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**CURRENT ACCOUNT AND RELATIVE PRICES:
Cointegration in the presence of Structural Breaks in Emerging Economies**

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Abstract

The aim of this study is to examine the long-run relationship between the current account and relative prices such as terms of trade and real exchange rate for the emerging economies. These variables have been exposed to large fluctuations for more than the last two decades nearly in all emerging economies. Therefore, structural breaks have to be taken into account in estimations. Therefore, the recent panel cointegration method developed by Westerlund (2006) was applied to the current account model allowing for structural breaks. The estimations of unit root tests proposed by Levin et al. (2002), Im et al. (2003) and by Hadri (2000) provided the evidence of the unit root existence in our series. The Hansen's (1992) stability test illustrated the instability exist in series except for the cases of India and Turkey. The Westerlund (2006) cointegration test estimations detected multiple structural shifts in every panel case; however, the hypothesis of cointegration in the panel could not be accepted by the Lagrange Multiplier statistics.

JEL: F32; F41

Key Words: Current account; Terms of trade; Panel cointegration; Structural breaks

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1. Introduction

The process of current account adjustment has always been at the centre of attention within the area of studies in international macroeconomics. This attention mainly has been due to the increase in the current account deficit in most of the emerging economies. Our focus is mainly on the process and the determinants associated with large current account reversals.

The behaviours of the developing country and the developed country are expected to vary along several dimensions. More specifically, the developing economy relies heavily on a narrow range of primary commodities for its export earnings. The instability of these earnings is important and highly dependent on the fluctuations in the relative prices of primary commodities. In addition, there are two major factors that play a role in the stability of a developing economy. First is the share of imported capital good intermediate inputs in domestic production and mainly exporting goods. Second is the fraction of the export revenues in using large foreign paybacks. Thus, the relative prices could play an important role in business cycle fluctuations and in determining the economic activities in developing countries.

The intertemporal approach views the current account balance as the outcome of forward-looking dynamic saving and investment decisions. The intertemporal approach of the current account behaviour was a common area of research in the 1980s (Buiter, 1981; Obstfeld, 1982; Sachs, 1981; and Svensson and Razin, 1983). Using current account as a percentage of GDP is important in terms of evaluating the success of economic performance in developing countries. In an open economy, a country's external balance is determined by the interplay between the country's expectations of future income (relative to those of its trading partners) and the cost of the necessary borrowing or lending that the country has to engage in with the purpose of smoothing its consumption over time. An increase in expected disposable income raises trade deficit causing an increase in current account deficit. The determinant of a country's current account is the citizens' desire to smooth consumption over time. Due to time-varying interest rates, exchange rates and other relative prices such as terms of trade, they affect the equilibrium of current account and therefore output to be modelled explicitly.

The objective of this paper is to examine the role of fluctuations in relative prices in determining the external balance for emerging economies using a stochastic dynamic model. We selected countries such as Argentina, Brazil, China, Korea, India, Indonesia, and Turkey

and we used annual data from 1982 to 2008. The model used in this paper includes various structural changes in a panel analysis. The question that will be asked in this paper is whether cointegration relationships exist between current account and relative price variables in the emerging economies of Argentina, Brazil, China, Korea, India, Indonesia and Turkey.

The rest of the paper is organised as follows. Section 2 briefly explains the literature review. Section 3 gives an overview of the economies of these countries and major political implications during the time period chosen in this paper. It also comments on the evolution of relative prices such as terms of trade and exchange rates. Section 4 describes the intertemporal current account model used in this paper and its testable implications. Section 5 explains the methodology for testing various structural breaks for the model developed in section 4. Section 6 provides detailed estimation results and finally section 7 concludes.

2. Literature Review

The main goal of this paper is to determine the behaviour of the current account and the evolution of current account and its determinants. Terms of trade shocks are regarded as a major force driving business cycle fluctuations in small open economies. Harberger (1950) and Laursen and Metzler (1950) (HLM) used a Keynesian model and showed that an exogenous rise in the terms of trade of a small open economy would increase real income – given a constant marginal propensity to consume of less than one. Accordingly, this would cause a rise in private savings and an improvement of the current account. This is called the HLM effect. Later, the discussion about the relation between terms of trade and current account continued with the transmission of disturbances in open economy macroeconomics (see, for example, Mussa 1979; Dornbusch 1980). Especially, when the oil price of an imported intermediate inputs caused deterioration in the current account (Findlay and Rodrigues 1977; Buiters 1978; Bruno and Sachs 1979). In these analyses the responses of current account is consistent with the HLM effect, although the effect is not always used in deriving it (see Svensson and Razin 1983).

The HLM effect has been examined within deterministic intertemporal models by Sachs (1981), Obstfeld (1982), and Swensson and Razin (1983). In the intertemporal current account models first introduced by Sachs (1981), he studied the behaviour of current account in the less developed countries in the 1970s. His argument was that the current account deficits were responses to terms of trade movements and, in a dynamic framework, the HLM effect depends on the duration of the shock. Only if the shock is temporary does the HLM

effect appear. Obstfeld (1980, 1982) also questioned the validity of the HLM effect. He argued that a deterioration of TOT will result in increased saving and an improvement, rather than a deterioration, in the current account. Svensson and Razin (1983) generalized the results of Sachs and Obstfeld by distinguishing between current and future changes in the terms of trade. They concluded that a temporary terms of trade deterioration in a small open economy implies a deterioration of its trade balance. A future terms of trade deterioration implies an improvement of the trade balance and, finally, a permanent terms of trade deterioration has an ambiguous effect, depending on the rate of time preference. Later, these intertemporal current account models have been used extensively in the literature.

Recent studies by Iscan (2000), Blanchard and Giavazzi (2002), Zanas (2004), Broda (2004), Huang and Meng (2007), Bauakez and Kano (2008), Santons-Paulino (2007), Cashin and Mc Dermott (2002) examined the relationship between the terms of trade and current account. Campa and Gavilan (2006) examined the current account in the euro area. Blanchard and Giavazzi (2002) found that the correlation between average output per capita and its external balance relative to GDP are positively related. Under the standard assumption in neoclassical growth models that low income countries have higher growth potential than higher income countries, which is consistent with the consumption smoothing hypothesis. They also found that the absolute value of correlation increases as the degree of economic integration rises.

3. Current Account and Terms of Trade in Developing Countries

The developing countries became large borrowers in the international markets that also caused unsustainable debt levels. Therefore, due to the close link between goods and financial markets, the theories assume that developing countries with higher rates of return should see an increase in investment. Also as countries with higher growth prospects, they should see a decrease in investment. So, on both counts, developing countries should run larger current account deficits and developed countries, on the contrary, should run larger current account surpluses. There is evidence to assume that it was the savings rather than investment that accounts for explaining the current account fluctuations (Blanchard and Giavazzi, 2002). This section looks at the concept of current account and the behaviours, and current account and relative prices.

[Insert Figure 1]

[Insert Figure 2]

Figure 1 shows the share of current account in GDP for the selected emerging economies. Similarly, all these countries have had large current account fluctuations for the last two decades and they had to deal with large current account deficits for most of the 1990s. Figure 2, on the other hand, shows terms of trade data for the selected countries. There is a clear visual observation that terms of trade have fluctuated nearly for two decades in all these countries. There are downward trends in all countries except Argentina and India.

4. An Intertemporal Model of the Current Account

A standard intertemporal current account model is used to examine the fluctuations in the current account. The model considers a small open economy where consumers smooth consumption over time (Campbell, 1987). Thus, the optimal consumption is based on the expectations of future output and relative prices. Current account balances in every period are the difference between optimal consumption and net output in that period. The model considers time-varying interest rates and relative prices (exchange rates and terms of trade) through the existence of traded and nontraded goods.

The effects of the terms of trade fluctuations on current account can be understood by examining a simple two-period deterministic model. Following Dornbush (1983), we consider a small country that can borrow and lend with the rest of the world at a time-varying interest rate. There are two types of goods: traded and non-traded goods. The representative household consumes a mix of tradable and non-tradable goods and has the following lifetime utility:

$$U(C_t) = \frac{C_t^{1-1/\sigma}}{1-1/\sigma} + \beta \frac{C_{t+1}^{1-1/\sigma}}{1-1/\sigma}, \quad 0 < \beta < 1, \quad \sigma > 0, \quad (1)$$

where β is the subjective discount factor and σ is the elasticity of intertemporal substitution.

Alternatively, the representative household makes the decisions regarding consumption and borrowing and solves an intertemporal maximisation problem choosing a path of consumption and debt that maximises discounted lifetime utility:

$$\max E_0 \sum_{t=0}^{\infty} \beta^t U(C_t^T, C_t^N) \quad (2)$$

$$s.t. Y_t - (C_t^T + P_t C_t^N) - I_t - G_t + r_t B_{t-1} = B_t - B_{t-1} \quad \forall t \quad (3)$$

where C_t^T and C_t^N denotes consumption by the households in traded and non-traded goods, P_t is the price of non-traded goods in terms of traded goods. Y_t denotes the value of current output, I_t is investment expenditure and G_t is government spending on goods and services, all measured in terms of traded goods. B_t denotes the stock of foreign assets at the beginning of the period, and r_t is the net world interest rate the country faces in terms of traded goods. The left-hand side of the budget constraint in equation (3) may be interpreted as the current account. Moreover, the total consumption index, C_t , takes the following Cobb-Douglas form:

$$C_t = \omega_1 (C_t^T)^\epsilon (C_t^N)^{1-\epsilon}, \quad 0 < \epsilon < 1, \quad (4)$$

where ϵ is the weight of the tradable good in consumption basket, and $\omega_1 \equiv \epsilon^{-\epsilon} (1-\epsilon)^{\epsilon-1}$ is a positive parameter. Following Bouakez and Kano (2008), we assume the tradable good to be the numeraire and normalise its price to 1.

$$P_t^c = Q_t^{1-\epsilon} \quad (5)$$

where P_t^c denotes the consumption based price index and Q_t is the price of the non-tradable good where the real exchange rate is used as a proxy. Following Obstfeld (1996) and (Obstfeld and Rogoff, 1996, p. 266), the tradable good takes the following Cobb-Douglas form (Bouakez and Kano, 2008, p. 262) :

$$C_t^T = \omega_2 X_t^\gamma M_t^{1-\gamma}, \quad 0 < \gamma < 1, \quad (5)$$

where X_t is consumption of exportable goods, M_t is consumption of importable goods, γ is the weight of exportable goods in the traded-good basket, and $\omega_2 \equiv \gamma^{-\gamma} (1-\gamma)^{\gamma-1}$ is a positive parameter. When we include the price of exportable goods P_t^x and the price of importable goods P_t^m to the model, then we must assume the following condition.

$$1 = (P_t^x)^\gamma (P_t^m)^{1-\gamma} \quad (6)$$

Furthermore, the terms of trade, P_t' , is defined as the relative price of exports in terms of imports and it can be expressed as a function of the price of an exportable goods.

$$P_t' \equiv P_t^x / P_t^m = (P_t^x)^{1/(1-\gamma)} \quad (7)$$

In our small open economy, we follow the assumptions of Bouakez and Kano (2008) where the only tradable assets are one-period risk-free international bonds and they are

indexed to the tradable consumption basket. Also, the representative household can borrow and lend freely in the international market at net world interest rate, r_{t+1} , to smooth consumption across two periods. Finally, this country can neither change the world real interest rate nor the terms of trade.

Therefore, when the representative household allocates its income to the consumption of goods at period t , this includes the consumption of non-tradable, exportable and importable and the purchase of international bonds. Also, the household receives interest payment on its holding of bonds in period $t+1$. We also assume that the following equation is valid.

$$P_t^x X_t + P_t^m M_t + Q_t C_t^N = P_t^c C_t \quad (8)$$

If NY^x and NY^n denote the exportable net outputs and non-tradable net outputs, respectively, then the household's intertemporal budget constraints take the following form (Bouakez and Kano, 2008, p. 263):

$$P_t^c C_t + \frac{1}{1+r_{t+1}} P_{t+1}^c C_{t+1} = P_t^x NY_t^x + Q_t NY_t^n + \frac{1}{1+r_{t+1}} [P_{t+1}^x NY_{t+1}^x + Q_{t+1} NY_{t+1}^n] \quad (9)$$

The first-order conditions for this problem derive the following optimal consumption profile:

$$1 = \beta(1+r_{t+1}) \left(\frac{P_t^c}{P_{t+1}^c} \right) \left(\frac{C_t}{C_{t+1}} \right)^{1/\sigma} \quad (10)$$

The current account, CA, is as follows:

$$CA_t \equiv B_{t+1} = P_t^x NY_t^x - C_t^T \quad (11)$$

Finally, Bouakez and Kano (2008) express the current account with intertemporal budget constraint as follows:

$$CA_t = P_t'^{(1-\gamma)} NY_t^x - \frac{1}{1 + \beta^\sigma (1+r_{t+1})^{\sigma-1} (Q_t / Q_{t+1})^{(\sigma-1)(1-\epsilon)}} \left[P_t'^{(1-\gamma)} NY_t^x + \frac{1}{1+r_{t+1}} P_{t+1}'^{(1-\gamma)} NY_{t+1}^x \right] \quad (12)$$

This expression shows the effects of changes in the world real interest rate, the real exchange rate, and the terms of trade on the current account. According to Obstfeld and Rogoff (1996) and Bouakez and Kano (2008), there are three distinctive effects of a rise in the world interest rate, r_{t+1} , on current account. First, if the world interest rate rises above its permanent level, this increases the consumption-based real interest rates together with the price of current consumption in terms of future consumption. Thus, a representative consumer

tilts its consumption towards the future and increases savings in the current period. This intertemporal substitution effect improves the current account. Second, if the price of future consumption is lower, a rise in consumption-based real interest rate increases current consumption, reducing savings and thereby worsening the current account. Lastly, if the world real interest rate increases, the market discount factor decreases together with the present value of lifetime income. The negative wealth effect reduces current consumption and improves current account.³

Additionally, there are two other variables that affect current account according to equation 12 and they are also used in this paper; real exchange rate and terms of trade. First, a rise in the real exchange rate increases the consumption-based real interest rate. So we may expect intertemporal substitution and income effects similar to those discussed above, causing a reduction in the consumption of tradable goods and therefore a reduction in total current consumption and improvement in the current account, or vice versa. Second, when there is an improvement in terms of trade, the present value of lifetime income increases and that leads to an increase in current household consumption. We assume that the marginal propensity to consume is less than 1; then, current consumption rises less than current income and the current account improves. This is called the HLM effect. So the relationship between the terms of trade and consumption is important in determining the existence of the HLM effect. If there is no relationship then there will be no HLM effect.

In the following sections, we will empirically test the long-run relationship between current account and the price of tradable goods and price of non-tradable goods. Therefore, the equation 12 will take the following reduced form as follows:

$$CA_{t,i} = \alpha + \eta P'_{t,i} + \mu Q_{t,i} + \varepsilon_{t,i} \quad (13)$$

where $CA_{t,i}$ denotes the current account at time t , for country i , P' is terms of trade, and Q is

the real exchange rate measured as $Q = \frac{XR}{CPI_i} CPI_{US}$, where XR is the currency in country i per

dollar, CPI_{US} is the consumer price index for the US and CPI_i is the consumer price index in country i . Finally, ε is the error term. It is important to draw attention to the point where the aim of this analysis is not to test the intertemporal model, but to use the model in determining the variables of current account, then to test the long-run relationship between current account

³ See Obstfeld (1996, chapter 1) and Bouakez and Kano (2008, p. 264) for further discussion of the effects of real interest rate fluctuation on consumption and current account.

and relative prices in emerging economies the determining the time of various structural breaks.

5. Methodology

Panel unit root tests

In our paper we used three different tests for the panel unit root. The first one was the Levin, Lin and Chu (LLC) test (Levin et al. 2002), which is based on orthogonalized residuals and the correction by the ratio of the long-run to the short-run variance of each variable. Although the LLC test has become a widely accepted panel unit root test, it has homogeneity restriction, allowing for heterogeneity only in the constant term of the ADF regression. The second applied test was the Im, Pesaran and Shin (IPS) test, which is a heterogeneous panel unit root test based on individual ADF tests. It was proposed by Im et al. (2003) as a solution to the homogeneity issue. This test allows for heterogeneity in both the constant and slope terms of the ADF regression. Finally, the third test used in our paper was again the heterogeneous panel unit root test, the PKPSS. This test was presented by Hadri (2000) as an extension of the test of Kwiatkowski et al. (1992), the KPSS (Kwiatkowski–Phillips–Schmidt–Shin) to a panel with individual and time effects and deterministic trends, which has as its null the stationarity of the series.

However, the considered unit root tests do not take into account the presence of any structural shifts in series. In the case where we consider developing countries such as Argentina, Brazil, China, Korea, India, Indonesia and Turkey, it is necessary to allow our estimations for structural shifts. The period considered in the study 1982-2008 is full of changes not only in current account determinants, but in most of the economic conditions of these countries as well. Therefore, additionally we employed panel unit root tests based on LM statistics. One of them is the Lagrange-Multiplier (LM) test designed by Schmidt and Phillips (1992), which is the univariate unit root test with no structural shifts, and which is employed in our study for comparison purposes. Another LM unit root test is proposed by Im et al. (2005), the panel extension of the Schmidt and Phillips (1992) test allowing for a structural shift in the trend of a panel and of every individual time series.

Amsler and Lee (1995) in their study provide evidence that the asymptotic distribution of the LM unit-root test does not change with the inclusion of dummy variables in the unit-root regression. Im et al. (2005) in their study of the panel version of the LM unit root test

illustrate that the size of distortions and loss of power in unit root tests remain insignificant when structural shifts are accommodated in cases when shifts do not exist. However, in the opposite situation, when unit root tests were applied to the time series without taking into account existing structural shifts' size distortions and loss of tests' power were found to be significant. In addition panel LM tests with structural breaks are not only robust when structural shifts exist in series, but at the same time they are more powerful than Dickey-Fuller type tests (for example, IPS). The break date in the Im et al. (2005) test is chosen using the minimum LM statistics of Lee and Strazicich (2003, 2004). In this method, the break date is selected when the t-statistic of possible break points is minimized.

Stability Test

The aim of our paper is to estimate the panel cointegration relationship between variables of equation 13 in the presence of multiple structural breaks. Before proceeding with panel cointegration we have to test the cointegration relationships in considered countries for the presence of structural shifts. To estimate parameter stability in cointegration relationships we employed Hansen's (1992) stability test. The test is based on the fully modified OLS residuals proposed by Phillips and Hansen (1990). The test does not require the selection of the structural shifts' location. However, a necessary requisite of the test is that variables have to be non-stationary.

Panel cointegration tests

McCoskey and Kao (1998) develop a residual-based Lagrange Multiplier test for the null hypothesis of cointegration in the panel data. The model they consider allows for varying slopes and intercepts across units:

$$y_{t,i} = \alpha_i + \beta_i x_{t,i} + e_{t,i} \quad (14)$$

where $e_{t,i} = \theta \sum_{j=1}^t u_{i,j} + u_{t,i}$.

The null hypothesis of cointegration is $H_0 : \theta = 0$, against the alternative $H_0 : \theta \neq 0$. Under the null hypothesis we have $e_{t,i} = u_{t,i}$ and the equation above is a system of cointegrated regressors. The panel test statistic is given by:

$$LM = \frac{\frac{1}{N} \sum_{i=1}^N \frac{1}{T^2} \sum_{t=1}^T S_{it}^{+2}}{s^{+2}} \quad (15)$$

where s^{+2} is the partial sum of estimated residuals:

$$S_{it}^+ = \sum_{j=1}^t e_{ij}^+ \quad \text{and} \quad s_{it}^{+2} = \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T e_{it}^{+2} \quad (16)$$

The residuals e_{it}^+ can be estimated using either the dynamic OLS (DOLS) estimator proposed by Saikkonen (1991), or the fully modified OLS (FMOLS) estimator of Phillips and Hansen (1990). Kao and Chiang (1999) show that the DOLS estimator is more powerful than the FMOLS estimator, and this is the method we employ. Whenever there is serial correlation, we use the Stock and Watson (1993) dynamic GLS (DGLS) estimator. McCoskey and Kao (1998) show that the standardised version of the LM statistic converges to a normally distributed random variable under the null of cointegration,

$$LM^* = \frac{\left[\sqrt{N} (LM - \mu_v) \right]}{\sigma_v} \Rightarrow N(0,1) \quad (17)$$

where μ_v and σ_v^2 are, respectively, the expected mean and standard deviation of a complex functional of Brownian motion, obtained by Monte Carlo simulation and tabulated in McCoskey and Kao (1998, Table 1).

The panel cointegration estimations of this paper allowing for structural breaks are based on the approach developed by Westerlund (2006). This is the panel cointegration test that allows accommodating multiple structural breaks in the level as well as in the trend of cointegrated regression. This test is based on the panel cointegration residual-based LM test proposed by McCoskey and Kao (1998), which does not allow for structural shifts. The advantage of Westerlund's test is that it allows for the possibility of known a priori multiple structural breaks or it allows for breaks the locations of which are determined endogenously from the series. At the same time this test allows for a possibility of structural breaks that may be placed at different locations in different individual series. Westerlund (2006) showed in his work that the test is free of nuisance parameters under the null hypothesis and that the number and location points of structural shifts do not affect the limiting distribution. The null of the test is

$H_0 : \phi_i = 0$ for all $i = 1, \dots, N$,

versus alternative hypothesis:

$H_1 : \phi_i \neq 0$ for $i = 1, \dots, N_1$, and $\phi_i = 0$ for $i = N_1 + 1, \dots, N$.

One of important advantages of this test is that the alternative hypothesis is not just a general rejection of the null like in the commonly used LM panel cointegration test of McCoskey and Kao (1998), but allows ϕ_i to differ across individual series.

6. Empirical Results

In general, our panel unit root estimations confirm that our variables are non-stationary or contain unit root I(1) (Table 1). A panel version of KPSS almost in all cases rejected the hypothesis of the stationarity of variables— current account, real exchange rate and terms of trade except for the current account where a time dummy was not included. The IPS test did not reject the null hypothesis of non-stationarity in cases of real exchange rate and terms of trade; however, the test rejected the hypothesis of unit root presence in the current account variable in the case with the constant and in the case where a time dummy was included. The LLC test in all cases confirmed that our variables contain a unit root; however, in the case where the trend dummy was not included the current account variable was found to be stationary. We may conclude that the inclusion of time dummies is important for all considered variables because the time period that we consider is full of continuous changes in taken economies. Therefore, based on the results of alternative panel unit root tests we have evidence to assume that all our considered variables contain a unit root.

The results of the LM unit root tests are reported in Tables 2, 3 and 4. In the case when structural breaks are not allowed, the null hypothesis of the unit root is accepted almost in all cases except for Brazil and for LM statistics for the panel test of the current account variable when trend was included.

However, in cases of unit root tests with structural breaks (Table 3 and 4), almost in all cases the unit root hypothesis was rejected as for univariate tests as well as for panel tests. It can be seen from the results that stationarity becomes apparent once structural shifts are allowed in the model. Table 3 reports the dates of the structural shifts found by the methodology of Lee and Strazcich (2004) while Table 4 presents the dates of two structural shifts for every country found by the methodology of Lee and Strazcich (2003). Thus, all variables appeared to have unit root in the absence of structural shifts and at the same time all

variables were found to be stationary in the case of structural shifts accommodation. In order to estimate cointegration relationships it is necessary that all variables be non-stationary. Given the results from the LLC, IPS, PKPSS and SP unit root tests we consider CA, RER and TOT variables as stationary; therefore it can be proceed with cointegration.⁴

The LM unit root test of individual series in Table 2 provided evidence of the unit root presence in selected variables; therefore Hansen's (1992) stability test can be applied. The stability test produces three test statistics: supF, meanF and Lc. The supF statistic tests for the null hypothesis of cointegration with no structural shift in the parameter vector against the alternative hypothesis of cointegration in the presence of sudden structural shifts. The meanF and Lc statistics test for a cointegration with constant parameters against alternative hypothesis of gradual variance in parameters, which is considered no cointegration. Particularly, the meanF statistic is used to capture the overall stability of the model.

The results of the stability tests are presented in Table 5. Almost in all cases supF statistics reject the null hypothesis of the stability of model parameters indicating the presence of structural change in parameters, except for the cases of India and Turkey, where test statistic is unable to reject the hypothesis of cointegration without structural shifts. The meanF statistic, in all cases except India and Turkey, rejects the hypothesis of cointegration in favor of the instability of the overall model in the considered countries. The Lc statistic is unable, however, to reject the hypothesis of constant parameters in most cases except for Brazil and India, where the null is rejected only at the 10% significance level. The results of the stability test do not provide clear evidence of the changes in the parameters of cointegration following the mixed results of meanF and Lc statistics. However, we found evidence of the presence of sudden structural shifts in the model in the cases of Argentina, Brazil, China, Korea and Indonesia.

On the basis of the unit root results, we apply the panel cointegration tests reviewed in Section 5. To calculate the McCoskey and Kao (1998) LM* panel statistic, we specify the following DOLS regression:

$$CA_{t,i} = \alpha_i + \beta_0 P'_{t-1} + \beta_1 Q_{t-1} + \sum_{j=-k}^{k_i} \Delta P'_{i,t-1} + \sum_{j=-k}^{k_i} \Delta Q_{t-1} + u_{t,i} \quad (18)$$

$i=1,\dots,N$, k_i =leads and lags of terms of trade, P' , and real exchange rate, Q . The DOLS estimator, used for calculating the McCoskey and Kao (1998) LM* statistic, includes leads

⁴ See for examples Beyer et al. (2009), Bagnai (2006), Daly and Siddiki (2008) where cointegration relationship between variables were applied to non-stationary variables according to conventional unit root tests without accomodation of structural breaks.

and lags of the changes in the explanatory variables. However, the DOLS method does not always produce residuals free from autocorrelation. Hence, Stock and Watson (1993) proposed a generalised dynamic least squares (DGLS) estimator that includes leads, lags and an autoregressive error process, thus encompassing Saikkonen's (1991) DOLS estimator.

We included one lead and lag and first order autoregressive term. Decisions regarding to the selection of leads and lags, and the autoregressive terms are selected according to AIC. According to the results of the LM panel test presented in Table 6, the null hypothesis of cointegration in the panel is rejected. However, there are many studies where conventional cointegration tests provide evidence of no cointegration, while cointegration tests with the allocation of structural breaks demonstrated the existence of cointegration among variables (see, for example, Beyer et al., 2009; Basher and Westerlund, 2009)

Table 7 presents the results of the panel cointegration test allowing for multiple structural shifts. In the test implementation, a maximum of five breaks were allowed. Panel A demonstrates the results of the test in which structural shifts are allowed in constant, while Panel B illustrates test results where structural shifts are allowed in both constant and trend of the regression. The results of Panel A and B do not differ significantly. The test was able to detect five significant structural shifts for Korea and four structural shifts for Turkey in both cases in the presence of constant and in the case with trend inclusion. For India and Indonesia three breaks were detected, while for Brazil and China only two breaks were detected in the model with constant and three breaks were detected in the case of trend inclusion. The location of break points mainly concentrated in the first half of the 1980s, which could have been a repercussion of the oil price shock of 1979. Another part of the breaks is concentrated around the end of the 1990s and the beginning of the 2000s, which coincided with the Asian financial crisis of 1997, when countries such as China, Korea and Indonesia were deeply affected. The economic crisis of Argentina in 1999-2002 with a decline in GDP is closely reflected by break points in Argentina as well. From the results the reflection of the economic crisis in Turkey in 2000-2001, the devaluation of the Turkish lira, the dramatic increase in inflation and the decline in GDP can be seen.

Statistics for panel LM test are 6.43 and 19.55 when we allow for a shift in a level and in a level and in trend, respectively. Thus, we cannot accept the null hypothesis of cointegration in both cases. In Table 5 the results of the stability test in which strong evidence of structural shifts presence in series in Argentina, Brazil, China, Korea and Indonesia were found were presented. For the cases of India and Turkey, we found mixed results for structural break presence in regressions. Therefore in order to avoid spurious results the panel test was applied

to the panel excluding series of India and Turkey, one by one and then excluding both of them.⁵ The detected break dates for the rest of series did not show any significant difference and LM statistics did not provide any evidence to support the null of cointegration in the panel.

7. Summary and Conclusion

The purpose of this paper is to contribute to the literature on the panel cointegration analysis of current account and its determinants applying panel cointegration analysis allowing for structural shifts. There is a limited number of studies on the cointegration estimations of current account regressions with structural shifts accommodation, see, for example, Bagnai (2006), Leachman and Francis (2002). The estimations of our tests provided evidence of the unit root presence in all our series using conventional panel unit root tests. The panel cointegration test of MacCoskey and Kao (1998) did not provide any evidence of cointegration relationship among variables. Therefore, in order to use the panel cointegration test allowing for structural shifts, it was necessary to estimate whether our series have structural breaks. For this purpose, Hansen's (1992) stability test was applied, where strong evidence was found for instability in the regressions of Argentina, Brazil, China, Korea and Indonesia, and mixed evidence was found for the instability in the India and Turkey models. Having evidence of the instability of the series, the panel cointegration test of Westerlund (2006) was applied with the accommodation of unknown multiple breaks. However, the results of this test did not supply any evidence of cointegration in the panel. As stability estimations of India and Turkey models provided mixed results of their parameter instability, panel cointegration test was run excluding these two countries, however results did not differ from full panel estimations, and did not illustrate any evidence of cointegration in the panel with the excluded cases of India and Turkey. Our study illustrated that there are no long run relationships between the Current Account, the Relative Prices and the Terms of Trade variables in the considered countries. However, the cointegration tests were applied on the basis of the results of conventional unit root tests, which indicated the non-stationarity of the considered variables. However, unit root tests allowing for structural breaks indicated that all variables in the panel are stationary. Therefore, the accommodation of structural breaks in unit root tests demonstrated the presence of structural breaks and absence of the unit root in the

⁵ The results are not provided in the paper for the purpose of space saving and available upon the request.

panel, indicating that there are no long-run relationships between the current account and its chosen determinants in the panel.

8. References

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9. Appendix: Data

We used a panel analysis. The selected countries were Argentina, Brazil, China, Korea, India, Indonesia, and Turkey. We used annual data between 1982 and 2008. The series for the current account variable, CA, was constructed for each country by subtracting the log of consumption from the log of net output. The net output was constructed by subtracting investment and government purchases from the GDP.

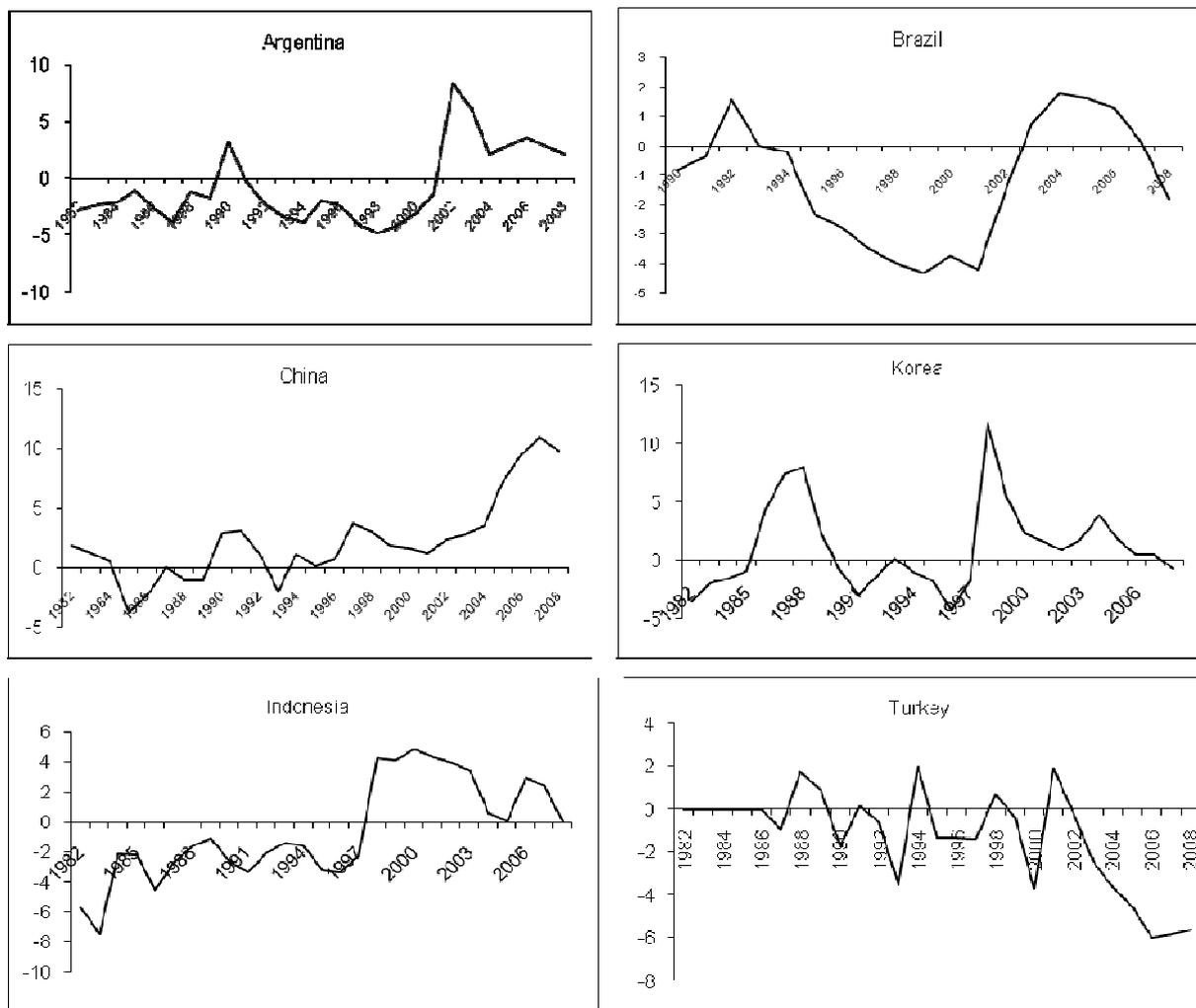
The series for the terms of trade, TOT, was the relative price of exports in terms of imports. The series for the real exchange rate, RER, was constructed by using the formula: $RER = XR(CPI_{US} / CPI_i)$, where XR is the currency in country i per dollar, CPI_{US} is the consumer price index for the US and CPI_i is the consumer price index in country i .

Finally, all series were obtained from the IMF, International Financial Statistics. They are all in log levels.

Figures and Tables

Figures

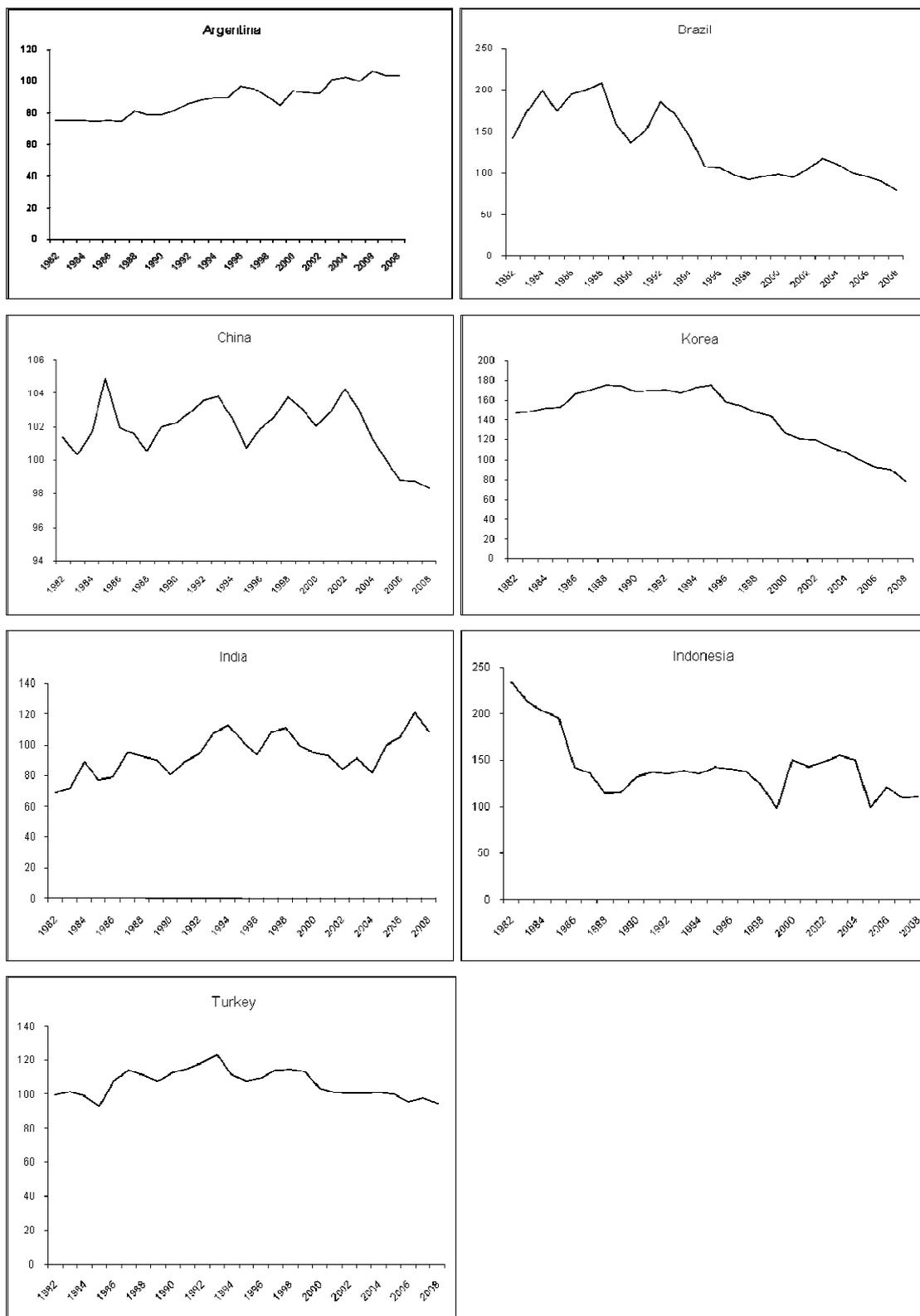
Figure 1. Current Account as a fraction of GDP.



Source: International Financial Statistics, IMF, 1982-2008⁶.

Figure 2. Terms of trade.

⁶ There is no available data for India.



Source: International Financial Statistics, IMF, 1982-2008.

Tables

Table 1. Panel Unit Root Tests

Test	<i>ca</i>	<i>rer</i>	<i>tot</i>
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LLC ^a	<i>ct</i>	-1.06	0.59	-1.49
	S	I(1)	I(1)	I(1)
	<i>c</i>	-2.34*	-1.63	2.19
IPS ^a	<i>ct</i>	-1.92*	1.45	-0.60
	S	I(0)	I(1)	I(1)
	<i>c</i>	-3.61*	-1.49	1.31
PKPSS ^β	<i>ct</i>	1.65*	6.12*	4.67*
	S	I(1)	I(1)	I(1)
	<i>c</i>	-0.04	6.58*	7.45*
	S	I(0)	I(1)	I(1)

* indicate significance at 5% significance level.

MacKinnon (1996) one-sided p-values are used.

^a null of non-stationarity (unit root)

^β null of stationarity

c and *ct* denote, respectively, that there is a constant and a constant with time dummy in the regression.

Table 2. Schmidt and Phillips (1992) LM unit root test

Country	CA		RER		TOT	
	<i>c</i>	<i>ct</i>	<i>c</i>	<i>ct</i>	<i>c</i>	<i>ct</i>
Argentina	-2.19	-2.41	-1.67	-2.64	-2.98	-2.49
Brazil	-3.27	-3.75*	-1.71	-1.71	-1.87	-1.98
China	-2.67	-2.53	-1.99	-1.53	-2.21	-1.45
Korea	-3.10	-3.89	-2.59	-1.79	-0.96	-1.29
India	-1.77	-2.00	-1.61	-1.58	-2.38	-2.34
Indonesia	-2.81	-3.41	-2.45	-1.67	-2.07	-1.86
Turkey	-2.36	-2.65	-2.06	-1.12	-1.82	-2.54
Panel LM statistic	-2.58	-4.05*	-0.24	1.12	-0.42	-0.09

* denotes significance at the 1% level

Table 3. Panel Unit Root Test with one structural break

Country	CA			RER			TOT		
	<i>LM</i>	<i>Break</i>	<i>Lag</i>	<i>LM</i>	<i>Break</i>	<i>Lag</i>	<i>LM</i>	<i>Break</i>	<i>Lag</i>
Argentina	-5.83*	2000	4	-3.52	1991	2	-3.87*	1989	1
Brazil	-4.83*	1994	0	-5.41*	1991	4	-2.99	1994	4
China	-3.32	1996	3	-4.67*	1990	2	-4.23*	1995	1
Korea	-4.13*	1992	1	-4.15*	1992	3	-3.77*	1998	3
India	-4.26*	1995	3	-4.65*	1992	3	-4.12*	2000	2
Indonesia	-4.13*	1994	0	-3.75*	1992	3	-4.25*	1992	3
Turkey	-4.31*	1996	4	-2.67	1990	0	-4.48*	1998	4
MinLM	-4.31*	1996	4	-2.67	1990	0	-4.48*	1998	4
LM	-10.21*			-9.13*			-8.40*		

statistic									
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* denotes significance at the 1% level

Table 4. Panel Unit Root Test with two structural break

Country	CA			RER			TOT		
	LM	Break	Lag	LM	Break	Lag	LM	Break	Lag
Argentina	-7.42*	1993,2000	4	-9.18*	1991,2001	4	-4.27*	1995,2001	1
Brazil	-5.73*	1990,1994	1	-5.74*	1991,2000	4	-3.83*	1994,1999	4
China	-4.60*	1992,1996	4	-5.55*	1988,1992	1	-6.08*	1995,2001	1
Korea	-5.66*	1988,1999	1	-5.40*	1988,2001	1	-5.31*	1995,1998	3
India	-7.21*	1988,1992	4	-6.27*	1992,2001	4	-5.21*	1999,2003	0
Indonesia	-5.07*	1990,1996	3	-4.95*	1991,1997	4	-7.37*	1993,1998	4
Turkey	-6.82*	1990,2002	0	-5.69*	1988,2000	1	-5.77*	1993,1998	4
MinLM	-6.82*	1990,2002	0	-5.69*	1988,2000	1	-5.77*	1993,1998	4
LM statistic	-16.98*			-17.10*			-14.27*		

* denotes significance at the 1% level

Table 5. Stability Tests in Cointegrated Relations

Country	SupF		MeanF		Lc	
	test	p-value	Test	p-value	test	p-value
Argentina	2178	0.01	123.36	0.01	0.15	0.20
Brazil	2084	0.01	456.19	0.01	0.51	0.09
China	84.58	0.01	8.13	0.01	0.21	0.20
Korea	60.69	0.01	8.29	0.01	0.38	0.20
India	8.27	0.20	3.60	0.20	0.53	0.08
Indonesia	970.01	0.01	485.81	0.01	0.19	0.20
Turkey	5.54	0.20	1.91	0.20	0.15	0.20

Table 6. McCoskey and Kao (1998) Panel Cointegration Test

Test	Statistics
LM*	14.43

Note: (a) The McCoskey and Kao (1998) LM* statistic is one-sided with a critical value of 1.64. Therefore large values ($LM^* > 1.64$) suggest rejection of the null hypothesis. The mean and variance used to calculate the LM* statistic are respectively 0.0850 and 0.0055 (MacCoskey and Kao, 1998, Table 2).

Table 7. Estimated Structural Breaks Using the Approach of Westerlund (2006).

Panel A breaks in constant						
Country	Breaks	Date				
Argentina	3	1983	1989	2003		
Brazil	2	1982	1991			
China	2	1982	1988			
Korea	5	1985	1989	1993	1997	2002
India	3	1983	1985	1993		
Indonesia	3	1983	1993	1999		

Turkey	4	1984	1990	1997	2003	
Lm	6.425					
Panel B breaks in constant and trend						
Country	Breaks	Date				
Argentina	3	1983	1989	2003		
Brazil	3	1983	1985	1991		
China	3	1983	1988	1998		
Korea	5	1985	1987	1991	1997	2002
India	3	1983	1987	1997		
Indonesia	3	1983	1992	1999		
Turkey	4	1984	1990	2000	2004	
Lm	19.553					