section 3 applies our approach to the real exchange rates of 17 OECD countries. Section 4 concludes.

## 2 The Methodology

This section is divided into 5 sub-sections. Sub-section 2.1 contains a brief review of the linear panel unit root test. The panel TAR model and the

[4] Unit root tests in heterogenous panels are traced back to Im, Pesaran, and Shin (2003). See Maddala and Wu (1999) for a Monte Carlo study; and Wu and Wu (2001) for an application to the long-run PPP.

estimation method can be found in the two subsequent sub-sections. Sub-section 2.4 is about the threshold test while sub-section 2.5 contains the nonlinear panel unit root test.[5]

## 2.1 The linear panel unit root test: a brief review

In this sub-section, we briefly discuss the procedure proposed by Pesaran (2007). Consider the following dynamic linear heterogenous model with $N$ cross-sectional units and $T$ time series units:

$$\Delta y_{i,t} = \alpha_i + \beta_i y_{i,t-1} + \gamma_i f_t + u_{i,t}, \tag{1}$$

$$u_{i,t} = \sum_{j=1}^{\tau} \delta_j \Delta y_{i,t-j} + e_{i,t}, \tag{2}$$

where $e_{i,t}$'s are $i.i.d.$ across $i$ and $t$.[6] If $\bar{\gamma} = \sum_{i=1}^{N} \gamma_i / N \neq 0$ and assume that $\beta_i = \beta$, which is the case under the null of unit root. See (17) and (19) below. Thus we have the so-called CADF (cross-sectionally augmented Dickey-Fuller) regression:

$$\Delta y_{i,t} = \alpha_i + \beta_i y_{i,t-1} + \sum_{j=1}^{\tau} \delta_j \Delta y_{i,t-j} + b_i \bar{y}_{t-1} + \sum_{j=0}^{\tau} c_{i,j} \bar{\Delta} y_{t-j} + \epsilon_{i,t}. \tag{3}$$

That is, for each $i$, one regresses $\Delta y_{i,t}$ on 1, $y_{i,t-1}$, $\Delta y_{i,t-j}$, $j = 1, \cdots, \tau$, $\bar{y}_{t-1}$, and $\bar{\Delta} y_{t-j}$, $j = 1, \cdots, \tau$. The CIPS[7] is defined as:

$$\text{CIPS} = \frac{1}{N} \sum_{i=1}^{N} \frac{\hat{\beta}_i}{\hat{se}\left(\hat{\beta}_i\right)}, \tag{4}$$

[5] It is unclear if the test in sub-section 2.5 should be called nonlinear panel unit root test, as under each regime, the model is linear. We call it nonlinear panel unit root test, only to distinguish it from the one discussed in sub-section 2.1.

[6] It should be noted that the common dynamic parameter $\delta_j^{(k)}$'s in (2) can be relaxed with some modification of the expression of the common factor. See, for instance, Pesaran (2007).

[7] This is the cross-sectionally augmented version of the IPS, where IPS stands for the test suggested in Im, Pesaran, and Shin (2003).

the cross-sectional average of the $t$-ratio of the OLS estimates of the $\beta_i$'s.[8]

## 2.2 The $K$-regime panel TAR model

In this paper, the $K$-regime panel TAR model is specified as:

$$\Delta y_{i,t} = \sum_{k=1}^{K} \left( \alpha_i^{(k)} + \beta_i^{(k)} y_{i,t-1} + \gamma_i^{(k)} f_t^{(k)} + u_{i,t}^{(k)} \right) \mathbf{1}\left( \underline{\lambda}^{(k)} < q_{t-1}^{(k)} \leq \overline{\lambda}^{(k)} \right), \tag{5}$$

where $K = 2, 3, \cdots$, and $\mathbf{1}(A) = 1$ if $A$ occurs and 0, otherwise.[9] Needless to say, when $K = 1$, the panel TAR model boils down to a panel AR model as that in equation (3).

With the *thresholds* $\underline{\lambda}^{(k)}$ and $\overline{\lambda}^{(k)}$, the *k-th threshold variable* $q_{t-1}^{(k)}$ is a univariate variable which is observable in period $t-1$. For instance, $q_{t-1}^{(k)} = y_{1,t-d}, d \geq 1$ (see, in a pure time-series context where $N = 1$, Tong (1978), or Teräsvirta (1994)). In our panel data context, we may consider $q_{t-1}^{(k)} = y_{g,t-d}, d \geq 1$ ($y_{g,t}$ may be called a "global variable" such as the world interest rate), or $q_{t-1}^{(k)} = t$ (see Lin and Teräsvirta (1994)). To proceed, we further assume:

$$u_{i,t}^{(k)} = \sum_{j=1}^{\tau} \delta_j^{(k)} \Delta y_{i,t-j} + e_{i,t}, \tag{6}$$

where $e_{i,t}$'s are $i.i.d.$ across $i$ and $t$. For each $t = 1, \cdots, T$ and each $k = 1, \cdots, K$, let $N_t^{(k)}$ be the number of cross-sectional units in Regime $k$ and we further assume

$$\bar{\gamma}_t^{(k)} = \frac{1}{N_t^{(k)}} \sum_{i \in \text{Regime } k} \gamma_i^{(k)} \to \bar{\gamma}^{(k)} \neq 0, \text{ as } N \to \infty. \tag{7}$$

[8] De Silva, Hadri, and Tremayne (2009) demonstrate the test by Pesaran (2007) is conditional on two important assumptions: (a) The average of the factor loading parameters needs to be different from zero and (b) there can only be one single factor in the model. When these assumptions are not satisfied, the test displays poor empirical performance. Hence, this will also be a drawback of the test introduced in this paper. The authors are grateful to an anonymous referee for pointing this out.

[9] As one can see in section 3 "Empirical results" below, some of the regimes are confined to a sub-sample. And "The estimation procedure", "The threshold test" and "The nonlinear panel unit root test" are modified accordingly.

## 2.3 The estimation procedure

Re-write (5) as:

$$\Delta y_{i,t} = \sum_{k=1}^{K} \sum_{\underline{\lambda}^{(k)} < q_{t-1}^{(k)} \le \bar{\lambda}^{(k)}} \Bigg( a_i^{(k)} + \beta_i^{(k)} y_{i,t-1} + \sum_{j=1}^{\tau} \delta_j^{(k)} \Delta y_{i,t-j} + b_i^{(k)} \bar{y}_{t-1}^{(k)} + \sum_{j=1}^{\tau} c_{i,j}^{(k)} \bar{\Delta} y_{t-j}^{(k)} \Bigg) + \epsilon_{i,t}, \tag{8}$$

where

$$\begin{aligned} \bar{y}_{t-1}^{(k)} &:= \frac{1}{N_t^{(k)}} \sum_{i \in \text{Regime } k} y_{i,t-1}, \\ \bar{\Delta} y_{t-j}^{(k)} &:= \frac{1}{N_t^{(k)}} \sum_{i \in \text{Regime } k} \Delta y_{i,t-j},\ j = 0, \cdots, \tau; \qquad (9) \\ a_i^{(k)} &:= \alpha_i^{(k)} - \lim_{N \to \infty} \frac{1}{\bar{\gamma}^{(k)} N_t^{(k)}} \sum_{i \in \text{Regime } k} \alpha_i^{(k)}, \\ b_i^{(k)} &:= \beta_i^{(k)} \gamma_i^{(k)} / \bar{\gamma}^{(k)}, \quad c_{i,0}^{(k)} := \gamma_i^{(k)} / \bar{\gamma}^{(k)}; \\ c_{i,j}^{(k)} &:= \delta_j^{(k)} \gamma_i^{(k)} / \bar{\gamma}^{(k)}, \quad j = 1, \cdots, \tau. \qquad (10) \end{aligned}$$

The estimation procedure for $K = 1$ can be found in sub-section 2.1. For $K = 2, 3, \cdots$, in view of (8), we consider the following procedure of estimation:

Step I: For each $k = 1, \cdots, K$, sort the $q_{t-1}^{(k)}$'s ($t = 1, \cdots, T$) in ascending order such that $q_{(1)}^{(k)} \le q_{(2)}^{(k)} \le \cdots \le q_{(T)}^{(k)}$.

Step II: For each $k = 1, \cdots, K$, let $\underline{\lambda}^{(k)}, \bar{\lambda}^{(k)} \in \Lambda = \{-\infty = q_{(0)}^{(k)}, q_{(1)}^{(k)}, q_{(2)}^{(k)}, \cdots, q_{(T)}^{(k)}, q_{(T+1)}^{(k)} = \infty\}$. For each $i = 1, \cdots, N$, perform the following linear regression:

$$\Delta y_{i,t} = a_i^{(k)} + \beta_i^{(k)} y_{i,t-1} + \sum_{j=1}^{\tau} \delta_j^{(k)} \Delta y_{i,t-j} + b_i^{(k)} \bar{y}_{t-1}^{(k)} + \sum_{j=1}^{\tau} c_{i,j}^{(k)} \bar{\Delta} y_{t-j}^{(k)} + \epsilon_{i,t}^{(k)}, \tag{11}$$

where $\underline{\lambda}^{(k)} < q_{t-1}^{(k)} \leq \overline{\lambda}^{(k)}$, whenever the estimators exist.[10] Denote the estimator by $\hat{\phi}_i^{(k)}(\lambda^{(k)})$ and the residual by $\hat{\epsilon}_{i,t}^{(k)}(\lambda^{(k)})$, where $\lambda^{(k)} = (\underline{\lambda}^{(k)}, \overline{\lambda}^{(k)})'$.
Step III: Let $\lambda_K = (\lambda^{(1)'}, \cdots, \lambda^{(K)'})'$ and $\mathcal{L} = \{\lambda_K : \hat{\phi}_i^{(k)}(\lambda^{(k)}) \text{ exists } \forall i, k\}$. Define $\hat{\lambda}_K = (\hat{\lambda}^{(1)'}, \cdots, \hat{\lambda}^{(K)'})' = \arg\sup_{\lambda \in \mathcal{L}} -\hat{S}_{NT}(\lambda)$, where $\hat{S}_{NT}(\lambda)$ is the residual sum of squares (RSS), defined as:

$$\hat{S}_{NT}(\lambda) = \sum_{k=1}^{K} \hat{S}_{NT}^{(k)}(\lambda), \;\; \hat{S}_{NT}^{(k)}(\lambda) = \sum_{i=1}^{N} \sum_{\underline{\lambda}^{(k)} < q_{t-1}^{(k)} \leq \overline{\lambda}^{(k)}} \left( \hat{\epsilon}_{i,t}^{(k)} \left( \lambda^{(k)} \right) \right)^2. \quad (12)$$

For $i = 1, \cdots, N$, the estimators chosen are:

$$\hat{\phi}_i \left( \hat{\lambda}_K \right) = \left( \hat{\phi}_i^{(1)} \left( \hat{\lambda}^{(1)} \right), \hat{\phi}_i^{(2)} \left( \hat{\lambda}^{(2)} \right), \cdots, \hat{\phi}_i^{(K)} \left( \hat{\lambda}^{(K)} \right) \right)', \quad (13)$$

and when no ambiguity arises, we write:

$$\hat{\phi}_i^{(k)} \left( \hat{\lambda}^{(k)} \right) = \hat{\phi}_i^{(k)} = \left( \hat{a}_i^{(k)}, \hat{\beta}_i^{(k)}, \hat{\delta}_{i,1}^{(k)}, \cdots, \hat{\delta}_{i,\tau}^{(k)}, \hat{b}_i^{(k)}, \hat{c}_{i,1}^{(k)}, \cdots, \hat{c}_{i,\tau}^{(k)} \right)', \quad k = 1, \cdots, K. \quad (14)$$

□

## 2.4 The threshold test

Our threshold test takes the following null and alternative hypotheses: For $K = 1, 2, \cdots,$

$$H_{0,I} : K \text{ regimes}, \quad H_{A,I} : K + 1 \text{ regimes}. \quad (15)$$

We use the following sup-statistic (Caner and Hansen, 2001) to determine the *number of thresholds*:

$$W_{NT}(K + 1 \mid K) = \frac{T \left[ \hat{S}_{NT} \left( \hat{\lambda}_K \right) - \hat{S}_{NT} \left( \hat{\lambda}_{K+1} \right) \right]}{\hat{S}_{NT} \left( \hat{\lambda}_{K+1} \right)}, \quad (16)$$

where $\hat{S}_{NT}(\hat{\lambda}_K)$ and $\hat{S}_{NT}(\hat{\lambda}_{K+1})$ are the RSS's, as defined in (12), for $K$ and $K+1$ regimes respectively. The asymptotic distribution of $W_{NT}(K+1 \mid K)$

[10] Should the sub-sample size of one of the $i$ in regime $k$ is smaller than the number of parameters, the estimators do not exist.

is nonstandard. The $p$-value and critical value, which in turn is robust to I(0)/I(1) series, are obtained via a *residual-based* bootstrap procedure.[11]

If we are unable to reject $H_{0,I}$ when $K = 1$, we will then conduct the panel unit root test in sub-section 2.1. Otherwise, we conduct the nonlinear panel unit root test using (5), which is the subject in the next sub-section.

## 2.5 The nonlinear panel unit root test

For each regime $k$, the unit root test is:

$$\begin{aligned} &H_{0,II} : \beta_i^{(k)} = 0, \forall i = 1, \cdots, N; \\ &H_{A,II} : -2 < \beta_i^{(k)} < 0, \exists i = 1, \cdots, N. \end{aligned} \tag{17}$$

It should be noted that in our analysis of PPP, we intend to examine whether, under the alternative, at least some countries obeys the PPP. In other words, our hypothesis sheds light on the validity of PPP in a number of countries.[12] Similar to (4), for regime $k, k = 1, \cdots K$, we define the nonlinear panel unit root test statistic, called nonlinear CIPS,[13]

$$\text{NCIPS}^{(k)} = \frac{1}{N} \sum_{i=1}^{N} \frac{\hat{\beta}_i^{(k)}}{\hat{se}\left(\hat{\beta}_i^{(k)}\right)}, \tag{18}$$

that is, the cross-sectional average of the $t$-ratio of the OLS estimates of the $\beta_i^{(k)}$'s in (11).

For sake of comparison we also consider a *full-sample* test in which one single regime ($K = 1$) is assumed. The nonlinear panel unit root test associated with (3) is stated as:

$$\begin{aligned} &H_{0,II} : \beta_i = 0, \forall i = 1, \cdots, N; \\ &H_{A,II} : -2 < \beta_i < 0, \exists i = 1, \cdots, N. \end{aligned} \tag{19}$$

[11] See Caner and Hansen (2001) for the details.

[12] As argued by Choi (2004), the rejection of the unit root null under the panel test is not informative as regard to the case of the stationary series in the panel. Instead, he suggests a sequential algorithm to separate the panel into $I(0)$ and $I(1)$ series. However, since interest usually lies in testing whether there is at least some evidence for PPP, a test of our $H_{0,II}$ versus $H_{A,II}$ is still informative.

[13] It should be noted that while the test statistic is called nonlinear CIPS, under each regime, the model is linear. See also footnote 5.

The full-sample NCIPS (FNCIPS) is thus defined as:

$$\text{FNCIPS} = \frac{1}{K}\sum_{k=1}^{K}\frac{1}{N}\sum_{i=1}^{N}\frac{\hat{\beta}_i^{(k)}}{\hat{se}\left(\hat{\beta}_i^{(k)}\right)}. \tag{20}$$

The $p$-values and critical values of these two statistics are also obtained via a residual-based bootstrap procedure.

## 3 Empirical Results

Section 3.1 below is our major study which uses data from 1973 to 1998, the year right before the introduction of Euro. This data span is commonly used in PPP studies as it permits the use of a number of developed European economies which subsequently entered the single currency. Comparisons of our testing results to the existing evidence for PPP can also be found. In subsection 3.2, we extend our data through 2012, which allows us to evaluate the impact of the launch of the euro as well as the global financial crisis of 2008.

### 3.1 Basic results and a comparison with results by other panel unit root tests

Refer to the panel TAR model in sub-section 2.2. We apply the model to 17 currencies of OECD countries consisting of Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland and the UK. The data are quarterly, spanning from 1973Q1 to 1998Q4. The sample ends at the period before the introduction of the euro in 1999, when the nominal exchange rates among the European countries became irrevocably fixed. The length of the data is subject to the fact when a number of the countries in our sample joined the European Monetary Union (EMU) and shared a common currency. The bilateral real exchange rate against the US dollar at time $t$ (in logarithm) is constructed as

$$y_{i,t} = s_{i,t} - p_{i,t} + p_{US,t}, \tag{21}$$

where $s_{i,t}$ is the logarithm of the nominal exchange rate of the $i$-th country ($i$th currency units per one US dollar), $p_{i,t}$ is the logarithm of the CPI of

the $i$-th country, and $p_{US,t}$ is the logarithm of the US CPI. In other words, a rise in $y_{i,t}$ implies a real depreciation of the $i$th currency. We use the data set considered in Choi (2004), Murray and Papell (2002) and Murray and Papell (2005) and Pesaran (2007). The data are available from http://www.econ.queensu.ca/jae/2007-v22.2/pesaran/. Figures 1a and 1b depict the logarithm of real exchange rates against the US dollar over the period 1973Q1 to 1998Q4. Two striking observations are made. First, the real exchange rates for these countries appear to move around their mean values relatively infrequently and are inconsistent with short-run mean-reverting behavior. Second, it appears that there was an obvious spike in the exchange rate series during the mid 1980s, and the fluctuations of the real exchange rates become substantially smaller in the early 1990s. In particular, since the Maastricht Treaty was signed in 1992, the real exchange rates have been relatively stable. For instance, Austria, France, Germany, Italy, the Netherlands, and the UK, the graphs suggest the real exchange rates have exhibited mean-reverting behavior in the last two decades or so. Table 1 shows the pairwise correlations between real exchange rates, which indicates that correlations are substantially positive. This suggests that cross-sectional dependence may be present in the data.

As a starting point, the real exchange rate series are jointly tested for unit roots using a number of conventional tests, namely, the LLC test (Levin, Lin, and Chu, 2002), the IPS test (Im, Pesaran, and Shin, 2003), the Fisher-ADF test (Maddala and Wu, 1999) and the PP test (Choi, 2001). The lag length $\tau$ is selected with the Schwarz Information Criterion (SIC) with the maximum length of 4. The test statistics are reported in Table 2. These findings are representative for those of studies using panel unit root tests *assuming* no cross-sectional dependence. It remains unclear whether the non-rejection of unit root is due to a failure of PPP, the low power of the tests, or the presence of cross-sectional dependence. Table 2 also reports the CIPS test of the linear model. The unit root hypothesis is not rejected at any sensible level. Indeed, the standard IPS statistic leads to rejection of the null, whereas the augmented one leads to non-rejection of the unit root hypothesis. This is consistent with the previous findings (for instance, O'Connell (1998)) that imposing the restriction of cross-sectional independence causes over-rejection of the unit root null. Though, on the other hand, it is unclear if the power of the CIPS test is affected, when nonlinearity is present.

Table 3 reports the results of a number of cross-sectional dependence

Table 1: Correlations of OECD real exchange rates

| | AUS | AUT | BEL | CAN | DEN | FIN | FRA | GER | ITA | JAP | NET | NOR | NAZ | SPA | SWE | SWI | UKI |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| AUS | 1.000 | | | | | | | | | | | | | | | | |
| AUT | 0.169 | 1.000 | | | | | | | | | | | | | | | |
| BEL | 0.510 | 0.874 | 1.000 | | | | | | | | | | | | | | |
| CAN | 0.719 | −0.011 | 0.255 | 1.000 | | | | | | | | | | | | | |
| DEN | 0.319 | 0.975 | 0.941 | 0.135 | 1.000 | | | | | | | | | | | | |
| FIN | 0.418 | 0.711 | 0.715 | 0.445 | 0.776 | 1.000 | | | | | | | | | | | |
| FRA | 0.443 | 0.909 | 0.966 | 0.210 | 0.959 | 0.754 | 1.000 | | | | | | | | | | |
| GER | 0.437 | 0.927 | 0.968 | 0.185 | 0.961 | 0.664 | 0.963 | 1.000 | | | | | | | | | |
| ITA | 0.205 | 0.827 | 0.711 | 0.259 | 0.848 | 0.813 | 0.810 | 0.755 | 1.000 | | | | | | | | |
| JAP | −0.324 | 0.758 | 0.405 | −0.372 | 0.632 | 0.395 | 0.477 | 0.506 | 0.572 | 1.000 | | | | | | | |
| NET | 0.482 | 0.894 | 0.990 | 0.223 | 0.947 | 0.716 | 0.964 | 0.975 | 0.715 | 0.439 | 1.000 | | | | | | |
| NOR | 0.501 | 0.865 | 0.900 | 0.393 | 0.924 | 0.905 | 0.919 | 0.880 | 0.841 | 0.470 | 0.906 | 1.000 | | | | | |
| NZA | 0.425 | 0.766 | 0.682 | 0.131 | 0.779 | 0.631 | 0.743 | 0.764 | 0.717 | 0.539 | 0.706 | 0.757 | 1.000 | | | | |
| SPA | 0.196 | 0.902 | 0.781 | 0.105 | 0.905 | 0.806 | 0.857 | 0.803 | 0.901 | 0.667 | 0.785 | 0.862 | 0.692 | 1.000 | | | |
| SWE | 0.670 | 0.692 | 0.886 | 0.562 | 0.811 | 0.816 | 0.877 | 0.822 | 0.750 | 0.169 | 0.853 | 0.906 | 0.613 | 0.744 | 1.000 | | |
| SWI | 0.026 | 0.957 | 0.766 | −0.168 | 0.904 | 0.649 | 0.818 | 0.823 | 0.755 | 0.832 | 0.801 | 0.770 | 0.701 | 0.871 | 0.541 | 1.000 | |
| UKI | 0.108 | 0.689 | 0.569 | −0.117 | 0.686 | 0.625 | 0.678 | 0.581 | 0.737 | 0.521 | 0.578 | 0.667 | 0.619 | 0.827 | 0.547 | 0.704 | 1.000 |

Notes: This table shows correlations of quarterly real exchange rates against the US dollar measured in logarithms for 17 OECD countries, spanning the period 1973Q1 to 1998Q4. Each column represents the real exchange rate of each individual country, ordered as follows: Australia (AUS), Austria (AUT), Belgium (BEL), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (GER), Italy (ITA), Japan (JAP), the Netherlands (NET), Norway (NOR), New Zealand (NZA), Spain (SPA), Sweden (SWE), Switzerland (SWI), and the UK (UKI).

Table 2: Linear panel unit root tests[a]

| Method | Statistics |
|---|---|
| LLC | 0.332 [0.630][b] |
| IPS | −1.974 [0.024] |
| Fisher-ADF | 43.114 [0.136] |
| PP | 45.827 [0.085] |
| CIPS[c] | −1.936 [0.300] |

Notes: [a] LLC: Levin, Lin, and Chu (2002); IPS: Im, Pesaran, and Shin (2003); Fisher-ADF: Maddala and Wu (1999); PP: Choi (2001). LLC and IPS are asymptotically standard normal; Fisher-ADF and PP are asymptotically $\chi^2_{2N}$, where $N = 17$ is the number of cross-sectional units.
[b] $p$-values are in brackets.
[c] The CIPS and its $p$-value are from Pesaran (2007), whose dataset is the same as ours.

Table 3: Cross-sectional dependence tests[a,b,c]

| | $t \leq$ 1992Q3 | 1992Q3 $< t \leq$ 1998Q4 | $t \leq$ 1998Q4 |
|---|---|---|---|
| LM | 273.508 [0.000][d] | 76.867 [0.000] | 331.718 [0.000] |
| CD | 62.101 [0.000] | 33.309 [0.000] | 65.771 [0.000] |
| $LM_{adj}$ | 197.788 [0.000] | 58.327 [0.000] | 238.829 [0.000] |

Notes: [a] Full sample: 1973Q1–1998Q4.
[b] LM: Breusch and Pagan (1980); CD: Pesaran (2004); $LM_{adj}$: Pesaran, Ullah, and Yamagata (2008).
[c] LM is asymptotically $\chi^2_1$; CD and $LM_{adj}$ are asymptotically standard normal.
[d] $p$-values are in brackets.

tests, namely, the Lagrange Multiplier (LM) test (Breusch and Pagan, 1980), the cross-sectional dependence (CD) test (Pesaran, 2004), and the bias-adjusted Lagrange Multiplier ($LM_{adj}$) test (Pesaran, Ullah, and Yamagata, 2008). All three tests suggest the presence of cross-sectional dependence.

Next we perform the threshold test, which is defined as in (16).[14] For the threshold variable, we consider both the time index $t$ and a "global variable" $y_{UK,t-d}$, $d \geq 1$ or $y_{Germany,t-d}$, $d \geq 1$. We also consider the first difference of the lagged dependent variables, but they do not produce smaller RSS than the level counterparts or fail to identify significant change points.

As one can see from the upper panel of Table 4, 2 regimes are strongly (at the 5% significance level) supported with change point at 1992Q3; while there is no significant evidence (at the 10% significance level) of 3 regimes with change points at 1992Q3 and 1979Q4. Given the structural change point at 1992Q3, we seek for further sample splitting by repeating the procedure. Recall that the estimation procedure in sub-section 2.3 can also be applied to a sub-sample. See also the discussion in footnote 9. More precisely, we examine whether the lagged UK or the lagged Germany dependent variable identifies any asymmetric adjustment prior to 1992Q3, and after then. As shown in the lower panel of Table 4, $y_{UK,t-1}$ further identifies two regimes prior to 1992Q3. We do not find any significant change points for four or more regimes with the choice of threshold variables. Complete results are available upon request. In sum, we are able to identify 3 regimes with threshold variables $t$ and $y_{UK,t-1}$. It should be noted that the results with other lagged dependent variables are not feasible as some estimates of $\beta_i$ are beyond the bound of $-2$ and 0 (see (17)), regardless of they having a smaller RSS or not. One regime is identified by $t \leq$ 1992Q3 and $y_{UK,t-1} \leq -0.304$, the second regime is identified by $t \leq$ 1992Q3 and $y_{UK,t-1} > -0.304$. Regime 1 covers most of periods prior to the end of 1992, whereas Regime 2 includes 1974Q1–1975Q1 (except for 1974Q3), 1976Q3–1977Q3, 1983Q2, and 1983Q4–1987Q1 (except for 1986Q3). The third regime is identified by $t >$ 1992Q3 (see Figure 2). It is clear from Figure 2 that for most of the time prior to 1992Q3, real exchange rates for the UK (as well as other currencies) fall into Regime 1, when countries virtually experienced appreciation. In contrast, Regime 2 is characterized by infrequent depreciation during the 1970s and the mid-1980s. Finally, the results on Regime 3 are consistent with the empirical results in Koedijk, Tims, and van Dijk (2004), and Lopez and Papell (2007).

The nonlinear unit root tests indicate that (see Table 5), under Regime 1, unit root is only marginally rejected or marginally not rejected; under Regime 2, it is not rejected at any sensible size; while under Regime 3, it is

[14] Based on Okui (2009), an AR(4) of (8) is selected.

Table 4: Threshold test[a]

| | | Bootstrapped critical values | | | |
|---|---|---|---|---|---|
| | Test statistic | 10% | 5% | 1% | Threshold variable = $t$ |
| $W_{NT}(2\|1)$ | 13.962 [0.022][b] | 12.442 | 13.072 | 14.541 | Threshold = (1992Q3) |
| $W_{NT}(3\|2)$ | 12.836 [0.308] | 14.286 | 14.944 | 16.634 | Threshold = (1992Q3, 1979Q4) |
| $W_{NT}(4\|3)$ | 12.540 [0.694] | 15.614 | 16.190 | 18.081 | Threshold = (1992Q3, 1979Q4, 1976Q3) |
| | | Bootstrapped critical values (given a threshold at 1992Q3) | | | |
| | Test statistic | 10% | 5% | 1% | Threshold variable = $y_{UK,t-1}$ |
| $W_{NT}(3\|2)$ | 13.418 [0.028] | 12.178 | 12.785 | 14.496 | Threshold = (−0.304) |
| $W_{NT}(4\|3)$ | 11.693 [0.150] | 12.074 | 12.669 | 13.861 | Threshold = (−0.304, − 0.367) |

Notes: [a] Full sample: 1973Q1–1998Q4.

[b] $p$-values are in brackets. The bootstrap uses 1,000 repetitions.

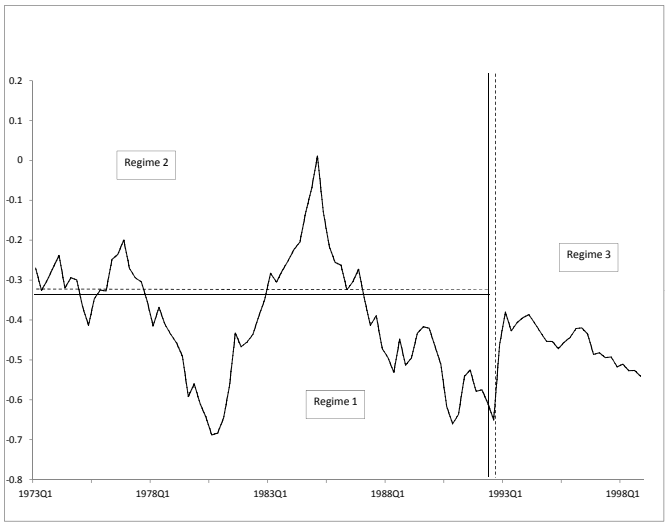



Figure 2: The UK real exchange rate ($y_{UK,t-1}$) vs time ($t$)

Notes: Regime 1: 1973Q1 $\leq t \leq$ 1992Q3 and $y_{UK,t-1} \leq -0.304$; Regime 2: 1973Q1 $\leq t \leq$ 1992Q3 and $y_{UK,t-1} > -0.304$; Regime 3: 1992Q3 $< t \leq$ 1998Q4.

strongly rejected. This suggests that before the establishment of the Maastricht Treaty in 1992, the real exchange rates are not mean-reverting and PPP may not hold well, when the UK currency (as well as that of other countries) is depreciated. That said, there is mild evidence that in the same period of time, the real exchange rates are mean-reverting and PPP holds, when the UK currency (as well as that of other countries) is appreciated. In contrast, after the establishment, the real exchange rates are mean-reverting and PPP holds.

A number of further interpretations are in order. After the collapse of Bretton Woods system, the OECD countries adopted exchange rate target zones in which the exchange rates are fixed but adjustable. According to Garber and Svensson (1995, p. 1866), "An individual central bank in a group of cental banks may peg its currency to a weighted average of the value of the group's currency. Even the major currency central banks frequently intervene in exchange markets to prevent the values of their currencies from moving excessively. Occasionally, they agree to impose informal upper and lower bounds on exchange rates." Thus, we would observe in Regime 2 there

Table 5: Nonlinear panel unit root test: separate regime[a]

| | $t \leq$ 1992Q3 | | | |
|---|---|---|---|---|
| | $y_{UK,t-1} \leq -0.304$ | $y_{UK,t-1} > -0.304$ | 1992Q3 $< t \leq$ 1998Q4 | $t \leq$ 1998Q4 |
| Regime | 1 | 2 | 3 | Full-sample |
| No. of quarters | 52 | 23 | 25 | 100 |
| | −1.772 | −1.358 | −3.559 | −2.229[b] |
| | [0.084][c] | [0.142] | [0.001] | [0.001] |
| Bootstrapped critical values[d] | | | | |
| 10% | −1.731 | −1.463 | −1.993 | −1.488 |
| 5% | −1.887 | −1.636 | −2.155 | −1.584 |
| 1% | −2.181 | −1.968 | −2.465 | −1.762 |

Notes: [a] Full sample: 1973Q1–1998Q4.

[b] The full-sample NCIPS, FNCIPS is the regime average of the cross-sectional averages of the $t$-statistics of the $\hat{\beta}_i^{(k)}$'s.

[c] $p$-values are in brackets. The bootstrap uses 1,000 repetitions.

[d] The bootstrapped critical value does not seem monotonic to the number of quarters, $T$. That may be due to small $T$ such as 23 or 25.

is a tendency to revert to the mean when the currency depreciates too much. This observation is consistent with the view of international goods market arbitrage in the presence of transactions costs. Prior to the Treaty, there is no agreement like the Maastricht of 1992 to require member countries follow the convergence criterion, the exchange rates were likely to fluctuate due to the massive speculative attacks and the weak economic fundamentals (see, for instance, Taylor and Peel (2000)).

We close this section with a discussion on the mean-reversion coefficient for Country $i$ in Regime $k$, $\hat{\beta}_i^{(k)}$, which is reported in Table 6. There is considerable variation in the $\hat{\beta}_i^{(k)}$'s. Compared to Regime 1, the mean-reversion coefficient for the UK is smaller than that in Regime 2. As shown in Figure 2, the depreciation in the sterling-dollar (with a higher $y_{UK,t-1}$) was translated into a high speed of mean-reversion. To put it differently, the mean-reversion coefficient for the UK changes from $-0.013$ to $-0.124$ (though the standard error is larger), the series gets closer to a white noise

Table 6: Mean-reversion coefficient: separate regime and individual country[a]

| | $t \leq$ 1992Q3 | | 1992Q3 $< t$ | |
|---|---|---|---|---|
| | $y_{UK,t-1} \leq -0.304$ | $y_{UK,t-1} > -0.304$ | $\leq$ 1998Q4 | $t \leq$ 1998Q4 |
| Regime | 1 | 2 | 3 | Full-sample |
| No. of quarters | 52 | 23 | 25 | 100 |
| Country | $\hat{\beta}_i^{(1)}$ | $\hat{\beta}_i^{(2)}$ | $\hat{\beta}_i^{(3)}$ | $\hat{\beta}_i$ |
| Australia | −0.066 (0.058)[b] | −0.068 (0.125) | −0.556 (0.120)*** | −0.058 (0.032) |
| Austria | −0.107 (0.069) | −0.284 (0.154) | −0.135 (0.045) | −0.202 (0.049)** |
| Belgium | −0.269 (0.106) | 0.034 (0.114) | −0.595 (0.122)*** | −0.050 (0.031) |
| Canada | −0.033 (0.025) | 0.065 (0.110) | −0.766 (0.159)*** | −0.030 (0.023) |
| Denmark | −0.025 (0.035) | −0.120 (0.065) | −0.168 (0.087) | −0.017 (0.017) |
| Finland | −0.210 (0.105) | −0.388 (0.283) | −0.478 (0.197) | −0.111 (0.069) |
| France | −0.100 (0.051) | −0.244 (0.087) | −0.364 (0.126) | −0.087 (0.031) |
| Germany | −0.087 (0.061) | −0.860 (0.169)*** | −0.790 (0.164)*** | −0.155 (0.046)* |
| Italy | −0.086 (0.036) | −0.292 (0.112) | −0.568 (0.096)*** | −0.079 (0.028) |
| Japan | −0.025 (0.042) | −0.085 (0.059) | −0.493 (0.095)*** | −0.087 (0.035) |
| Netherlands | −0.086 (0.038) | −0.228 (0.100) | −0.616 (0.166)** | −0.042 (0.024) |
| New Zealand | −0.384 (0.092)** | −0.332 (0.187) | −0.727 (0.216)* | −0.113 (0.056) |
| Norway | −0.036 (0.028) | −0.200 (0.262) | −0.578 (0.079)*** | −0.052 (0.026) |
| Spain | −0.114 (0.046) | −0.555 (0.322) | −0.285 (0.076)** | −0.113 (0.042) |
| Sweden | −0.118 (0.058) | −0.073 (0.068) | −0.592 (0.125)*** | −0.058 (0.034) |
| Switzerland | −0.230 (0.061)** | −0.193 (0.089) | −0.413 (0.124)* | −0.077 (0.030) |
| UK | −0.013 (0.058) | −0.124 (0.186) | −0.362 (0.115)* | −0.085 (0.044) |
| Country average | −0.117 | −0.232 | −0.499 | −0.083 |
| Simulated critical values[c] | | | | |
| 10% | −3.046 | −3.042 | −3.043 | −3.050 |
| 5% | −3.502 | −3.495 | −3.497 | −3.518 |
| 1% | −4.534 | −4.514 | −4.524 | −4.558 |

Notes: [a] Full sample: 1973Q1–1998Q4.

[b] The estimated standard errors are in brackets. *, ** and *** denote significance at 10%, 5% and 1% level respectively.

[c] These (finite-sample) simulated critical values for the $t$-statistic $\hat{\beta}_i^{(k)}/\hat{se}(\hat{\beta}_i^{(k)})$ are interpolated from those for the CADF unit root tests (Pesaran (2007), Table I(b)), which assume normal errors.

and the speed of mean-reversion is higher. In fact, this is also the case for other OECD countries except Belgium, Canada, New Zealand, Sweden and Switzerland. For Belgium and Canada, the $\hat{\beta}_i^{(2)}$'s (in Regime 2)

are, unexpectedly, positive (it is statistically insignificant though). For New Zealand, Sweden and Switzerland, the differences in the $\hat{\beta}_i^{(k)}$, $k = 1, 2$'s do not seem to be substantial. Further, all the standard errors in Regime 2 are relatively larger, partially due to fewer observations in that regime. Compared to Regime 2, the mean-reversion coefficient for the UK is in turn smaller than that in Regime 3. The increase in the speed of mean-reversion signifies that the long-run PPP may hold after the Maastricht Treaty was signed. The mean-reversion coefficient for the UK changes from $-0.124$ to $-0.362$, the series gets further closer to a white noise and the speed of mean-reversion is higher. This is also the case for other OECD countries except Austria, Germany and Spain, though the speed is, nevertheless, faster than that in Regime 1 for these three countries. These findings for Regime 3 are consistent with previous studies on the impact of the European integration process. See, for instance, Lopez and Papell (2007) and Christidou and Panagiotidis (2010). And as concluded in Christidou and Panagiotidis (2010, pp. 1120–1121) "The effect of the Maastricht Treaty and the subsequent criteria seem to have a more significant effect as countries made an attempt to reduce inflation in order to participate in the single currency".

As one can see from the $t$-tests in Table 6,[15] prior to the establishment of the Maastricht Treaty, the results do not show much evidence for mean reversion. The unit root null can be rejected at the 5% significance level for New Zealand and Switzerland in Regime 1, and Germany in Regime 2. In the post-Maastricht Treaty period (Regime 3), however, the statistics reveal more evidence for mean reversion, with rejection of the unit root null in 13 out of the 17 countries. In particular, the null is strongly rejected at the 1% level for Australia, Belgium, Canada, Germany, Italy, Japan, Norway, and Sweden. Finally, the $t$-tests for the full-sample reject the unit root null in only 2 out of the 17 countries at the 10% (or smaller) significance level.

Pesaran (2007) considers a CADF test and a CIPS test. His empirical results fail to reject the unit root tests, thus suggesting non-mean reversion. Subsequently, several scholars suggest some potential avenues for further investigation on the empirical validity of PPP.

Lopez and Papell (2007) find strong evidence of convergence to PPP within the Eurozone starting in 1996. Moreover, they provide stronger evidence for the PPP, with the process of convergence starting in 1993 and a

[15] These $t$-tests are different from the CADF tests considered in Pesaran (2007), as the latter do not consider the possible nonlinearity.

rejection of the unit root hypothesis beginning in 1998. In a recent study, Wu, Cheng, and Hou (2011) apply panel unit root tests to country subgroups that have been shaped to feature similar characteristics. Their findings indicate the validity of PPP is conditional on regional groups, in which openness, growth rates, inflation and nominal exchange rate volatility play a pivotal role in the strength of PPP. While our findings are generally consistent with previous empirical findings of stronger evidence for PPP prior to the launch of the euro, we contribute the literature by identifying distinct exchange rate regimes which lead to different strength of support for PPP. Both Papell (2006) and Cerrato et al. (2011) use the CIPS-based tests assuming a stationary common factor across individual units which in turn account for cross-sectional dependence. Both papers also allow for a nonlinear adjustment to the long-rum equilibrium under the alternative hypothesis, and heterogeneous serial correlation processes across the real exchange rates in the test specification. Their specification of nonlinear adjustment is somewhat different from ours though. Our TAR specification in (5) is flexible to allow for shifts not only in slopes, but also in levels and factor loadings. In contrast, Cerrato et al. (2011) only consider a nonlinear adjustment in mean reversion under the alternatives, with deterministic components in means. Further, we consider two types of structural breaks. One is the time index and the other is a global variable. The latter is particularly important for the highly integrated euro area as the dominant economy, such as the UK and Germany, would likely have significant impact on the member exchange rates. Further, the test equations in Cerrato et al. (2011) involve third-order Taylor's expansion and thus the finite-sample performance of the proposed inference procedure is not known. In terms of empirical findings, our TAR procedures identify distinct exchange rate regimes which lead to different strength of support for PPP, whereas the test in Cerrato et al. (2011) rejects the unit root null more frequently than Pesaran (2007) and therefore yields stronger support for the long-run PPP.

### 3.2 Impact of the euro and the 2008 global financial crisis

In this sub-section, the sample period spans from 1973Q1 to 2012Q4. For the sub-period from 1999Q1 to 2012Q4, the dollar exchange rates of the Euro Area countries are calculated by $s_i = s_{euro} + s_j$ where $s_{euro}$ is the log of the euro price of a dollar and $s_j$ is the log of a Eurozone country's currency conversion rate of a euro. The conversion rates, for various Euro Area coun-

tries are available at: (http://www.ecb.europa.eu/euro/intro/html/index.en.html). Again, Figures 3a and 3b depict the logarithm of real exchange rates against the US dollar through 2012Q4, and Table 7 shows the pairwise correlations between real exchange rates.

The threshold test results in Table 8 clearly indicate that 4 regimes are strongly significant at the 5% significance level. These distinct regimes are identified by the time index $t$ and the global variables, $y_{\mathrm{UK},t-1}$. More precisely, apart from the 3 regimes in our our major study.[16] Moreover, there is one more regime, 2000Q4–2012Q4, which is arguably due to the economic downturn and the restructuring of financial markets in the aftermath of the global financial crisis.

With regard to the nonlinear unit root tests on the extended data, the results shown in Table 9 are qualitatively similar to those found in the shorter panel ending in 1998Q4.

## 4 Conclusions

Previous empirical studies on the validity of PPP suffered from lack of degrees of freedom (in, say, purely time-series models), ignorance of possible structural change, ignorance of possible asymmetric adjustment, and/or ignorance of cross-sectional dependence or heterogeneity (in, say, panel-data models). Integrating Pesaran (2007) test of linear panel unit root model with cross-sectional dependence or heterogeneity, and Caner and Hansen (2001) TAR unit root test, this paper investigates whether the real exchange rates of 17 OECD countries during the post-Bretton Woods period exhibit mean-reversion.

Our results suggest that using time index and the lagged UK real exchange rate as thresholds, three distinct regimes are identified. We identify a case of influential currency arrangements in effect as significant threshold in the dynamics of PPP exchange rates during the recent floating exchange rate period. More precisely, we document the significant effect of currency realignments in the post-Bretton Woods period and the Maastricht Treaty in 1992 in the analysis of the PPP tenet across OECD real exchange rates.

[16] It should be noted that due to some sampling errors or otherwise, with this extended data set, the 3 regimes are a bit different. The threshold of $y_{\mathrm{UK},t-1}$ is now 0.0701. And the other two regimes are 1973Q1–1993Q2 (rather than 1973Q1–1992Q3) and 1993Q3–2000Q3 (rather than 1992Q4–1998Q4).

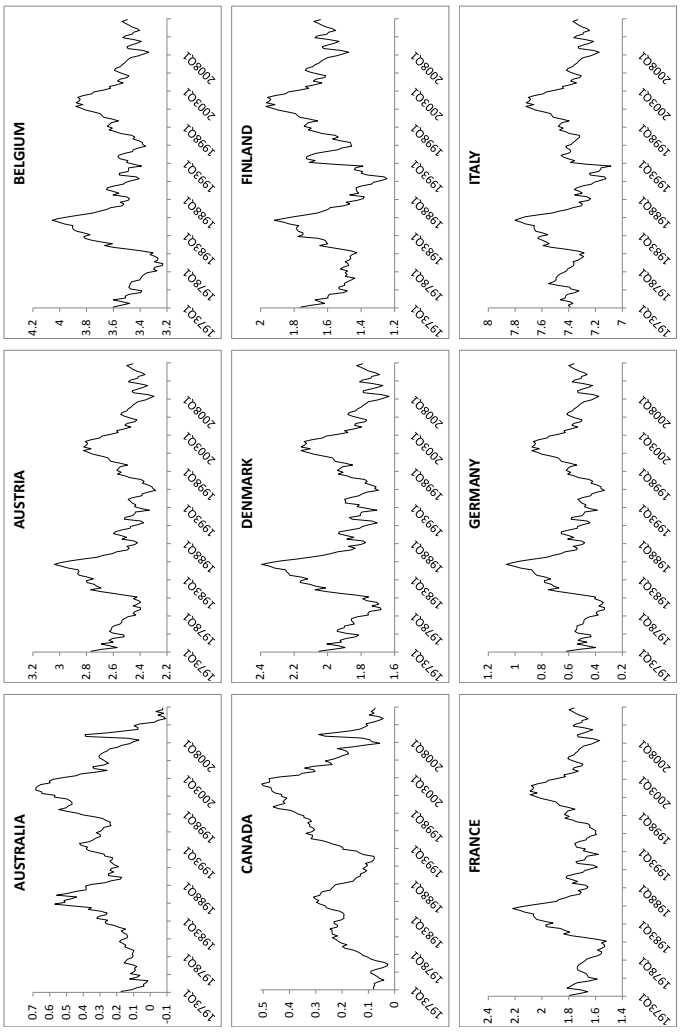



Figure 3a: Real exchange rates relative to the US dollar 1973Q1–2012Q4

Notes: Each plot depicts the real exchange rate of one of the 17 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, and the UK.

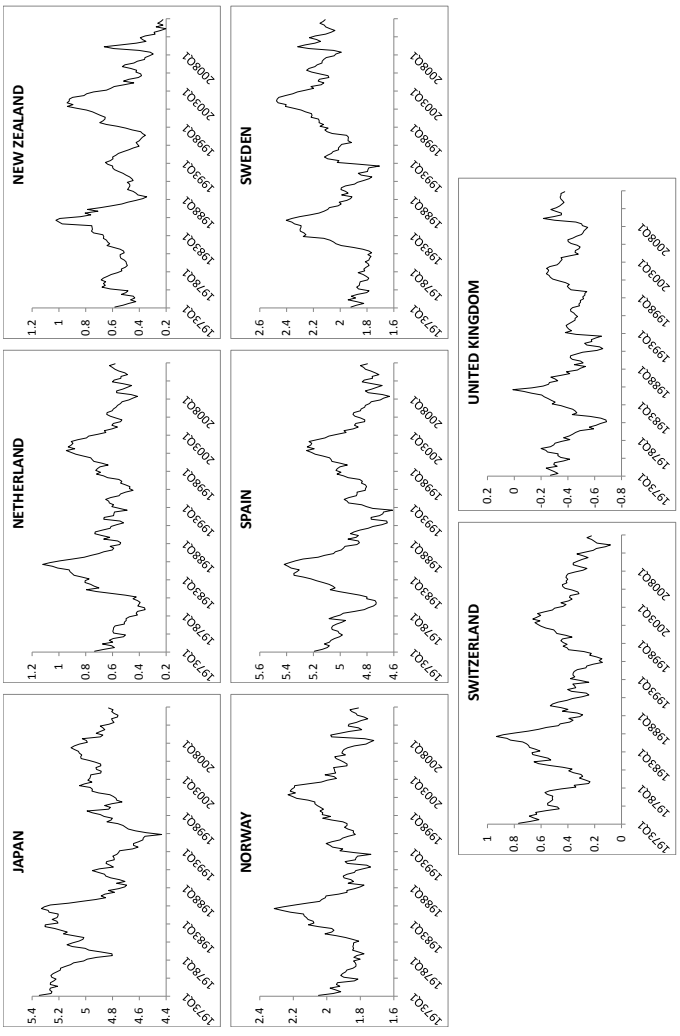



Figure 3b: Real exchange rates relative to the US dollar 1973Q1–2012Q4

Notes: Each plot depicts the real exchange rate of one of the 17 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, and the UK.

Table 7: Correlations of OECD real exchange rates

| | AUS | AUT | BEL | CAN | DEN | FIN | FRA | GER | ITA | JAP | NET | NOR | NAZ | SPA | SWE | SWI | UKI |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| AUS | 1.000 | | | | | | | | | | | | | | | | |
| AUT | 0.443 | 1.000 | | | | | | | | | | | | | | | |
| BEL | 0.631 | 0.893 | 1.000 | | | | | | | | | | | | | | |
| CAN | 0.857 | 0.316 | 0.496 | 1.000 | | | | | | | | | | | | | |
| DEN | 0.515 | 0.981 | 0.930 | 0.384 | 1.000 | | | | | | | | | | | | |
| FIN | 0.571 | 0.681 | 0.756 | 0.627 | 0.695 | 1.000 | | | | | | | | | | | |
| FRA | 0.601 | 0.904 | 0.971 | 0.492 | 0.923 | 0.812 | 1.000 | | | | | | | | | | |
| GER | 0.571 | 0.925 | 0.974 | 0.442 | 0.936 | 0.733 | 0.971 | 1.000 | | | | | | | | | |
| ITA | 0.514 | 0.878 | 0.793 | 0.522 | 0.887 | 0.786 | 0.859 | 0.818 | 1.000 | | | | | | | | |
| JAP | −0.153 | 0.606 | 0.309 | −0.252 | 0.511 | 0.218 | 0.333 | 0.381 | 0.419 | 1.000 | | | | | | | |
| NET | 0.615 | 0.918 | 0.987 | 0.473 | 0.949 | 0.737 | 0.964 | 0.975 | 0.809 | 0.325 | 1.000 | | | | | | |
| NOR | 0.705 | 0.877 | 0.907 | 0.640 | 0.907 | 0.853 | 0.917 | 0.885 | 0.883 | 0.316 | 0.915 | 1.000 | | | | | |
| NAZ | 0.718 | 0.795 | 0.716 | 0.528 | 0.808 | 0.582 | 0.738 | 0.741 | 0.786 | 0.366 | 0.751 | 0.822 | 1.000 | | | | |
| SPA | 0.470 | 0.927 | 0.811 | 0.388 | 0.933 | 0.699 | 0.845 | 0.815 | 0.922 | 0.520 | 0.834 | 0.869 | 0.774 | 1.000 | | | |
| SWE | 0.604 | 0.556 | 0.778 | 0.617 | 0.594 | 0.863 | 0.806 | 0.754 | 0.634 | 0.047 | 0.718 | 0.757 | 0.459 | 0.528 | 1.000 | | |
| SWI | 0.364 | 0.934 | 0.764 | 0.181 | 0.903 | 0.550 | 0.768 | 0.786 | 0.777 | 0.729 | 0.799 | 0.777 | 0.759 | 0.879 | 0.376 | 1.000 | |
| UKI | 0.191 | 0.638 | 0.567 | 0.081 | 0.620 | 0.597 | 0.661 | 0.591 | 0.669 | 0.382 | 0.571 | 0.608 | 0.509 | 0.706 | 0.548 | 0.574 | 1.000 |

Notes: This table shows correlations of quarterly real exchange rates against the US dollar measured in logarithms for 17 OECD countries, spanning the period 1973Q1 to 2012Q4. Each column represents the real exchange rate of each individual country, ordered as follows: Australia (AUS), Austria (AUT), Belgium (BEL), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (GER), Italy (ITA), Japan (JAP), the Netherlands (NET), Norway (NOR), New Zealand (NZA), Spain (SPA), Sweden (SWE), Switzerland (SWI), and the UK (UKI).

Table 8: Threshold test[a]

| | Bootstrapped critical values | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | Test statistic | 10% | 5% | 1% | Threshold | Cut Point | SubSample (%) | | | | |
| $W_{NT}(2\|1)$ | 12.294 [0.011] | 10.577 | 11.061 | 12.323 | Time | 2000Q4 | 0.1282 | 0.5769 | 0.2949 | | |
| $W_{NT}(3\|2)$ | 16.580 [0.000] | 10.149 | 10.559 | 11.579 | UK | 0.0701 | 0.1282 | 0.3974 | 0.1795 | 0.2949 | |
| $W_{NT}(4\|3)$ | 11.904 [0.023] | 10.766 | 11.313 | 12.519 | Time | 1993Q2 | 0.1282 | 0.2372 | 0.1603 | 0.1795 | 0.2949 |
| $W_{NT}(5\|4)$ | 9.216 [0.138] | 9.435 | 9.834 | 11.107 | UK | | | | | | |
| $W_{NT}(5\|4)$ | 7.477 [0.680] | 10.472 | 10.942 | 12.251 | Time | | | | | | |

Notes: [a] Full sample: 1973Q1–2012Q4.

[b] $p$-values are in brackets. The bootstrap uses 1,000 repetitions.

Table 9: Nonlinear panel unit root test: separate regime[a]

| Regime | 1 | 2 | 3 | 4 | Full-sample |
|---|---|---|---|---|---|
| No. of quarters | 42 | 30 | 33 | 51 | 156 |
| | −0.919 | −2.726 | −1.715 | −1.953 | −1.778 |
| | [0.697] | [0.001] | [0.100] | [0.158] | [0.001] |
| | Bootstrapped critical values[d] | | | | |
| 10% | −1.654 | −1.883 | −1.714 | −2.061 | −1.413 |
| 5% | −1.801 | −2.037 | −1.857 | −2.197 | −1.475 |
| 1% | −2.096 | −2.338 | −2.186 | −2.482 | −1.596 |

Notes: [a] Full sample: 1973Q1–2012Q4.

[b] The full-sample NCIPS, FNCIPS is the regime average of the cross-sectional averages of the $t$-statistics of the $\hat{\beta}_i^{(k)}$'s.

[c] $p$-values are in brackets. The bootstrap uses 1,000 repetitions.

[d] The bootstrapped critical value does not seem monotonic to the number of quarters, $T$. That may be due to small $T$ such as 23 or 25.

In the pre-Treaty period, there is some evidence for long-run PPP when the sterling-dollar appreciates, while the evidence is weak when the sterling-dollar depreciates. From an economic perspective, the real and financial integration in Europe seemed to lend support to the PPP hypothesis. Our results based on the nonlinear panel unit root tests echo the findings of Choi (2004) that the strength of the evidence of PPP depends crucially on the presence of nonlinear dynamics in the real exchange rates, or a potential mix of stationary and non-stationary series in the panel.

Although this study goes directly two-step beyond that of Li and Ling (2012) by the panel-data and nonstationary setting, it should be noted that this literature on the multiple-regime TAR models have suffered from the lack of a large sample theory, notably the limiting distributions of the estimated thresholds. The inference for thresholds in the current paper is based on a residual-based bootstrap. Establishing a large sample theory is a job to be called for in future studies.

## Appendix A: Monte Carlo Simulation

In order to assess the performance of estimating the panel TAR model, the

threshold test as well as the nonlinear panel unit root test, we carry out a number of Monte Carlo simulations. The results are reported in Tables 10 through 13. We consider the following 2-regime dynamic panel data model (see (5)):

$$\Delta y_{i,t} = \left(\beta^{(1)} y_{i,t-1} + \gamma_i^{(1)} f_t + u_{i,t}^{(1)}\right) \mathbf{1}(t \leq T/2) + \left(\beta^{(2)} y_{i,t-1} + \gamma_i^{(2)} f_t + u_{i,t}^{(2)}\right) \mathbf{1}(t > T/2), \tag{A.1}$$

$i = 1, \cdots, N, t = 1, \cdots, T$. In other words, we set $q_{t-1}^{(1)} = q_{t-1}^{(2)} = t$ with $(\underline{\lambda}^{(1)}, \overline{\lambda}^{(1)})' = (-\infty, T/2)'$ and $(\underline{\lambda}^{(2)}, \overline{\lambda}^{(2)})' = (T/2, \infty)'$ (the structural change points). $f_t \overset{i.i.d.}{\sim} N(0,1)$, $\gamma_i^{(k)} \overset{i.i.d.}{\sim} N(2,1)$, and $u_{i,t}^{(k)} \overset{i.i.d.}{\sim} N(0,1)$ for $k = 1, 2$.

Under the null of unit root, $H_{0,II}: \beta^{(1)} = \beta^{(2)} = 0$. Three classes of alternatives are considered: (i) under $H_{A,II}^{(i)}$, $\beta^{(1)} \in \{-0.1, -0.2, -0.3, -0.4\}$ and $\beta^{(2)} = 0$, (ii) under $H_{A,II}^{(ii)}$, $\beta^{(1)} = \beta^{(2)}$, while $\beta^{(2)} = \{-0.1, -0.2, -0.3, -0.4\}$, and (iii) under $H_{A,II}^{(iii)}$, $\beta^{(1)} = 0$ and $\beta^{(2)} \in \{-0.1, -0.2, -0.3, -0.4\}$. For each $T$, $N \in \{25, 50, 100, 200, 400\}$, the number of repetitions, $M = 1,000$.

Denote $\beta_t = \beta^{(1)}\mathbf{1}(t \leq T/2) + \beta^{(2)}\mathbf{1}(t > T/2)$. For each repetition $m$, denote $\hat{\beta}_{t,m} = \hat{\beta}_m^{(1)}\mathbf{1}(t \leq \hat{\lambda}_m) + \hat{\beta}_m^{(2)}\mathbf{1}(t > \hat{\lambda}_m)$. The bias and the mean squared error (MSE) are defined as:

$$\text{Bias} = \frac{1}{M}\sum_{m=1}^{M}\frac{1}{T}\sum_{t=1}^{T}\left(\hat{\beta}_{t,m} - \beta_t\right), \tag{A.2}$$

$$\text{MSE} = \frac{1}{M}\sum_{m=1}^{M}\frac{1}{T}\sum_{t=1}^{T}\left(\hat{\beta}_{t,m} - \beta_t\right)^2. \tag{A.3}$$

The bias and the MSE are summarized in Tables 10 and 11 respectively. One can see that the MSE declines over $N$ and $T$. Also, when $\beta^{(1)}$ is closer to zero, the MSE becomes larger. In addition, we also find that the bias tends to be a constant when $T$ is fixed.[17,18]

[17] This might possibly due to the fixed-effects transformation in each regime. See also Harris and Tzavalis (1999).

[18] We also calculate the MSE for $\hat{\beta}_{t,m}^{(1)}$ and $\hat{\beta}_{t,m}^{(2)}$ for regimes 1 and 2. The MSE in each regime also declines with $N$ and $T$. The detailed results are available upon request.

Table 10: MSE under $H_{0,II}$, $H_{A,II}^{(i)}$, $H_{A,II}^{(ii)}$ and $H_{A,II}^{(iii)}$

| | N=25 | 50 | 100 | 200 | 400 |
|---|---|---|---|---|---|
| | $(\beta^{(1)}, \beta^{(2)}) = (0, 0)$ | | | | |
| T=25 | 0.153 | 0.150 | 0.151 | 0.149 | 0.150 |
| 50 | 0.056 | 0.056 | 0.054 | 0.055 | 0.055 |
| 100 | 0.019 | 0.018 | 0.018 | 0.018 | 0.018 |
| 200 | 0.005 | 0.005 | 0.005 | 0.005 | 0.005 |
| 400 | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.1, 0)$ | | | | |
| T=25 | 0.135 | 0.132 | 0.132 | 0.130 | 0.130 |
| 50 | 0.046 | 0.043 | 0.042 | 0.042 | 0.042 |
| 100 | 0.013 | 0.011 | 0.010 | 0.010 | 0.010 |
| 200 | 0.002 | 0.002 | 0.002 | 0.002 | 0.002 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.2, 0)$ | | | | |
| T=25 | 0.121 | 0.120 | 0.118 | 0.116 | 0.116 |
| 50 | 0.037 | 0.035 | 0.033 | 0.032 | 0.032 |
| 100 | 0.009 | 0.008 | 0.007 | 0.007 | 0.007 |
| 200 | 0.002 | 0.002 | 0.002 | 0.002 | 0.002 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.3, 0)$ | | | | |
| T=25 | 0.111 | 0.109 | 0.107 | 0.105 | 0.105 |
| 50 | 0.032 | 0.029 | 0.027 | 0.026 | 0.027 |
| 100 | 0.007 | 0.007 | 0.006 | 0.006 | 0.006 |
| 200 | 0.002 | 0.002 | 0.002 | 0.001 | 0.001 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.4, 0)$ | | | | |
| T=25 | 0.102 | 0.101 | 0.099 | 0.096 | 0.096 |
| 50 | 0.028 | 0.026 | 0.025 | 0.023 | 0.024 |
| 100 | 0.006 | 0.006 | 0.006 | 0.006 | 0.006 |
| 200 | 0.002 | 0.002 | 0.001 | 0.001 | 0.001 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |

| | N=25 | 50 | 100 | 200 | 400 |
|---|---|---|---|---|---|
| | $(\beta^{(1)} = \beta^{(2)}) = (-.1, -.1)$ | | | | |
| T=25 | 0.120 | 0.116 | 0.115 | 0.112 | 0.113 |
| 50 | 0.042 | 0.041 | 0.040 | 0.041 | 0.042 |
| 100 | 0.015 | 0.016 | 0.016 | 0.017 | 0.018 |
| 200 | 0.004 | 0.004 | 0.004 | 0.004 | 0.005 |
| 400 | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| | $(\beta^{(1)} = \beta^{(2)}) = (-.2, -.2)$ | | | | |
| T=25 | 0.095 | 0.090 | 0.090 | 0.087 | 0.088 |
| 50 | 0.034 | 0.033 | 0.032 | 0.033 | 0.034 |
| 100 | 0.013 | 0.014 | 0.014 | 0.015 | 0.016 |
| 200 | 0.003 | 0.003 | 0.003 | 0.003 | 0.003 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)} = \beta^{(2)}) = (-.3, -.3)$ | | | | |
| T=25 | 0.075 | 0.071 | 0.070 | 0.068 | 0.069 |
| 50 | 0.027 | 0.026 | 0.025 | 0.026 | 0.028 |
| 100 | 0.010 | 0.011 | 0.011 | 0.012 | 0.013 |
| 200 | 0.002 | 0.002 | 0.002 | 0.002 | 0.003 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)} = \beta^{(2)}) = (-.4, -.4)$ | | | | |
| T=25 | 0.059 | 0.056 | 0.055 | 0.054 | 0.055 |
| 50 | 0.022 | 0.020 | 0.020 | 0.020 | 0.022 |
| 100 | 0.008 | 0.009 | 0.009 | 0.009 | 0.010 |
| 200 | 0.002 | 0.002 | 0.002 | 0.002 | 0.002 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |

| | N=25 | 50 | 100 | 200 | 400 |
|---|---|---|---|---|---|
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.1)$ | | | | |
| T=25 | 0.114 | 0.111 | 0.110 | 0.108 | 0.108 |
| 50 | 0.033 | 0.030 | 0.028 | 0.028 | 0.027 |
| 100 | 0.007 | 0.006 | 0.006 | 0.006 | 0.006 |
| 200 | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.2)$ | | | | |
| T=25 | 0.090 | 0.087 | 0.084 | 0.083 | 0.083 |
| 50 | 0.024 | 0.022 | 0.021 | 0.021 | 0.020 |
| 100 | 0.006 | 0.005 | 0.005 | 0.005 | 0.005 |
| 200 | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.3)$ | | | | |
| T=25 | 0.078 | 0.077 | 0.075 | 0.075 | 0.075 |
| 50 | 0.021 | 0.020 | 0.019 | 0.019 | 0.019 |
| 100 | 0.005 | 0.005 | 0.005 | 0.005 | 0.005 |
| 200 | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.4)$ | | | | |
| T=25 | 0.073 | 0.072 | 0.070 | 0.070 | 0.070 |
| 50 | 0.020 | 0.019 | 0.019 | 0.019 | 0.018 |
| 100 | 0.005 | 0.005 | 0.005 | 0.005 | 0.005 |
| 200 | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| 400 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |

Table 11: Bias under $H_{0,II}$, $H_{A,II}^{(i)}$, $H_{A,II}^{(ii)}$ and $H_{A,II}^{(iii)}$

| | N=25 | 50 | 100 | 200 | 400 |
|---|---|---|---|---|---|
| | $(\beta^{(1)}, \beta^{(2)}) = (0, 0)$ | | | | |
| T=25 | −0.370 | −0.371 | −0.372 | −0.370 | −0.372 |
| 50 | −0.211 | −0.211 | −0.207 | −0.208 | −0.209 |
| 100 | −0.111 | −0.110 | −0.109 | −0.109 | −0.109 |
| 200 | −0.056 | −0.056 | −0.055 | −0.055 | −0.055 |
| 400 | −0.023 | −0.023 | −0.022 | −0.022 | −0.022 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.1, 0)$ | | | | |
| T=25 | −0.347 | −0.346 | −0.346 | −0.344 | −0.346 |
| 50 | −0.186 | −0.181 | −0.180 | −0.180 | −0.181 |
| 100 | −0.082 | −0.081 | −0.079 | −0.080 | −0.080 |
| 200 | −0.036 | −0.036 | −0.036 | −0.036 | −0.036 |
| 400 | −0.015 | −0.015 | −0.015 | −0.015 | −0.015 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.2, 0)$ | | | | |
| T=25 | −0.326 | −0.327 | −0.325 | −0.323 | −0.324 |
| 50 | −0.163 | −0.158 | −0.157 | −0.154 | −0.156 |
| 100 | −0.070 | −0.072 | −0.070 | −0.071 | −0.071 |
| 200 | −0.036 | −0.036 | −0.036 | −0.036 | −0.036 |
| 400 | −0.015 | −0.015 | −0.015 | −0.015 | −0.014 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.3, 0)$ | | | | |
| T=25 | −0.307 | −0.308 | −0.305 | −0.305 | −0.305 |
| 50 | −0.146 | −0.143 | −0.142 | −0.141 | −0.142 |
| 100 | −0.068 | −0.069 | −0.068 | −0.069 | −0.069 |
| 200 | −0.035 | −0.035 | −0.035 | −0.035 | −0.035 |
| 400 | −0.014 | −0.014 | −0.014 | −0.014 | −0.014 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.4, 0)$ | | | | |
| T=25 | −0.291 | −0.291 | −0.288 | −0.287 | −0.287 |
| 50 | −0.136 | −0.134 | −0.135 | −0.134 | −0.135 |
| 100 | −0.067 | −0.068 | −0.067 | −0.067 | −0.068 |
| 200 | −0.034 | −0.034 | −0.034 | −0.034 | −0.034 |
| 400 | −0.014 | −0.014 | −0.014 | −0.014 | −0.014 |

| | N=25 | 50 | 100 | 200 | 400 |
|---|---|---|---|---|---|
| | $(\beta^{(1)} = \beta^{(2)}) = (-.1, -.1)$ | | | | |
| T=25 | −0.326 | −0.323 | −0.323 | −0.320 | −0.321 |
| 50 | −0.174 | −0.169 | −0.168 | −0.168 | −0.170 |
| 100 | −0.083 | −0.082 | −0.081 | −0.082 | −0.082 |
| 200 | −0.039 | −0.038 | −0.038 | −0.037 | −0.038 |
| 400 | −0.014 | −0.014 | −0.013 | −0.013 | −0.013 |
| | $(\beta^{(1)} = \beta^{(2)}) = (-.2, -.2)$ | | | | |
| T=25 | −0.287 | −0.284 | −0.284 | −0.280 | −0.282 |
| 50 | −0.149 | −0.144 | −0.143 | −0.143 | −0.145 |
| 100 | −0.071 | −0.070 | −0.070 | −0.070 | −0.071 |
| 200 | −0.033 | −0.032 | −0.032 | −0.032 | −0.032 |
| 400 | −0.013 | −0.012 | −0.012 | −0.012 | −0.011 |
| | $(\beta^{(1)} = \beta^{(2)}) = (-.3, -.3)$ | | | | |
| T=25 | −0.252 | −0.250 | −0.249 | −0.247 | −0.248 |
| 50 | −0.130 | −0.126 | −0.125 | −0.125 | −0.127 |
| 100 | −0.063 | −0.062 | −0.061 | −0.062 | −0.062 |
| 200 | −0.030 | −0.029 | −0.028 | −0.028 | −0.028 |
| 400 | −0.011 | −0.011 | −0.011 | −0.011 | −0.010 |
| | $(\beta^{(1)} = \beta^{(2)}) = (-.4, -.4)$ | | | | |
| T=25 | −0.223 | −0.220 | −0.220 | −0.218 | −0.220 |
| 50 | −0.115 | −0.110 | −0.110 | −0.110 | −0.112 |
| 100 | −0.056 | −0.055 | −0.054 | −0.055 | −0.055 |
| 200 | −0.027 | −0.026 | −0.025 | −0.025 | −0.025 |
| 400 | −0.010 | −0.010 | −0.010 | −0.010 | −0.009 |

| | N=25 | 50 | 100 | 200 | 400 |
|---|---|---|---|---|---|
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.1)$ | | | | |
| T=25 | −0.312 | −0.312 | −0.311 | −0.308 | −0.308 |
| 50 | −0.159 | −0.153 | −0.150 | −0.148 | −0.147 |
| 100 | −0.071 | −0.068 | −0.067 | −0.067 | −0.067 |
| 200 | −0.033 | −0.033 | −0.032 | −0.033 | −0.032 |
| 400 | −0.013 | −0.013 | −0.013 | −0.013 | −0.013 |
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.2)$ | | | | |
| T=25 | −0.268 | −0.267 | −0.263 | −0.262 | −0.261 |
| 50 | −0.133 | −0.129 | −0.127 | −0.127 | −0.126 |
| 100 | −0.063 | −0.063 | −0.062 | −0.062 | −0.062 |
| 200 | −0.032 | −0.031 | −0.031 | −0.031 | −0.031 |
| 400 | −0.013 | −0.013 | −0.013 | −0.013 | −0.012 |
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.3)$ | | | | |
| T=25 | −0.244 | −0.244 | −0.242 | −0.242 | −0.242 |
| 50 | −0.125 | −0.121 | −0.120 | −0.120 | −0.119 |
| 100 | −0.061 | −0.061 | −0.060 | −0.060 | −0.060 |
| 200 | −0.031 | −0.030 | −0.030 | −0.030 | −0.030 |
| 400 | −0.012 | −0.012 | −0.012 | −0.012 | −0.012 |
| | $(\beta^{(1)}, \beta^{(2)}) = (0, -.4)$ | | | | |
| T=25 | −0.230 | −0.231 | −0.230 | −0.230 | −0.229 |
| 50 | −0.120 | −0.116 | −0.116 | −0.115 | −0.115 |
| 100 | −0.059 | −0.059 | −0.058 | −0.058 | −0.058 |
| 200 | −0.030 | −0.030 | −0.029 | −0.030 | −0.029 |
| 400 | −0.012 | −0.012 | −0.012 | −0.012 | −0.012 |

Table 12: Empirical power of linearity test

| | Common structural change, $q_{t-1}^{(1)} = q_{t-1}^{(2)} = t$ | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | N=25 | 50 | 100 | 200 | 400 | 25 | 50 | 100 | 200 | 400 |
| | $\beta^{(1)} = -0.1, \beta^{(2)} = 0$ | | | | | $\beta^{(1)} = 0, \beta^{(2)} = -0.1$ | | | | |
| T=25 | 0.058 | 0.054 | 0.057 | 0.055 | 0.061 | 0.228 | 0.268 | 0.315 | 0.310 | 0.349 |
| 50 | 0.080 | 0.065 | 0.060 | 0.074 | 0.077 | 0.401 | 0.478 | 0.530 | 0.613 | 0.615 |
| 100 | 0.160 | 0.162 | 0.217 | 0.239 | 0.245 | 0.683 | 0.828 | 0.876 | 0.906 | 0.909 |
| 200 | 0.607 | 0.734 | 0.782 | 0.876 | 0.867 | 0.979 | 0.996 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| | $\beta^{(1)} = -0.2, \beta^{(2)} = 0$ | | | | | $\beta^{(1)} = 0, \beta^{(2)} = -0.2$ | | | | |
| T=25 | 0.079 | 0.066 | 0.069 | 0.078 | 0.085 | 0.528 | 0.628 | 0.685 | 0.692 | 0.718 |
| 50 | 0.189 | 0.190 | 0.186 | 0.236 | 0.252 | 0.840 | 0.873 | 0.922 | 0.948 | 0.941 |
| 100 | 0.620 | 0.736 | 0.845 | 0.832 | 0.861 | 0.984 | 0.999 | 1.000 | 1.000 | 1.000 |
| 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| | $\beta^{(1)} = -0.3, \beta^{(2)} = 0$ | | | | | $\beta^{(1)} = 0, \beta^{(2)} = -0.3$ | | | | |
| T=25 | 0.113 | 0.106 | 0.118 | 0.127 | 0.136 | 0.732 | 0.812 | 0.860 | 0.873 | 0.878 |
| 50 | 0.400 | 0.472 | 0.483 | 0.594 | 0.611 | 0.965 | 0.988 | 0.997 | 0.999 | 1.000 |
| 100 | 0.959 | 0.988 | 0.999 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| | $\beta^{(1)} = -0.4, \beta^{(2)} = 0$ | | | | | $\beta^{(1)} = 0, \beta^{(2)} = -0.4$ | | | | |
| T=25 | 0.170 | 0.171 | 0.201 | 0.229 | 0.247 | 0.868 | 0.908 | 0.946 | 0.963 | 0.964 |
| 50 | 0.668 | 0.765 | 0.792 | 0.849 | 0.845 | 0.997 | 1.000 | 1.000 | 1.000 | 1.000 |
| 100 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |

Next we summarize the power of the threshold test in Table 12. While the power by and large increases with $N$ and $T$, it increases when either $\beta^{(1)}$ or $\beta^{(2)}$ moves further away from zero. For instance, when $\beta^{(1)} = -0.1$ (and $\beta^{(2)} = 0$), the power is only 0.160 when $N = 25$ and $T = 100$. When $\beta^{(1)} = -0.4$, it is remarkably 1.00 for the same $N$ and $T$. Our results suggest that the threshold test in (16) still performs well after taking the cross-sectional dependence into account.

Finally, we summarize the critical value and the power of the nonlinear panel unit root test in (18) in Table 13. Similar to Table 12, as expected, the power increases with $N$ and $T$, and it is more powerful when both regimes are mean-reverting.

Table 13: Critical values and empirical power of the nonlinear panel unit root test against $H_{A,II}^{(i)}$, $H_{A,II}^{(ii)}$ and $H_{A,II}^{(iii)}$

| | | | | | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | | N=25 | 50 | 100 | 200 | 400 | | | | | |
| | | | | | | | $(\beta^{(1)}, \beta^{(2)}) = (0, 0)$ | | | | | | | | | |
| | | | | | | | T=25 | −2.140 | −2.074 | −2.034 | −2.011 | −1.992 | | | | |
| | | | | | | | 50 | −2.195 | −2.136 | −2.071 | −2.054 | −2.039 | | | | |
| | | | | | | | 100 | −2.201 | −2.138 | −2.097 | −2.082 | −2.057 | | | | |
| | | | | | | | 200 | −2.211 | −2.139 | −2.088 | −2.071 | −2.058 | | | | |
| | | | | | | | 400 | −2.200 | −2.153 | −2.102 | −2.086 | −2.066 | | | | |
| | | $(\beta^{(1)}, \beta^{(2)}) = (-.1, 0)$ | | | | | | $(\beta^{(1)} = \beta^{(2)}) = (-.1, -.1)$ | | | | | | $(\beta^{(1)}, \beta^{(2)}) = (0, -.1)$ | | |
| T=25 | 0.076 | 0.082 | 0.071 | 0.082 | 0.090 | | T=25 | 0.113 | 0.137 | 0.140 | 0.139 | 0.158 | T=25 | 0.050 | 0.054 | 0.056 | 0.062 | 0.068 |
| 50 | 0.096 | 0.093 | 0.142 | 0.149 | 0.170 | | 50 | 0.330 | 0.378 | 0.533 | 0.580 | 0.686 | 50 | 0.090 | 0.080 | 0.123 | 0.126 | 0.133 |
| 100 | 0.239 | 0.339 | 0.374 | 0.421 | 0.503 | | 100 | 0.907 | 0.985 | 0.997 | 1.000 | 1.000 | 100 | 0.359 | 0.509 | 0.612 | 0.681 | 0.761 |
| 200 | 0.747 | 0.927 | 0.987 | 0.999 | 1.000 | | 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 200 | 0.965 | 0.997 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| | N=25 | 50 | 100 | 200 | 400 | | | N=25 | 50 | 100 | 200 | 400 | | N=25 | 50 | 100 | 200 | 400 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.2, 0)$ | | | | | | | $(\beta^{(1)} = \beta^{(2)}) = (-.2, -.2)$ | | | | | | $(\beta^{(1)}, \beta^{(2)}) = (0, -.2)$ | | | | |
| T=25 | 0.107 | 0.143 | 0.132 | 0.151 | 0.161 | | T=25 | 0.254 | 0.315 | 0.383 | 0.423 | 0.477 | T=25 | 0.103 | 0.120 | 0.122 | 0.124 | 0.155 |
| 50 | 0.248 | 0.274 | 0.416 | 0.434 | 0.498 | | 50 | 0.861 | 0.961 | 0.994 | 1.000 | 1.000 | 50 | 0.421 | 0.539 | 0.725 | 0.781 | 0.849 |
| 100 | 0.792 | 0.917 | 0.978 | 0.996 | 1.000 | | 100 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 100 | 0.970 | 0.994 | 1.000 | 1.000 | 1.000 |
| 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.3, 0)$ | | | | | | | $(\beta^{(1)} = \beta^{(2)}) = (-.3, -.3)$ | | | | | | $(\beta^{(1)}, \beta^{(2)}) = (0, -.3)$ | | | | |
| T=25 | 0.179 | 0.244 | 0.252 | 0.266 | 0.319 | | T=25 | 0.474 | 0.655 | 0.759 | 0.829 | 0.881 | T=25 | 0.260 | 0.355 | 0.443 | 0.497 | 0.569 |
| 50 | 0.513 | 0.661 | 0.799 | 0.849 | 0.903 | | 50 | 0.996 | 1.000 | 1.000 | 1.000 | 1.000 | 50 | 0.843 | 0.942 | 0.994 | 0.995 | 1.000 |
| 100 | 0.996 | 1.000 | 1.000 | 1.000 | 1.000 | | 100 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 100 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| | $(\beta^{(1)}, \beta^{(2)}) = (-.4, 0)$ | | | | | | | $(\beta^{(1)} = \beta^{(2)}) = (-.4, -.4)$ | | | | | | $(\beta^{(1)}, \beta^{(2)}) = (0, -.4)$ | | | | |
| T=25 | 0.282 | 0.370 | 0.420 | 0.459 | 0.526 | | T=25 | 0.758 | 0.920 | 0.979 | 0.998 | 0.999 | T=25 | 0.526 | 0.699 | 0.768 | 0.830 | 0.876 |
| 50 | 0.809 | 0.926 | 0.978 | 0.997 | 1.000 | | 50 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 50 | 0.977 | 0.998 | 1.000 | 1.000 | 1.000 |
| 100 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 100 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 100 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 200 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |
| 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 | 400 | 1.000 | 1.000 | 1.000 | 1.000 | 1.000 |

## Appendix B: A Residual-Based Bootstrap Procedure

We first exposit a residual-based bootstrap procedure for the threshold test in which the null and alternative hypotheses (see (15)) are:

$$H_{0,I}: K \text{ regimes}, \quad H_{A,I}: K+1 \text{ regimes}. \tag{B.1}$$

The bootstrap procedure is divided into the following steps:

1. Step for error. Consider the threshold variables $\{q_{t-1}^{(k)}, k = 1, \cdots, K+1\}$. Obtain, under $(K+1)-$regimes, the estimated thresholds $\hat{\lambda}_{K+1,A} = (\hat{\lambda}_A^{(1)'}, \cdots, \hat{\lambda}_A^{(K+1)'})'$ and the estimated $\hat{\phi}_{iA}^{(k)} = (\hat{a}_{iA}^{(k)}, \hat{\beta}_{iA}^{(k)}, \hat{\delta}_{1A}^{(k)}, \cdots, \hat{\delta}_{\tau A}^{(k)}, \hat{b}_{iA}^{(k)}, \hat{c}_{iA,1}^{(k)}, \cdots, \hat{c}_{iA,\tau}^{(k)})'$, $i = 1, \cdots, N$; $k = 1, \cdots, K+1$. See (14). By (11) and footnote 9, we calculate the residual for $i = 1, \cdots, N$; $t = 1, \cdots, T$:[19]

$$\hat{\epsilon}_{it} = \Delta y_{i,t} - \sum_{k=1}^{K+1} \sum_{\hat{\lambda}_A^{(k)} < q_{t-1}^{(k)} \leq \hat{\lambda}_A^{(k)}} \left( \hat{a}_{iA}^{(k)} + \hat{\beta}_{iA}^{(k)} y_{i,t-1} + \sum_{j=1}^{\tau} \hat{\delta}_{jA}^{(k)} \Delta y_{i,t-j} + \hat{b}_{iA}^{(k)} \bar{y}_{t-1}^{(k)} + \sum_{j=1}^{\tau} \hat{c}_{iA,j}^{(k)} \bar{\Delta} y_{t-j}^{(k)} \right). \tag{B.2}$$

2. Step for common correlated effect. Consider the threshold variables $\{q_{t-1}^{(k)}, k = 1, \cdots, K\}$. Obtain, under $K-$regimes, the estimated thresholds $\hat{\lambda}_K = (\hat{\lambda}^{(1)'}, \cdots, \hat{\lambda}^{(K)'})'$ and the estimated $\hat{\phi}_i^{(k)} = (\hat{a}_i^{(k)}, \hat{\beta}_i^{(k)}, \hat{\delta}_1^{(k)}, \cdots, \hat{\delta}_\tau^{(k)}, \hat{b}_i^{(k)}, \hat{c}_{i,1}^{(k)}, \cdots, \hat{c}_{i,\tau}^{(k)})'$, $i = 1, \cdots, N$; $k = 1, \cdots, K$. See (14). By (11) and footnote 9, we calculate the following term for $i = 1, \cdots, N$; $t = 1, \cdots, T$; $k = 1, \cdots, K$:

$$\hat{g}_{i,t-1}^{(k)} = \hat{b}_i^{(k)} \bar{y}_{t-1}^{(k)} + \sum_{j=1}^{\tau} \hat{c}_{i,j}^{(k)} \bar{\Delta} y_{t-j}^{(k)}. \tag{B.3}$$

3. Step for DGP. Let the initial values $y_{i,1}^b = y_{i,1}$ and $\Delta y_{i,2-j}^b = \Delta y_{i,2-j}$, $i = 1, \cdots, N$; $j = 1, \cdots, \tau$. For each $i = 1, \cdots, N$ and $t =$

[19] Here we consider $(K+1)-$regimes so that the residual is as clean as possible.

$2, \cdots, T$, *recursively* generate (under $K-$regimes):

$$\Delta y_{i,t}^b = \hat{\epsilon}_{it}^b + \sum_{k=1}^{K} \sum_{\hat{\underline{\lambda}}^{(k)} < q_{t-1}^{(k)} \leq \hat{\bar{\lambda}}^{(k)}} \left( \hat{a}_i^{(k)} + \hat{\beta}_i^{(k)} y_{i,t-1}^b + \sum_{j=1}^{\tau} \hat{\delta}_j^{(k)} \Delta y_{i,t-j}^b + \hat{g}_{i,t-1}^{(k)b} \right),$$

$$y_{i,t}^b = y_{i,t-1}^b + \Delta y_{i,t}^b, \tag{B.4}$$

where $\hat{\epsilon}_{it}^b$ is randomly drawn from $\{\hat{\epsilon}_{ls},\ l = 1, \cdots, N;\ s = 1, \cdots, T\}$ (see (B.2)) and $\hat{g}_{i,t-1}^{(k)b}$ is randomly drawn from $\{\hat{g}_{i,s-1}^{(k)}, s \in \text{regime } k\}$ (see (B.3)).[20]

4. Step for threshold test statistic. Given the bootstrap sample, repeat the estimation procedure in sub-section 2.3 and obtain the bootstrap versions of $\hat{S}_{NT}(\hat{\lambda}_K)$ and $\hat{S}_{NT}(\hat{\lambda}_{K+1})$, denoted by $\hat{S}_{NT}^b(\hat{\lambda}_K^b)$ and $\hat{S}_{NT}^b(\hat{\lambda}_{K+1}^b)$. Following (16), the bootstrap version of the sup-statistic is:

$$W_{NT}^b(K+1 \mid K) = \frac{T\left[\hat{S}_{NT}^b\left(\hat{\lambda}_K^b\right) - \hat{S}_{NT}^b\left(\hat{\lambda}_{K+1}^b\right)\right]}{\hat{S}_{NT}^b\left(\hat{\lambda}_{K+1}^b\right)}. \tag{B.5}$$

5. Step for critical value. Repeat $B$ times Step 2 to Step 5, where $B$ is a sufficiently large number. Sort $W_{NT}^b(K+1|K),\ b = 1, \cdots, B$ and the upper 1%, 5% and 10% percentiles are the bootstrapped critical values for the threshold test.

The bootstrap procedure for NCIPS and FNCIPS in equations (18) and (20) (under $K$-regimes) are similar, with $\hat{\beta}_i^{(k)} = 0$, $i = 1, \cdots, N$; $k = 1, \cdots, K$ in (B.4) of Step 3. In Steps 1 and 2, instead of using equation (13), we use equation (19).

[20] This is based on the fact that $\hat{g}_{i,t-1}^{(k)}$ is *regime- and individual-* specific.

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# 非線性長期購買力平價說: 使用橫斷面相依異質追蹤資料的檢定

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過去相當多研究長期購買力平價說 (PPP) 的文獻, 在進行單根檢定時, 俱使用追蹤資料以提升其檢定力。本文在橫斷面相依之追蹤資料架構下, 採用非線性門檻自我迴歸 (TAR) 模型, 並針對不同匯率狀態進行追蹤資料的單根檢定。我們提出一個嶄新檢定程序, 結合兩種現有方法的優點: (一) 單根下 TAR 模型的統計推論, 以及 (二) 擴展的 Dickey-Fuller 追蹤資料迴歸式的單根檢定。本文分析資料包含 17 個 OECD 國家在近期浮動匯率期間的實質匯率。結果發現三個不同的匯率狀態, 特別是在馬斯垂克條約 (Maastricht Treaty) 後期, 支持長期購買力平價的證據較爲強烈; 在馬斯垂克條約前期, 當英鎊 (兌美元) 升值時, 存在一些長期購買力平價的證據, 然而在英鎊貶值時, 證據較爲薄弱。即使將資料延長至 2012 年, 這些實證結果仍具穩健性。

**關鍵詞**: 不對稱調整, 橫斷面相依性, 長期購買力平價, 實質匯率, 門檻自我迴歸
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