

Herwartz, Helmut; Xu, Fang

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Reviewing the sustainability/stationarity of current account imbalances with tests for bounded integration

by Helmut Herwartz and Fang Xu

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Christian-Albrechts-Universität Kiel

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Reviewing the sustainability/stationarity of current account imbalances with tests for bounded integration

Helmut Herwartz* and Fang Xu†

May 2006

Abstract: We investigate for 26 OECD economies if their current account imbalances are driven by stochastic trends. Standard ADF results are contrasted with tests accounting for the bounded support of the current account. Neglecting the latter feature might give misleading results in the sense that ADF based conclusions are biased towards the rejection of unit root features. The current account imbalances are found to be bounded nonstationary for most OECD economies. Panel based test statistics confirm the bounded nonstationarity for these series.

Keywords: Current account, bounded unit root tests

JEL Classification: C12, C22, C23, F32, E21, E22

*Institute for Statistics and Econometrics, Christian-Albrechts-University of Kiel, Olshausenstr. 40-60, D-24118 Kiel, Germany, Herwartz@stat-econ.uni-kiel.de.

†Corresponding author. Institute for Statistics and Econometrics, Christian-Albrechts-University of Kiel, Olshausenstr. 40-60, D-24118 Kiel, Germany, FangXu@stat-econ.uni-kiel.de. The authors gratefully acknowledge financial support of Deutsche Forschungsgemeinschaft (HE2188/1-1).

1 Introduction

The sustainability of current account (CA) imbalances has been one of the most discussed issues among economists and politicians since the early 1980s. Husted (1992), Coakley, Kulasi and Smith (1996) and Taylor (2002) have shown that the long-run intertemporal budget constraint (solvency constraint) implies a stationary CA. Time series of CA imbalances are mostly constructed as a share of GDP, which are subjected to unit root testing e.g. by means of Augmented Dickey-Fuller (ADF) tests. The stationarity of CA is still facing a lively discussion. Empirical findings in favor of nonstationary imbalances are often explained to mirror power deficiencies of unit root tests applied to short time spans of data. To address the potential of power loss, Coakley and Kulasi (1997) have applied panel unit root tests (Im, Pesaran and Shin 2002). It appears that the evidence from panel unit root approaches is supportive for the view that the CA is stationary.

As a major caveat of the latter contributions, however, one may argue that the bounded nature of CA series measured as a share of GDP has not yet been taken into account. Traditional unit root tests are formalized to distinguish between stationary processes ($I(0)$) and processes driven by stochastic trends ($I(1)$) that can grow or decrease to any level. Time series as the CA imbalance to GDP are ‘bounded’ by construction. Furthermore, the CA imbalance could be bounded via policy control or economic crisis. The latter argument can be lightened up by the distinction between “solvency” and “sustainability”. If a “drastic” policy shift is needed or a crisis is likely to emerge under continuation of the current policy stance or of private sector behavior, then this CA imbalance lacks sustainability. Sustainability is a more restrictive concept than solvency as argued by Milesi-Ferretti and Razin (1996). As a consequence, a stationary CA detected via rejection of the $I(1)$ hypothesis can be in line with the solvency constraint but does not necessarily imply sustainability. For the latter the ‘bounded’ impact of the policy control or economic crisis on the series shall be taken into account. Thus, from an econometric perspective bounded stationarity is the more natural counterpart of sustainability as is the mere rejection of an unbounded $I(1)$ model.

Recently, the issue of testing for stochastic trends governing bounded processes has been addressed by Cavaliere (2005). An important result from this work is that bounds may alter the asymptotic distribution of the ADF statistic, such that inferential results drawn from standard unit root testing suffer from invalid significance levels. Through testing for bounded integration mean reverting behavior of stationary series could be separated from (spurious) reverting dynamics of nonstationary but bounded processes.

In this paper we analyse the CA balance via the bounded ADF unit root test suggested by Cavaliere (2005). Annual data for 26 OECD countries spanning the period 1971 to 2002 are used. To provide panel statistics, the Fisher test is applied

as in Maddala and Wu (1999). In addition, we test the null hypothesis of bounded integration for domestic investment and saving, exports and imports all measured as a share of GDP. The CA balance can be approximately decomposed either as domestic saving minus investment, or exports minus imports. If the latter series are (bounded) nonstationary, a (bounded) stationary CA might reflect a cointegration relation e.g. between saving and investment. Such a cointegrating relation has been put forth as a possible solution for the so-called “Feldstein-Horioka puzzle” (Feldstein and Horioka 1980), stating that an empirically high correlation between domestic saving and investment is at odds with international capital mobility.

Our findings support the view that the CA balance is a bounded nonstationary process and rejections of the I(1) model might be due to the existence of bounds in the sense of policy controls or crises. Thus, high CA deficits as observed for some OECD members could be in line with solvency while violating sustainability features. Secondly, saving and investment (exports and imports) are not cointegrated. The high association between saving and investment may be partially due to the existence of the bounds on the CA imbalances, and thus might not imply a low capital mobility as argued by Feldstein and Horioka (1980).

The remainder of the paper is organized as follows: in the next Section we sketch the bounded ADF unit root test and the Fisher test. Empirical results are provided in Section 3. Section 4 summarizes the main findings and concludes.

2 Methodology

Let $\{Y_t\}_0^T$ denote a stochastic process which is obtained from mapping the sample paths of a common AR(1) process $\{\mathcal{Y}_t\}_0^T$,

$$\mathcal{Y}_t = \rho \mathcal{Y}_{t-1} + \varepsilon_t, \quad t = 1, \dots, T, \quad \varepsilon_t \sim \text{i.i.d.}(0, \sigma^2), \quad (1)$$

on the interval $[\underline{b}, \bar{b}]$. When $\rho = 1$, $\{Y_t\}$ is called “bounded integrated of order 1”, briefly BI(1). For a detailed discussion of the assumptions underlying the process and the derivation of the test of the BI(1) null hypothesis ($\rho = 1$) against bounded stationarity ($\rho < 1$), the reader is referred to Cavaliere (2005). Consider the common ADF regression

$$\Delta Y_t = (\rho - 1)Y_{t-1} + \sum_{l=1}^p \alpha_l \Delta Y_{t-l} + e_t. \quad (2)$$

In case $Y_0 = 0$ the asymptotic distribution of the t -statistic of $\hat{\rho} - 1$ in (2), denoted λ , is obtained under the BI(1) null hypothesis as

$$\lambda = \frac{B_{\underline{c}}^{\bar{c}}(1)^2 - 1}{(4 \int_0^1 B_{\underline{c}}^{\bar{c}}(s)^2 ds)^{1/2}}, \quad (3)$$

where $B_{\underline{c}}^{\bar{c}}$ is a regulated Brownian motion with bounds at $\underline{c} = \underline{b}(\sigma^2 T)^{-1/2}$, $\bar{c} = \bar{b}(\sigma^2 T)^{-1/2}$. If $Y_0 \neq 0$, the result in (3) holds with (\underline{c}, \bar{c}) replaced by $(\underline{c} - c_0, \bar{c} - c_0)$, where $c_0 = Y_0(\sigma^2 T)^{-1/2}$.

We obtain the BI(1) process $\{Y_t\}_0^T$ by reflecting $Y_{t-1} + \varepsilon_t$ at \underline{b} and \bar{b} . To formalize the latter mapping under the null hypothesis, consider

$$Y_t := Y_{t-1} + \varepsilon_t + \underline{\xi}_t - \bar{\xi}_t, \quad (4)$$

where $\varepsilon_t = \Delta \mathcal{Y}_t$ is defined in (1). The bounded random walk $\{Y_t\}_0^T$ is obtained by setting $\underline{\xi}_t := [2\underline{b} - (Y_{t-1} + \varepsilon_t)]\mathbb{I}\{Y_{t-1} + \varepsilon_t < \underline{b}\}$ and $\bar{\xi}_t := [2\bar{b} - (Y_{t-1} + \varepsilon_t)]\mathbb{I}\{Y_{t-1} + \varepsilon_t > \bar{b}\}$. Once the bounds are reached, the reflected process mirrors the original process around the bounds within the interior state.

Apart from testing the BI(1) hypothesis for single economies, we also adopt a panel perspective. To obtain a suitable test statistic at the pooled level from single country test results, we apply the Fisher test (Maddala and Wu 1999). The latter criterion is derived from a cross sectional set of equations as (2) obtaining test statistics λ_i , $i = 1, \dots, N$, with N being the number of cross sectional entities. Under the null hypothesis $H_0 : \rho_i = 1 \forall i$ and the assumption of cross-sectional independence, a Fisher statistic is obtained as

$$\Lambda = 2 \sum_{i=1}^N \ln(p_i(\lambda_i)) \xrightarrow{d} \chi^2(2N), \quad (5)$$

where $p_i(\lambda_i)$ is the p -value associated with λ_i under the BI(1) null hypothesis.

3 Results

We investigate the prevalence of stochastic trends governing CA, the trade balance, domestic investment and saving, exports and imports for all OECD countries except Czech Republic, Poland, Slovak Republic and Luxembourg. The latter are not contained in the sample owing to data nonavailability or as in the case of Luxembourg for likely peculiar determinants of domestic savings. All series are measured as percentages to GDP and drawn from the World Development Indicators CD-Rom 2004 published by the World Bank. The sample comprises yearly observations made over the period 1971 to 2002. For 4 countries data of domestic saving and investment for 2002 are not available as well as data of exports and imports for 2002 for 3 countries. These missing values are estimated by means of univariate autoregressive models of order 1 with intercept. Besides, we approximate the CA balance as domestic saving minus domestic investment rather than national saving minus domestic investment in order to be consistent with the main empirical discussions of the Feldstein-Horioka puzzle.

Before contrasting test results for the I(1) and BI(1) null hypothesis we sketch the derivation of the latter which requires Monte Carlo simulation. Firstly, the nuisance parameters $\underline{c}_i, \bar{c}_i, c_{i0}$ are estimated for each country i as $\hat{\underline{c}}_i = \underline{b}_i(\hat{\sigma}_i^2 T_a)^{-1/2}$, $\hat{\bar{c}}_i = \bar{b}_i(\hat{\sigma}_i^2 T_a)^{-1/2}$ and $\hat{c}_{i0} = Y_{i0}(\hat{\sigma}_i^2 T_a)^{-1/2}$, where $T_a = 32$ and $\hat{\sigma}_i$ is the OLS estimator of the standard deviation of the residuals in the regression $Y_{it} = \rho Y_{it-1} + e_{it}$. By construction of the variables, their natural bounds are $[-100, 100]$. However, since CA dynamics are likely subjected to policy control there might be hidden bounds which are considerably smaller in absolute value. For the latter reason we choose the relevant bounds as $\underline{b}_i = \min_t(Y_{it})$ and $\bar{b}_i = \max_t(Y_{it})$. Bounded random walks of length $T_s = 1000$ starting with $Y_0 = 0$ are simulated with 10000 replications. For simulation purposes the bounds for the random walk in country i are $[(\min_t(Y_{it}) - Y_{i0})(\hat{\sigma}_i^2 T_a)^{-1/2} T_s^{1/2}, (\max_t(Y_{it}) - Y_{i0})(\hat{\sigma}_i^2 T_a)^{-1/2} T_s^{1/2}]$. The (cross section specific) distribution in (3) is approximated by the distribution of ADF statistics obtained from the bounded random walks¹.

The results of (bounded) unit root tests for the CA balance are shown in Table 1. Throughout, our discussion of test results is conditional on the 5% significance level. For the estimated bounds of the regulated Brownian motion, Ireland has the highest upper bound (1.34) and the U.S. (-1.18) has the smallest lower bound. For all 26 OECD economies the critical values from the BI(1) process (cv_b) are considerably smaller in comparison with the corresponding critical values implied by a conventional I(1) model ($cv_{ub} \approx -1.94$). As can be seen from the p -values in the middle three columns of Table 1, for 10 out of 26 countries the I(1) hypothesis is rejected (p_{ub}). However, the BI(1) hypothesis is only rejected for Korea (p_b). Thus, for 9 of the initially 10 significant test statistics the mean reverting behavior of the CA appears spurious and might be due to the bounded nature of the data.

It might be argued that the diagnosis of bounded nonstationarity for the CA is crucially depending on the choice of bounds. To address this issue we simulate bounded processes having $[\underline{c}_i^* - \hat{c}_{i0}, \bar{c}_i^* - \hat{c}_{i0}]$, $\underline{c}_i^* = \underline{c}_i - \delta$, $\bar{c}_i^* = \bar{c}_i + \delta$ and select the parameter δ such that the bounded model obtains the same p -value for the ADF statistic as the unbounded model ($p_b^* = p_{ub}$). From the latter implicit bounds, we estimate the implied counterparts for the observable data and provide the latter as $[\underline{b}_i^*, \bar{b}_i^*]$ for each country in the right hand side panel of Table 1. To have the same p -values of the ADF statistics for the 10 economies diagnosed as unbounded stationary (rejections implied by DF distribution), the implied bounds are between 1.56 (Italy) to about 2.18 (UK) times wider in comparison with their empirical counterparts. Most of these bounds likely reach above some country specific sustainable CA levels.

¹The empirical results with simulated sample size $T_s = 2000$ are similar as those from $T_s = 1000$. Moreover all results documented in this work are qualitatively identical if censoring is used as an alternative mapping scheme, or the variance parameter σ_i^2 is estimated by imposing $\rho = 1$. For space considerations we do not provide results obtained along the latter lines.

Considering Mexico, for instance, the latter implicit interval equalizing p-values of testing the I(1) and BI(1) model is $[-11.92, 16.46]$. However, a CA deficit about 5 percentage of GDP in Mexico has already induced a financial crisis at the end of 1994.

As reported in Table 2, testing the BI(1) null hypothesis for domestic saving, bounded nonstationarity cannot be rejected for all 26 OECD countries except Japan. Testing the I(1) hypothesis for domestic investment, stationary behavior is diagnosed for Germany and Japan. Considering bounds, while the German investment series is still found stationary, Japanese investment is classified as bounded nonstationary. To summarize, the domestic saving and investment are bounded nonstationary processes for almost all OECD economies considered. Taking the observed maximum and minimum values of domestic savings and investment as their true bounds might be criticized since these variables are likely less under the policy control as the CA balance. However, since these series are mostly nonstationary, taking account of bounds strengthens the empirical evidence in favor of their nonstationarity.

Neglecting the net capital gains on foreign assets, the CA balance can be (approximately) decomposed as exports minus imports. Results of BI(1) and I(1) tests for the latter series and the trade balance (exports minus imports) are also shown in Table 2. For 11 out of 26 economies, the trade balance is found stationary when testing the I(1) null hypothesis (p_{ub}). For 10 of the latter economies the BI(1) null hypothesis cannot be rejected (p_b). The time series of exports and imports are found nonstationary and bounded nonstationary, respectively, for all economies considered.

Finally, results from Fisher tests are shown in Table 3. As one may expect for the standard panel unit root tests, the CA and the trade balance appear stationary at the aggregate level. When testing the BI(1) model at an aggregate level, none of the series is diagnosed as bounded stationary. It might be argued that CA account imbalances in the OECD likely show contemporaneous correlation of underlying innovations. The latter feature has been ignored when constructing the Fisher statistics (Banerjee and Zanghieri 2003). Simulating the bounded processes from innovations that exhibit the empirical covariance features of cross sectional ADF residuals, however, turns out to even further support the null hypothesis of bounded integration governing CA and trade balance dynamics. Empirical results on the latter issue are not shown for space considerations but available from the authors upon request.

4 Conclusions

In this paper we contrast for 26 OECD countries results obtained from testing the I(1) and BI(1) null hypotheses by means of the ADF statistic. Taking account of bounds weakens throughout the evidence against the prevalence of stochastic trends. At the pooled level the null hypothesis of bounded integration cannot be rejected for the CA

and trade balance thereby providing evidence against a unit cointegration relation linking domestic saving and investment or exports and imports.

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Table 1: Unit root tests for CA

Country	ADF	\underline{b}_i	\bar{b}_i	$\underline{\hat{c}}_i - \hat{c}_{i0}$	$\bar{\hat{c}}_i - \hat{c}_{i0}$	cv_b	p_{ub}	p_b	\underline{b}_i^*	\bar{b}_i^*	$\underline{\hat{c}}_i^* - \hat{c}_{i0}$	$\bar{\hat{c}}_i^* - \hat{c}_{i0}$
AUS	-2.35	-3.56	3.17	-0.57	0.31	-2.82	0.02	0.19	-6.90	6.51	-1.01	0.75
AUT	-2.80	-2.88	1.25	-0.68	0.15	-3.18	0.01	0.07	-5.07	3.44	-1.12	0.59
BEL	-0.44	-3.36	4.40	-1.10	0.45	-2.29	0.52	0.79	-10.55	11.59	-2.54	1.89
CAN	-0.87	-1.82	5.22	-0.52	0.53	-2.21	0.34	1.00	-8.10	11.51	-1.46	1.47
DEN	-0.64	-4.46	6.81	-0.31	0.97	-2.85	0.44	0.73	-16.39	18.75	-1.67	2.33
FIN	-0.48	-5.97	9.24	-0.36	1.08	-2.57	0.52	0.77	-21.93	25.20	-1.88	2.60
FRA	-1.62	-3.20	2.99	-0.65	0.48	-2.22	0.10	0.33	-5.83	5.62	-1.13	0.96
GER	-0.35	-4.98	3.94	-0.65	1.18	-1.98	0.57	0.81	-11.60	10.56	-2.01	2.54
GRC	-0.58	-9.82	-2.18	-0.42	0.61	-2.37	0.46	1.00	-20.40	8.41	-1.84	2.03
HUN	-3.21	-9.16	4.57	-0.17	0.61	-3.22	0.00	0.05	-15.20	10.61	-0.51	0.95
ISL	-3.44	-9.74	5.31	-0.18	0.67	-3.41	0.00	0.05	-15.01	10.58	-0.48	0.97
IRL	-0.30	-16.88	16.00	-0.51	1.34	-2.17	0.58	0.79	-43.65	42.77	-2.01	2.84
ITA	-2.12	-4.17	4.93	-0.57	0.50	-2.21	0.03	0.07	-6.73	7.49	-0.87	0.80
JPN	-1.12	-0.91	3.88	-0.72	0.25	-3.14	0.25	0.65	-5.78	8.75	-1.70	1.23
KOR	-2.86	-11.25	13.46	-0.03	1.11	-2.14	0.00	0.01	-23.34	25.55	-0.59	1.67
MEX	-2.09	-5.03	9.58	-0.29	0.77	-2.98	0.04	0.24	-11.92	16.46	-0.79	1.27
NLD	-0.38	-1.58	6.14	-0.04	1.07	-2.16	0.54	0.70	-15.17	19.72	-1.98	3.01
NOR	-0.94	-9.16	17.33	-0.29	0.87	-2.94	0.30	0.63	-33.77	41.95	-1.37	1.95
NZL	-2.92	-12.08	3.46	-0.85	0.18	-3.33	0.00	0.08	-17.85	9.22	-1.23	0.56
PRT	-0.54	-15.79	-1.46	-0.73	0.27	-3.07	0.48	0.98	-37.01	19.76	-2.21	1.75
SPA	-2.68	-4.73	1.96	-0.69	0.16	-3.22	0.01	0.09	-8.34	5.58	-1.15	0.62
SWE	-0.42	-1.55	7.25	-0.36	0.76	-2.60	0.54	0.99	-14.70	20.39	-2.04	2.44
SWI	-0.54	-3.34	6.47	-0.23	0.96	-3.32	0.49	0.77	-15.85	18.99	-1.75	2.48
TUR	-0.85	-7.49	2.40	-0.29	0.32	-3.28	0.34	1.00	-26.77	21.69	-1.47	1.50
UK	-1.98	-4.98	2.80	-0.89	0.20	-3.38	0.04	0.21	-9.58	7.40	-1.53	0.84
US	0.42	-4.08	0.98	-1.18	0.26	-3.13	0.80	1.00	-12.58	9.48	-3.60	2.68

Results from I(1) and BI(1) tests. ADF-test statistics for the 26 OECD countries are shown with the observed bounds $[\underline{b}_i, \bar{b}_i]$ and the corresponding estimated bounds $[\underline{\hat{c}}_i - \hat{c}_{i0}, \bar{\hat{c}}_i - \hat{c}_{i0}]$ for the regulated Brownian motion. 5% critical values are shown for testing the the BI(1) hypothesis cv_b . The corresponding 5% critical value cv_{ub} for testing I(1) hypothesis is about -1.94. Respective p-values are p_{ub}, p_b . p-values smaller than 0.05 are highlighted in bold face. $[\underline{\hat{c}}_i^* - \hat{c}_{i0}, \bar{\hat{c}}_i^* - \hat{c}_{i0}]$ are the implicit obtaining $p_b^* = p_{ub}$. The corresponding nominal bounds are $[\underline{b}_i^*, \bar{b}_i^*]$. The considered countries are AUS-Australia, AUT-Austria, BEL-Belgium, CAN-Canada, DEN-Denmark, FIN-Finland, FRA-France, GER-Germany, GRC-Greece, HUN-Hungary, ISL-Iceland, IRL-Ireland, ITA-Italy, JPN-Japan, KOR-Korea (Rep.), MEX-Mexico, NLD-Netherlands, NOR-Norway, NZL-New Zealand, PRT-Portugal, SPA-Spain, SWE-Sweden, SWI-Switzerland, TUR-Turkey, UK-United Kingdom and US-United States.

Table 2: Further unit root tests

Country	Saving			Investment			Trade Balance			Export			Import		
	ADF	p_{ub}	p_b	ADF	p_{ub}	p_b	ADF	p_{ub}	p_b	ADF	p_{ub}	p_b	ADF	p_{ub}	p_b
AUS	-1.35	0.17	0.19	-0.72	0.41	0.73	-2.35	0.02	0.19	0.76	0.88	0.98	1.38	0.96	1.00
AUT	-1.16	0.22	0.34	-1.03	0.27	0.50	-2.80	0.01	0.07	2.38	1.00	1.00	1.77	0.98	1.00
BEL	-0.98	0.29	0.28	-1.12	0.24	0.43	-0.44	0.52	0.79	1.34	0.95	1.00	1.10	0.93	1.00
CAN	-0.16	0.63	1.00	-0.51	0.50	1.00	-0.82	0.36	1.00	1.58	0.97	0.97	2.03	0.99	1.00
DEN	0.14	0.72	1.00	-0.83	0.36	0.61	-0.64	0.44	0.73	1.58	0.97	1.00	0.69	0.86	1.00
FIN	-0.42	0.54	0.96	-1.07	0.27	0.52	-0.48	0.52	0.77	0.96	0.91	0.99	0.23	0.76	1.00
FRA	-1.41	0.15	0.31	-1.18	0.22	0.43	-1.62	0.10	0.33	1.58	0.97	1.00	0.68	0.86	1.00
GER	-1.24	0.20	0.19	-2.00	0.04	0.04	-0.35	0.57	0.81	2.62	1.00	1.00	1.35	0.96	1.00
GRC	-1.07	0.26	0.50	-0.92	0.32	0.71	-0.58	0.46	1.00	0.47	0.82	0.94	0.68	0.86	0.97
HUN	-1.03	0.28	0.43	-1.14	0.23	0.42	-3.21	0.00	0.05	1.08	0.93	0.99	0.90	0.90	1.00
ISL	-0.97	0.30	0.53	-1.35	0.17	0.33	-3.44	0.00	0.05	-0.01	0.68	1.00	-0.61	0.46	0.75
IRL	2.11	0.99	1.00	-0.25	0.59	1.00	-0.37	0.55	0.77	4.28	1.00	1.00	1.54	0.97	1.00
ITA	-0.99	0.29	0.59	-0.67	0.43	0.99	-2.12	0.03	0.07	1.04	0.92	1.00	0.60	0.85	1.00
JPN	-2.40	0.01	0.01	-2.11	0.03	0.22	-2.02	0.04	0.24	-0.33	0.57	1.00	-0.29	0.58	1.00
KOR	0.46	0.82	0.84	-0.29	0.58	1.00	-2.86	0.00	0.01	0.50	0.82	0.95	0.51	0.83	1.00
MEX	-0.31	0.58	0.89	-0.22	0.61	1.00	-2.09	0.04	0.24	0.63	0.85	0.97	1.60	0.97	0.99
NLD	-0.32	0.57	0.88	-1.35	0.16	0.15	-0.38	0.54	0.70	0.71	0.87	1.00	0.42	0.80	1.00
NOR	-0.04	0.67	1.00	-1.22	0.20	0.48	-0.94	0.30	0.63	0.17	0.73	1.00	-0.88	0.33	0.81
NZL	-0.58	0.47	1.00	-0.59	0.46	1.00	-2.92	0.00	0.08	0.75	0.88	1.00	0.32	0.78	1.00
PRT	-0.76	0.39	0.92	-0.23	0.60	1.00	-0.53	0.49	0.98	0.37	0.79	1.00	0.46	0.82	1.00
SPA	-0.55	0.48	0.64	-0.41	0.53	1.00	-2.68	0.01	0.09	0.70	0.87	0.88	1.73	0.98	1.00
SWE	-0.60	0.46	0.65	-1.03	0.28	0.44	-0.42	0.54	0.99	1.44	0.96	1.00	0.85	0.90	1.00
SWI	-1.21	0.21	0.19	-1.65	0.10	0.09	-0.54	0.49	0.77	1.64	0.98	1.00	0.33	0.78	1.00
TUR	-0.02	0.68	0.81	-0.24	0.60	0.95	-0.85	0.34	1.00	1.17	0.94	0.99	1.48	0.96	1.00
UK	-1.38	0.16	0.23	-0.68	0.42	1.00	-1.98	0.04	0.21	0.18	0.73	1.00	0.27	0.76	1.00
US	-0.97	0.30	0.63	-0.30	0.57	1.00	0.42	0.80	1.00	0.42	0.80	0.84	1.80	0.99	0.99

ADF-statistics and p -values p_{ub} and p_b . For further notes see also Table 1.

Table 3: Panel unit root tests

Variable	I(1) Test	BI(1) Test	Variable	I(1) Test	BI(1) Test
Current Account	130.647 (0.00)	58.626 (0.25)	Trade Balance	134.288 (0.00)	60.657 (0.19)

Fisher test results. p -values appear in parentheses below the test statistic Λ from (5).