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Robustness of the Balassa-Samuelson effect: evidence from developing and emerging economies

Florian MORVILLIER*

Abstract

This paper aims at investigating the robustness of the Balassa-Samuelson (BS) effect to alternative proxies for a panel of 38 developing and emerging economies over the period 1980-2016. We examine the internal and external versions of the BS hypothesis using a total of five different measures. Relying on the Cross Sectional-Distributed Lag (CS-DL) approach, we show that the internal version of the BS hypothesis holds only if the labor productivity differential between the tradable and non-tradable sectors is used rather than the Gross Domestic Product Per worker. We also find evidence of a positive and robust effect of the relative price of the non-traded to traded goods on the real exchange rate. Overall, our findings highlight that while the verification of the internal version of the BS effect depends on the proxy considered for productivity, the validity of the external version is a general and robust result.

JEL codes: F31, F41

Keywords: Balassa–Samuelson effect, real exchange rate, relative prices.

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1 Introduction

Sizeable currency misalignments are shown to be harmful for growth (see Aguire and Calderon, 2005; Gala, 2007). This issue is particularly critical for Emerging Markets and Developing Countries (EMDEs), which can rely on a competitive real exchange rate (RER) policies to promote their economic development. A proper assessment of their equilibrium exchange rates is thus more important as commonly thought as it allows an appropriate monitoring of the evolution of their currency misalignments. Such exercise is, however, affected by significant measurement issues surrounding the consideration of the Balassa-Samuelson (BS, hereafter) effect, the main driving force behind the long-run dynamic of the RER (Couharde et al., 2018). According to the BS effect (Balassa, 1964; Samuelson, 1964) a relatively larger productivity growth in the domestic traded goods compared to the non-traded goods sector causes wage increases in the tradable goods sector. Assuming wage equalization across sectors it leads to a rise in the relative price of non-traded to traded goods, pushing up the general price level and leading therefore to a RER appreciation. In fact, the BS effect relies on two components (Canzoneri et al., 1999): internal and external components (Égert et al., 2003; García-Solanes and Torrejón-Flores, 2009). The internal version is the positive relationship between the productivity differential and the relative price of non-tradables, while according to the external version higher relative price of non-tradables should appreciate the RER. The problem is that the investigations of the BS effect suffer from two main measurement issues. On the one hand, the productivity differential and the relative price of non-tradables to tradables have to be carefully appraised as they are the two key ingredients of the BS effect. On the other hand, an adequate econometric framework is also required to properly estimate the extent of the BS effect, especially in the context of panel data. However, despite their great importance, these different issues have not yet been addressed in the literature examining EMDEs. This paper aims at filling this gap.

We provide an in-depth investigation of the BS hypothesis for EMDEs by simultaneously addressing the two aforementioned issues on a sample of 38 developing and emerging economies over the period 1980-2016. As the relative price of non-traded to traded goods stands at the heart of the BS effect, we begin our empirical investigation by assessing its determinants. Although some evidence have been provided for advanced economies (see Lane and Milesi-Ferretti, 2002; Galstyan and Lane, 2009; Galstyan and Velic, 2018), to the best of our knowledge, this issue has not yet been addressed for EMDEs. It is however of primary importance as the relative price of non-tradables accounts for a significant amount of the exchange rate variation (Burstein et al., 2006; Ouyang and Rajan, 2013; Cheung et al, 2015). We then assess the internal and external versions using a battery of proxies. Specifically, we address the measurement uncertainty previously discussed by employing two alternative measures for the productivity differential, and three for the relative price of non-tradables. On the one hand, we use the Gross Domestic Product Per Worker (GDP PW) and the labor productivity differential between tradable and non-tradable sectors as a measure of productivity differential. On the other hand, the relative price of non-tradables to tradables is appraised employing three and six sectors' value-added deflators. For the sake of robustness, fixed and time-varying weights are used in the aggregation process of the sectors into non-tradables and tradables. Moreover, for comparison purposes with existing studies, we also examine the relevance of the GDP Per Capita (GDP PC) based-proxy for the external version of the BS hypothesis.

Thus, we go further than previous examinations focusing on EMDEs (Bahmani-Oskooee and Nasir, 2004; Ricci et al., 2013; Wang et al., 2016) as we systematically investigate the robustness of the internal and external versions of the BS effect to the proxies. Indeed, the existing literature dealing with EMDEs considers a unique measure in their investigation of the BS hypothesis such as GDP PC (Wang et al., 2016), GDP PW (Bahmani-Oskooee and Nasir, 2004) and productivity of tradables relative to tradables (Ricci et al., 2013). Moreover, to the best of our knowledge, we are also the first to examine the robustness of the BS effect to alternative weighting scheme in the aggregation process of the sectors into non-tradables and tradables. Furthermore, from a methodological viewpoint, panel data investigations of the BS hypothesis extensively rely on Dynamic Ordinary Least Square (DOLS) and/or Fully Modified OLS (FMOLS) estimators which assume cross-sectional independence (see Choudhri and Kahn, 2004; Ricci et al., 2013; Wang et al., 2016). This issue is particularly important because the ignorance of the cross-sectional dependence in the data can lead to misleading inference and even inconsistent estimation (see Pesaran, 2015; Reese and Westerlund, 2016; among others). To tackle this issue, we rely on the Cross-Sectionally augmented Distributed Lag (CS-DL) estimator recently developed by Chudik et al. (2016). This estimator is particularly suitable as it allows us to estimate long-run effects in panel data models with cross-sectionally dependent errors and presents good small sample performance. To sum up, we contribute to the existing literature in three ways. First, we provide an original analysis of the relative price of non-tradables to tradables' determinants for EMDEs. Second, we investigate the robustness of the BS effect using a large battery of measures. Finally, we make use of a new methodology allowing us to tackle the issue of cross-sectional dependence.

Our estimations provide evidence that government consumption expenditures are a robust driver of the relative price of non-tradables, which is also affected -but to a lesser extent- by the trade balance. Our results show quite strong support in favor of the internal version of the BS effect if we use the labor productivity differential between the tradable and non-tradable sectors: a 1% increase in the labor productivity differential raises the relative price of non-traded to traded goods by about 0.15% to 0.35%. Furthermore, our findings strongly support the external version of the BS hypothesis. Indeed, no matter the sectoral value-added deflators or the aggregation schemes examined, a significantly positive effect on the Real Effective Exchange Rate (REER) is obtained. At least, a 1% increase in the price deflator leads to an appreciation of the currency going from 0.18% to a maximum of 0.32%. We also show that GDP PC is clearly inadequate to detect the existence of BS hypothesis for our sample of countries.

Our paper is organized as follows. In Section 2, we discuss the conceptual issues associated with the BS effect. Section 3 discusses how to measure the BS effect. Section 4 provides an overview of the empirical evidence for the BS effect. Section 5 presents the data and methodology. Section 6 reports and discusses our empirical results. Finally, Section 7 concludes the paper.

2 Conceptual issues

This section is devoted to the presentation of the conceptual issues associated with the Balassa-Samuelson effect. Section 2.1 presents the Balassa-Samuelson theoretical framework. Section 2.2 discusses the recent theoretical developments of the BS framework.

2.1 The Balassa-Samuelson theoretical framework

In this section, we develop the Balassa-Samuelson theoretical framework, which is a useful benchmark for our empirical investigation.

2.1.1 The intercountry relative price of non-tradable in terms of tradable goods

One way of giving a very simplified representation of the BS effect is to take the example of two open economies—a catching-up economy and an advanced country—that produce two types of goods: tradable goods (T) and non-tradable goods (NT). If we suppose that the price index (p, expressed in logs) in each economy is a geometric average of traded and non-traded goods prices, log-differentiating the expressions for prices yields:

- In the catching-up economy:

$$\overset{\bullet}{p} = \gamma \overset{\bullet}{p}_{NT} + (1-\gamma)\overset{\bullet}{p}_T \tag{1}$$

- In the more advanced country:

$$\overset{\bullet}{p}^* = \gamma \overset{\bullet}{p}^*_{NT} + (1-\gamma)\overset{\bullet}{p}^*_T \tag{2}$$

where foreign variables (i.e., variables related to the trading partners) are flagged with a star and the • denotes the rate of change. γ is the share of non-tradable goods in the consumption basket, assumed to be the same in the two economies for simplicity.

The real exchange rate (q, expressed in logs) between the two economies is, by definition, the nominal exchange rate adjusted by price levels:

$$\overset{\bullet}{q} = \overset{\bullet}{s} + \overset{\bullet}{p} - \overset{\bullet}{p}^* \tag{3}$$

where s is the exchange rate defined in units of the currency of the advanced country per units of the currency of the catching-up economy.

Then, substituting (1) and (2) into (3) gives:

$$\overset{\bullet}{q} = \overset{\bullet}{q}_T + \gamma[(\overset{\bullet}{p}_{NT} - \overset{\bullet}{p}_T) - (\overset{\bullet}{p}^*_{NT} - \overset{\bullet}{p}^*_T)] \tag{4}$$

Equation (4) indicates that the appreciation of the real exchange rate in the catching-up economy can be explained by two factors: (i) the increase in the relative price of tradable goods, $\overset{\bullet}{q}_T$, and (ii) the increase in the relative price of non-tradable in terms of tradable goods. If Purchasing Power Parity (PPP) holds only for tradable goods, then the appreciation of the real exchange rate in the catching-up economy will stem from faster rise in the prices of non-tradables relative to tradables compared to the advanced economy. Equation (4) corresponds to the external version of the BS effect.

The faster rise in the relative price of non-tradable in terms of tradable goods may come from a variety of factors. For Balassa (1964) and Samuelson (1964), it results from higher relative productivity gains in the tradable sector in the catching-up economy.

2.1.2 The intercountry productivity differential between the tradable and the non-tradable sectors

Assuming that the production functions in the two sectors are of Cobb-Douglas type and the same in both countries, as in De Gregorio et al. (1994) and Rogoff (1996), we have:

$$Y_T = A_T L_T^{\alpha_T} K_T^{1-\alpha_T} \tag{5}$$

$$Y_{NT} = A_{NT} L_{NT}^{\alpha_{NT}} K_{NT}^{1-\alpha_{NT}} \tag{6}$$

where Y designates output, L labor and K capital. α represents the share of labor in the sectors' value-added, and A denotes the total productivity of factors. Under perfect competition, prices in each sector are thus given by:

$$P_T = \frac{1}{A_T} W^{\alpha_T} R^{1-\alpha_T} \alpha_T^{-\alpha_T} (1-\alpha_T)^{-(1-\alpha_T)} \tag{7}$$

$$P_{NT} = \frac{1}{A_{NT}} W^{\alpha_{NT}} R^{1-\alpha_{NT}} \alpha_{NT}^{-\alpha_{NT}} (1-\alpha_{NT})^{-(1-\alpha_{NT})} \tag{8}$$

where W is the unit cost of labor and R the rate of return on capital. If we consider the case of a small open economy with perfect capital mobility and P_T as the numeraire, then PPP in the tradable goods sector ensures that the rate of return in tradables (R) is equal to its world value. Log-differentiating the expressions for prices yields:

$$\dot{p}_{NT} = -\dot{A}_{NT} + \alpha_{NT}\dot{w} \tag{9}$$

$$\dot{p}_T = -\dot{A}_T + \alpha_T \dot{w} = 0 \tag{10}$$

where variables in lowercase are expressed in logarithmic terms. Solving for the difference, the increase in the relative price of non-tradable in terms of tradable goods can be written in the catching-up economy as:

$$\dot{p}_{NT} - \dot{p}_T = \frac{\alpha_{NT}}{\alpha_T}\dot{A}_T - \dot{A}_{NT} \tag{11}$$

Equation (11) corresponds to the internal version of the BS hypothesis. A positive productivity differential between the tradable and non-tradable sectors is expected to rise the relative non-tradable price. Assuming for simplicity that $\frac{\alpha_{NT}}{\alpha_T}$ is the same in the two economies, we have:

$$\overset{\bullet}{p}_{NT}^{\ *} - \overset{\bullet}{p}_{T}^{\ *} = \frac{\alpha_{NT}}{\alpha_T}\overset{\bullet}{A}{}_T^* - \overset{\bullet}{A}{}_{NT}^* \tag{12}$$

Substituting equations (11) and (12) into Equation (4) yields to the following expression for the real exchange rate:

$$\overset{\bullet}{q} = \overset{\bullet}{q}_T + \gamma\left[\frac{\alpha_{NT}}{\alpha_T}(\overset{\bullet}{A}_T - \overset{\bullet}{A}{}_T^*) - (\overset{\bullet}{A}_{NT} - \overset{\bullet}{A}{}_{NT}^*)\right] \tag{13}$$

Then, if PPP holds only for tradable goods, the appreciation of the real exchange rate in the catching-up economy will stem from faster relative productivity growth in the tradable goods sector compared to that of the advanced country.

2.2 Theoretical developments

2.2.1 BS effect and the new trade theory

One strand of the theoretical literature makes use of the development of the "new trade theory" to enhance the basic BS framework through the endogenization of three key ingredients:

- goods' tradability (Bergin et al., 2006),
- spatial location of firms (Mejean, 2008),
- inclusion of Terms Of Trade (TOT) adjustment (Choudhri and Schembi, 2010).

As shown by Bergin et al. (2006), one possible way to capture the evolving pattern of the BS effect is to endogenize goods' tradability. To do so, they propose an updated BS framework constituted of a continuum of goods differentiated by their productivity. In their model, goods' tradability results from firms' choice which decide to trade a good only if it is sufficiently productive to support the associated trade costs. The time varying pattern of the BS effect is captured by the existence of two additional channels associated with a productivity rise in traded goods. On the one hand, following a productivity rise in the traded sector, wages in the domestic sector increase and some goods become non-traded because they are no longer enough beneficial to be exportable. It reinforces thus the BS effect through the increase of the share of non-traded goods in the domestic basket pushing upwards the cross-country effect of wage rate. On the other hand, the BS effect is mitigated by the reduction in the traded goods' share which lowers the relative productivity effect. While this model constraints traded goods' producer firms to produce either locally or in the foreign economy, Mejean (2008) builds a framework where the spatial location of these firms is endogenously determined. In her model, a rise in the domestic productivity of the tradable sector affects the RER through three different channels: the BS effect, a "terms-of-labor" effect and an expenditure switching effect. The "terms-of-labor" effect reinforces the BS effect. To take advantage of the productivity improvement in the domestic tradable sector, firms decide to produce locally exerting upward pressures on domestic wages, reinforcing the BS effect. The “expenditure switching effect” acts in opposition with the BS effect. Following the location choice of firms to go in the local market, the share of domestically produced goods in consumption increases, reducing thus the non-tradable price because consumers are able to save on trade costs as there is a higher number of domestic firms. To investigate which effect dominates, Mejean (2008) performs a panel cointegration analysis. She finds that the "terms-of-labor" effect dominates the "expenditure switching effect". Choudhri and Schembi (2010) also consider the role played by firms entries and its potential effect on TOT.[1] Indeed, to understand the failure of the BS effect, a two-country model with differentiated traded and non-traded goods is built where the role of TOT is explicitly modelised. An improvement in traded productivity can either appreciates or depreciates the RER because of the ambiguous effect on TOT. For example, following a rise in the home traded productivity firms entry in this sector, lower the price of the home traded good. Hence, two antagonist effects are at play: (i) the RER appreciates via the relative price of non-traded goods, but (ii) depreciates *via* the TOT channel.

2.2.2 BS effect and imperfect competition

The traditional assumption of perfect competition in the BS framework is relaxed by Coto-Martinez and Reboredo (2014). In their amended BS model, the price of non-traded and traded goods is no longer determined only by marginal costs but also by firms' mark-ups. As firms benefit from market power, they are able to fix a price above their marginal costs. Their framework shows that mark-up variations produce changes in prices provided that mark-up movements in one sector are not offset by those prevailing in the other sector. Their model is empirically verified on a sample of 12 OECD countries over the period 1970-2006. Using Mean Group (MG) and Pooled Mean Group (PMG) estimators, they confirm their theoretical prediction as a rise in mark-up differentials between tradable and non-tradable goods positively affects the relative price of non-traded goods. More specifically, a 1% increase in the mark-up differentials leads to a decrease of relative price by about 0.81%. While imperfect competition is assumed for both sectors by Coto-Martinez and Reboredo (2014), Bénassy-Quéré and Coulibaly (2014) assume monopolistic competition only in the non-tradable sector. They investigate the role played by product market regulation on the RER dynamic of some European economies. Consistent with Coto-Martinez and Reboredo's model (2014), they show that market power affects the RER. This theoretical prediction is confirmed by an empirical analysis for 12 European countries over the period 1985-2006, where product market regulation is shown to positively affect the RER.

2.2.3 BS effect and the labor market

The transcription of Balassa (1964) and Samuelson (1964) ideas in a modern framework relies on the following assumptions regarding the labor factor: (i) perfect labor mobility across the tradable and non-tradable sectors, (ii)

[1] In their model, TOT are computed as the ratio between the price of the exported and imported varieties. They also assume that TOT have a positive effect on the RER, meaning that the income effect dominates the substitution effect.

homogeneity of the labor factor and (iii) perfect competition. In front of the inability of the BS hypothesis to fully explain the RER dynamic, these assumptions have been continuously relaxed through the following ways: inclusion of firing and hiring costs (Shen and Xu, 2011), consideration of heterogeneity in the labor factor (Doan and Gente, 2014), inclusion of imperfect labor mobility across sectors (Cardi and Restout, 2015), labor wedge[2] (Berka et al., 2018) and sector varying labor market unionization rate (Berka and Steenkamp, 2018). While the costs associated to job search are eluded from the basic BS framework, Sheng and Xu (2011) extend it to an environment with search and unemployment where firms and workers face firing and hiring costs. Hence, obviously the filling up of a vacant job takes time for firms. The traditional BS framework emerges as a special case of their model where hiring and firing costs are equal to zero. Such labor market frictions influence the BS effect as part of the relative productivity rise will be used by firms to cover these frictional costs. Hence, wages increase is lower and the BS effect is mitigated. To examine the empirical relevance of their theoretical framework, Shen and Xu (2011) employ a dummy variable measuring "relative labor market inefficiency"[3] which is interacted with GDP PC. They show that economies with high labor market frictions present lower effect for productivity, confirming thus their framework. More recently, Berka et al. (2018) build an amended BS framework where labor market distortions are taken into account. In their new-Keynesian Dynamic Stochastic General Equilibrium (NK DSGE) 2 sectors framework, it is shown that a labor supply shock affects the RER through the TOT channel. Indeed, as a labor supply shock raises the relative wages, export prices increase leading thus to a RER appreciation. In their empirical investigation on the eurozone, they find evidence of a BS effect if and only if they control for labor wedge through Unit Labor Cost (ULC). The positive effect expected for the labor wedge proxy is confirmed by the significantly positive sign for the ULC.[4] While in the previous model, the source of the labor wedge is not explicit, Berka and Steenkamp (2018) amend the specification proposed by Berka et al. (2018) with varying labor market unionization rate. Their amended BS framework is supported by their estimation as an increase in the sectoral wage mark-ups due to labor market institution appreciates the RER.[5]

Cardi and Restout (2015) also relax the hypothesis of perfect labor mobility across sectors in a two sector small open economy model allowing for a limited substitutability in hours worked. Constructing traded and non-traded productivity from the Solow residuals of a production function with capital stock and employment, they provide evidence contrary to the BS effect. These findings are reconciled with this hypothesis showing that countries with higher intersectoral labor reallocation experience higher relative price increase and lower wage decrease. While the assumption of labor homogeneity is common, Doan and Gente (2014) relax this hypothesis in a two-sector specific model where only capital is mobile across sectors. They assume that the traded sector is intensive in skilled workers, while the non-traded sector is unskilled labor intensive. In their model, a rise in traded productivity entails

[2]Berka et al. (2018) defined the labor wedge as the difference between the marginal product of hours worked in production and the marginal rate of substitution between labor and consumption.

[3]The labor market inefficiency variable is measured as hiring and firing costs in number of weeks of salary. The dummy is equal to one if a country has higher hiring and firing costs than those of the UK which is 1 by default.

[4]It is worth mentioning that the positive effect of ULC has already been highlighted by Mejean (2008).

[5]They use different proxies of structural labor market differences: measures of concentration of unions at aggregate and sectoral levels, union density rate and replacement rate.

two effects: (i) a decrease in non-tradable goods production, (ii) a rise in the skilled wages and reduction in the unskilled wages since the traded sector is labor intensive. In their model, a reversed BS effect can occur if the share of unskilled workers wage in total wages exceeds the unskilled labor share diminishing the global demand. Thus, countries exporting skilled labor-intensive goods can face a RER appreciation following a traded productivity shock depending on the share of skilled to unskilled workers in the economy. Considering skilled workers as the proportion of persons having completed tertiary education, they can confirm their model prediction through a cointegration analysis.

2.2.4 Introduction of the distribution sector

The tradable and non-tradable sectors obviously constitute an aggregate of different sectors. One of them is the distribution sector which has received a special interest in the literature. Traditionally, the wholesale and retail trade has been considered as a non-tradable sector. Devereux (1999) was the first to endogenize the distribution sector in the BS framework. He shows that an increase in the productivity of the traded sector leads to a rise in consumption; it expands the distribution sector's size and lowers the price for goods distribution. The traded goods price faced by the consumers is thus lower and the RER can depreciate. Shortly after Devereux (1999), MacDonald and Ricci (2005) develop a model where the effects of a rise in productivity in the distribution sector are explicitly modeled. In their framework, two antagonist effects are at play. On the one hand, a rise in productivity in the distribution sector is associated with lower price for consumers, depreciating the RER. On the other hand, the RER appreciates as higher productivity in the distribution sector lowers the price of tradables by lowering the cost of distributing intermediate inputs. Given the absence of clear-cut predictions from their model, they investigate which effect dominates through an empirical analysis. To this aim, they rely on a sample of 9 OECD countries over the period 1970-1991. They show that productivity in the distribution sector relative to the US is positively correlated with the RER. In other words, the use of services from the distribution sector to deliver intermediate goods in the production of tradables has a larger impact on the RER than the use of distribution services to deliver final goods to consumers.

2.2.5 BS effect and OverLapping Generation (OLG) model

In the face of the failure of the BS effect for some developing economies, several amendments to the benchmark model have been proposed. One common feature of these studies is to consider how countries' macroeconomic characteristics can affect the effect of the differential productivity on the RER. Examples of characteristics examined are: (i) constraints on capital inflows (Gente, 2006; Christopolous et al., 2012), and (ii) saving rate and population growth (Doan and Gente, 2013). An efficient way to include these different factors is to rely on OverLapping Generation (OLG) models. Building on an OLG model, Gente (2006) highlights the relevance of these different macroeconomic variables for the RER dynamic. Indeed, she shows the dependency of the RER to the following variables: the rate of time preference, the age dependency ratio and external constraints. More precisely, to shed some light on the RER determinants of Asian economies, she develops a two sector small open economy OLG model where the country faces a constraint on capital inflows. If the economy is constrained, the RER is determined by both demand and supply shocks. The theoretical effect of higher traded productivity on the RER is ambiguous. The mechanism proposed by Gente (2006) is the following: higher traded productivity affects the return of each sector and increases the interest rate. Then, we observe a lower aggregate demand leading to a RER depreciation. This capital inflows constraint creates thus a gap between domestic and world returns on capital to explain the mechanism. While capital flows are explicitly constrainted, their interaction with the BS effect is not examined in Gente's model. That is why Christopoulous et al. (2012) consider this issue. Their results lend support for the theoretical implications of their model: the RER appears to be mainly driven by productivity and NFA in countries that face external constraints and exclusively by productivity in countries with perfect access to international capital markets. Using a sample of 21 countries over the 1974-2004 period, they find evidence of a positive effect of the productivity gap on the non-traded to traded relative price ratio using the Groningen Growth and Development Center (GGDC) database. Furthermore, Doan and Gente (2013) also develop an OLG semi-small open economy model to investigate the relationship between the RER and countries' savings. They show that in a low-saving country and/or a high-population-growth country, a rise in productivity may appreciate or depreciates the RER.

3 How to measure the BS effect?

Section 3.1 discusses the concept of tradability. The examination of the internal version of the BS effect requires the specification of the determinants of the relative price of non-traded goods which are reviewed in Section 3.2. Finally, Section 3.3 surveys the determinants of the RER required in the examination of the external version of the BS effect.

3.1 Discussion on tradability

Tradability lies at the heart of the Balassa-Samuelson theory as the latter considers traded and non-traded goods. This concept of tradability receives both empirical and theoretical considerations. From a theoretical point of view, Betts and Kehoe (2001) provide an original definition of tradability. Rather than following the traditional tradable/non-tradable dichotomy, they prefer to rely on the degree of tradability of a good as in Obstfled and Rogoff (1996). In their approach, tradability is defined by the degree of substitutability in consumption between units of the same good produced in different countries and by the transaction cost that must be incurred to consume goods outside their country of origin. They argue that the traditional dichotomy of goods into purely tradables and non-tradables is empirically inappropriate. Tradability also receives empirical considerations. De Gregorio et al. (1994) are among the first to propose a sectoral classification between tradable and non-tradable sectors, which is currently the most widely used in empirical studies. Following De Gregorio et al. (1994), a sector is defined as tradable if the ratio of its total exports to total production exceeds a 10% threshold.[6,7] This approach presents the merit

[6] De Gregorio et al. (1994) consider 14 OECD countries over the period 1970-1985. In their analysis, the following sectors are considered: agriculture, mining, manufacturing, transportation and other services.

[7] In the following, we refer to this ratio as the tradability ratio.

to be simple and easily implementable. However, it has the drawback to impose the same sectoral classification for each country across the whole period examined. Aware of this limitation, Dumrongrittikul (2012) proposes an original approach which is applied to a panel of 33 advanced and emerging economies over the period 1970 to 2008. Contrary to De Gregorio et al. (1994), Dumrongrittikul (2012) computes the tradability ratio for each sector in a given country at a specific year.[8] He considers three different cases in his analysis. First, as in De Gregorio et al. (1994), if the tradability ratio for a year is below the 10% threshold, the sector examined is classified as non-tradable this year. If the tradability ratio is above 20%, the sector is categorized as tradable over this year. In situations in which the ratio is between 10% and 20%, he implements the Conzale-Soriano price test to help to assess the sector's tradability. In the first step, the domestic sectoral price is regressed on the world sectoral price. Then, using the residuals from the previous step, two different error correction models are estimated using the domestic and world sectoral prices as endogenous variables. If the null of no cointegration is rejected in at least one specification, the sector is classified as tradable; otherwise the domestic sectoral price is regressed on the world sectoral price. Finally, if the world sectoral price significantly affects the domestic sectoral price, the sector examined is assumed to be tradable; otherwise non-tradable. While "exogenous" thresholds are set by De Gregorio et al. (1994) and Dumrongrittikul (2012), Bems (2008) assumes that the wholesale and retail trade sector is the benchmark as this sector is known to produce non-tradable outputs.[9] According to Bems (2008), a sector is classified as non-tradable (resp. tradable) if its tradability ratio is lower (resp. higher) than the one of the wholesale and retail trade sector in the same country. Using the same approach as Bems (2008) and a 10% critical threshold, Lombardo and Ravena (2012) find substantial variation in the effective threshold, introducing thus important cross-country variability in sectors' tradability. The interest of this question of tradability is reinforced today seeing the growing shares of export services in total exports for advanced and developing economies. The share of services in total exports grew from 3% in 1970 to 23% in 2017 (Loungani et al., 2017).[10] However, not all services are equally traded across the world and significant differences between sectors exist. Loungani et al. (2017) find that Telecommunication, Computer and Information services as well as financial services are among the most dynamic sectors. The tradability of the components of the services sector remain thus an open question. In this vein, the sensitivity of the tradability of the components of the services sector to a specific threshold has been investigated by Piton (2019) on a sample of advanced economies. She shows that a threshold of 15% would exclude financial and insurance activities and information and communication from the tradable sector. Using a threshold of 20% would also exclude professional, scientific and technical activities from the tradable sector as well as information and communication.

[8] Dumrongrittikul (2012) considers the following sectors in his empirical analysis: agriculture, hunting, forestry and fishing; mining, utilities; manufacturing; construction; wholesale and retail trade, hotels, restaurants; transport, storage, communication and other activities.

[9] Its interest lies in the tradable content of investment.

[10] It is worth mentioning that the large majority of global trade in services is made by advanced economies (80% in 2014 according to Loungani et al. (2017) although we observe a sustained growth of export share in developing economies which was equal to 3% back to 1980.).

3.2 Determinants of the relative price of non-traded goods

3.2.1 Government consumption expenditures

In traditional macroeconomic frameworks, government consumption expenditures are assumed to be biased towards non-tradable goods (Froot and Rogoff, 1991). A rise in government consumption expenditures is thus associated with an increase in the non-traded goods' demand pushing-up the relative price of non-traded to traded (see Froot and Rogoff, 1991; among others). This determinant received strong empirical support in the literature, as notably shown by Galstyan and Lane (2009) and Galstyan and Velic (2018).

3.2.2 Government public debt

The public debt has been recently proposed as a new determinant of the relative price of non-tradables to tradables by Galstyan and Velic (2018). They build a theoretical framework where higher public debt can be financed through two instruments: borrowing and labor taxation. They show that if higher level of public debt is issued through the second instrument, it can affect the relative price of non-traded goods. Indeed, following a rise in labor taxation, labor supplies in the tradable and non-tradable sectors are contracted. The effect of public debt is thus ambiguous and depends on the relative factor intensity. In the case where the non-tradable sector's output is more labor intensive, we expect a positive effect of the public debt on the relative price of non-traded goods as the decrease in labor supply is higher in the non-tradable sector than in the tradable one. On the contrary, if the tradable sector is more intensive, the public debt effect is negative. Disentangling the two effects is thus an empirical issue that is investigated by Galstyan and Velic (2018). To this aim, controlling for the traditional determinants of the relative prices of non-tradables, they examine the effects of government public debt and labor taxation.[11] Considering 15 advanced economies over the period 1980–2007, they find a positive effect of labor taxation and general government gross debt on the relative price of non-tradables. Their results are consistent with a non-tradable sector exhibiting higher labor share than the traded sector.

3.2.3 Government investment

By its effects on productivity, larger stock of public capital can influence the relative price of non-tradable to tradable goods. To formalize the underlying mechanisms, Galstyan and Lane (2009) build a two-sector small open-economy model including the stock of public capital in the production functions of the tradable and non-tradable sectors. Due to its heterogeneous effects across sectors, government investment has an ambiguous impact on the non-tradable relative price. In their theoretical framework, different situations emerge depending on which sectors benefit the more from higher public capital stock. In the case where government investment mainly rises productivity in the non-tradable sector, it leads to an increase in the relative supply of non-traded goods causing a decrease in the relative price. On the contrary, if productivity gains are more concentrated in the tradable sector, the relative supply

[11] Following Eurostat methodology, labor taxation is measured as the implicit tax rate on labor. It is calculated as the "sum of all direct and indirect taxes and social contributions, divided by the total economic remuneration of employees working in the economic territory".

of tradable goods increases pushing up the relative price of non-traded goods. This theoretical indetermination is addressed by an empirical investigation, where a panel of 19 advanced countries is examined over the period 1980-2004. Galstyan and Lane (2009) confirm the existence of an ambiguous effect of government investment (expressed as share of GDP) as they obtain significant coefficients ranging from -3.63 to 8.47 on different sub-samples. Galstyan and Velic (2018) also consider this determinant in their empirical analysis. Instead of government investment, they use the public capital stock and provide evidence of a positive effect for this variable.

3.2.4 Trade balance

The trade balance is expected to influence the relative price of non-traded to traded goods dynamic through three mechanisms (Lane and Milesi-Ferretti, 2002). First of all, as a trade balance surplus is associated with an absorption level below domestic production, the demand for non-traded goods is lower. Furthermore, an improvement in the trade balance also leads to a negative wealth effect. We thus expect an increase in labor supply in both sectors, decreasing the production costs for the non-tradable sector; contributing to the decrease of the relative price. Moreover, as argued by Lane and Milesi-Ferretti (2002), the necessity to sustain a trade balance surplus also works through labor force movements from the non-tradable sector to the tradable one. From an empirical perspective, different outcomes are proposed by the literature. Lane and Milesi-Ferretti (2002) successfully confirm their predictions as higher trade balance surplus decreases the relative price of non-tradables to tradables. More recently, Galstyan and Lane (2009) fail to find support for this variable while Galstyan and Velic (2018) provide evidence of a negative effect.

3.2.5 Relative income

The level of GDP PC of a country is expected to influence the relative price of non-traded goods through the Penn-effect i.e. the positive association between per capita income and price level. Evidence in favor of this hypothesis is provided by Lane and Milesi-Ferretti (2002) showing a positive effect of relative income on the relative price of non-tradables. The impact is expected to be positive as it occurs through a positive wealth effect which increases demand for non-traded goods. The relationship between both variables can also be subject to non-linearity as shown by Hassan (2016). Indeed, the price-income relationship turns out to be significantly negative in poor countries, while it is positive for richer economies.

3.3 Determinants of the RER

This section is devoted to a quick overview of the main RER determinants for EMDEs. Our determinants selection builds extensively on the seminal paper from Edwards (1988).[12] He proposes a three-goods sector economy to apprehend real exchange rate determination for developing economies. In the long run, he shows that only real variables affect the RER dynamic of the 12 developing economies examined. His theoretical framework provides

[12] Theoretical frameworks applicable for developing and emerging market economies are also proposed by Elbadawi (1994) and Elbadawi and Soto (2004).

thus a list of variables useful to investigate the behaviour of the RER (import tariffs, terms of trade, the composition of government consumption and capital flows).

3.3.1 Trade openness

Countries' trade openness has been examined as a RER determinant because it proxies for trade liberalization (IMF, 2013). As further trade liberalization is expected to lower the domestic price level, trade openness should be negatively signed. For EMDEs experiencing significant variation in their trade openness degrees, this variable seems to be particularly relevant (Elbadawi, 1994; Dufrénot and Yehoue, 2005). Dufrénot and Yehoue (2005) obtain an elasticity equals to -0.30 for middle-income countries, while a lower effect is evidenced for low-income countries.

3.3.2 Terms Of Trade

The impact of TOT on the RER is ambiguous as two antagonists effects are at play, an income and a substitution effect (Couharde et al., 2018). The income effect leads to a positive wealth effect and increases non-traded goods' demand. In order to ensure internal balance, a real depreciation is needed. In the case of the substitution effect, producers are expected to move their production towards the tradable sector which leads to a wage increase in this sector. Assuming wage equalization across both sectors, we expect an increase in the overall price level, leading thus to an appreciation of the domestic currency.[13] Although ambiguous, the appreciating effect of higher terms of trade on the RER is a standard result (Elbadawi, 2004). For example, empirical evidence provided by Dufrénot and Yehoue (2005) show that the income effect dominates the substitution effect on a sample of low and middle income countries.

3.3.3 Government consumption expenditures

The government consumption expenditures are expected to affect the RER through the well-known Froot-Rogoff effect (Froot and Rogoff, 1991). As government consumption expenditures are biased towards non-traded goods, a positive effect is expected for this determinant. This variable has been investigated through several studies (see Edwards, 1988; De Gregorio and Wolf, 1994; Chinn, 1997). Examining 73 developing countries over the 1970-2004 period, Elbadawi (2004) finds that a lower fraction of government expenditures on traded goods leads to RER appreciation.

3.3.4 Fertility rate

The fertility rate is expected to influence the RER through the savings channel. A higher fertility rate is associated with lower savings due to the increase in the young dependency ratio. At the same time, it also increases domestic investment due to the decline in the future equilibrium capital stock. The current account is then deteriorated and the domestic currency appreciates (see Higgings, 1998; Rose et al., 2009; Hassan et al., 2011). Moreover, as the

[13] In line with the TOT effect, is worth mentioning that part of the literature investigates the existence of commodity currencies (see Cashin et al. (2004) among others).

young's consumption is biased towards non-traded goods (Rose et al., 2009), a higher fertility rate is thus again expected to be positively signed. Examining a panel of 87 countries over the period 1975–2005, Rose et al. (2009) provide robust evidence of a positive sign for this determinant.

3.3.5 Government investment

In addition to investigate the effect of government investment on the non-traded to traded relative price, Galstyan and Lane (2009) also examine how relative government investment affects the RER. They failed to find a significant effect of relative government investment on the RER of advanced economies over the period 1980-2004.

3.3.6 Foreign Direct Investment

The theoretical effect of the Foreign Direct Investment (FDI) on the RER is ambiguous. Following Kosteletou and Liargovas (2000), we assume a trade integrated model in the presence of a small open economy, which is a price taker. Within this framework, the FDI effect depends on the way capital inflows are used in the domestic economy. These effects are likely to work through two main transmission channels. On the one hand, if capital inflows are used for a spending finance purpose, it raises the traded and non-traded goods' demand. As the traded goods' price is fixed in the world economy, higher FDI appreciates the RER. On the other hand, FDI also affects capital accumulation in the non-tradable and tradable sectors which raises productivity in both sectors. The net effect of this channel thus depends on which sector benefits the more from the productivity increase.

3.3.7 Net Foreign Asset position

The connection between real exchange rates and net foreign assets derives from the intertemporal budget constraint which links external assets, real exchange rate and trade balance together, as documented by Lane and Milesi-Ferretti (2002). When a country runs a current account deficit, it is building up liabilities to the rest of the world. Solvency requires that the country be willing and able to (eventually) generate sufficient current account surpluses to repay what it has borrowed to finance the current account deficits. Therefore, a country running a current account deficit (borrowing more) may have an overvalued currency. Indeed, it should register a more depreciated real exchange rate in order to restore external equilibrium. Conversely, net creditor countries may have an undervalued currency and experience real exchange rate appreciations: their trade deficit will be indeed offset by investment income on their net foreign asset position.

3.3.8 Investment

The relevance of domestic investment (as GDP share) as a RER determinant has been proposed by Edwards (1989). The effect of this variable depends on the composition of investment terms of tradable and non-tradable goods. If higher investment rate raises the relative share of tradable goods, an exchange rate depreciation is expected. A currency appreciation is possible if an increasing investment rate is associated with a rise in the relative non-tradable share. Edwards (1989) finds that a higher investment rate depreciates the RER.

4 Overview of the evidence

The end of the 1990s and the beginning of the 2000s gave birth to a growing literature investigating the BS effect. At the heart of these studies are either heterogeneous groups of countries, or specific geographical areas. For example, Drine and Rault (2003) and Iyke and Odhiambo (2017) investigate the existence of the BS effect in Africa. The first ones consider a sample of 16 Middle East and North Africa (MENA) countries over the period 1960-1999. They use GDP PC to measure the productivity differential. While they failed to find evidence in favor of the productivity bias hypothesis for 11 of the 16 economies examined, their panel cointegration analysis provides support for the BS effect. Considering a smaller sample of countries over a longer period of time, Iyke and Odhiambo (2017) confirm Drine and Rault (2003)'s findings in a panel framework.

Besides the scarce studies existing for the African continent, the Asia-Pacific region has been, on the contrary, a source of great interest in the literature. This enthusiasm is explained by the sustained growth experienced by South-East Asian economies over the last decades of the previous century, making these countries ideal candidates for the investigation of the BS effect. The very influential study of Chinn (2000) is among the first to document the main determinants of the RER dynamic in this region. He examines 9 Asian countries over the period 1970-1992. The relative productivity between the tradable and non-tradable sectors is measured using the difference in the value-added per worker in both sectors.[14] His time-series analysis validates the existence of the BS effect for 5 of the 9 Asian economies examined (Indonesia, Japan, Korea, Malaysia and the Philippines). Similar results are provided by Bahmani-Oskooee and Ree (1996) for the Korean won, and by Ito et al. (1996) for the yen and the Taiwan dollar. Updating the dataset from Chinn (2000) to a most recent period, King and Thomas (2008) show that his findings continue to be valid. While previous authors only consider an external version of the BS hypothesis, Drine and Rault (2004) propose a more in-depth investigation for a panel of 9 Asian economies. They test the three following hypotheses: the positive relationship between productivity differential and relative price of tradables, the positive relationship between relative prices of tradables and REER and the verification of the PPP. All in all, they confirm the last two hypotheses, while the first one is rejected. While Chinn (2000) and King and Thomas (2008) do not assess the contribution of each determinant to the RER's variation, Tsen (2011) performs a generalized forecast error variance decomposition to analyse the relative importance of the terms of trade, productivity differential,[15] real oil price and reserve differential for Hong Kong, Japan and Korea. His results show that productivity differential accounts for about 1%, 2% and 8% of the RER variation respectively for Japan, Hong Kong and Korea. To circumvent the potential drawbacks associated with the use of labor productivity measures, Kakkar and Yan (2012) prefer to rely on Total Factor Productivity (TFP) measures. To this aim, TFP series for the tradable and non-tradable sectors are constructed for 6 Asian economies[16] as the Solow residuals (1957) from a Cobb-Douglas production function. Their results provide support for both the internal and external versions of the BS hypothesis.

[14] The tradable sector consists of manufacturing, while the non-tradable one is composed of services, construction, mining, and transportation.

[15] Tsen (2011) approximates the productivity differential by the ratio of GDP in volume to manufacturing employment relative to the US.

[16] Hong Kong, Indonesia, Korea, Malaysia, Singapore and Thailand.

Given their recent history, the Central Eastern European Countries (CEECs) constitute a relevant research laboratory for who seeks to investigate the BS hypothesis.[17,18] Indeed, during the initial transition period,[19] the apparition of three key elements makes these economies a particularly relevant case of study: the establishment of free-market economies, the creation of new currencies and the productivity rebound. It is thus quite naturally that the BS effect has been proposed as the prime explanation for CEECs' RER movements during the transition period. Defining relative productivity as the ratio between industrial and services productivity, De Broeck and Sløk (2006) confirm the existence of the BS effect for the European Union (EU) acceding countries. Using time-series and panel econometric techniques, Egert (2002) also finds that the BS effect works well for transition economies. However, although explaining part of their exchange rate variation, this effect fails to fully explain its dynamic (Egert, 2002; Klau and Mihaljek, 2004; Egert et al., 2006). Thereupon, Egert et al. (2006) argue that: "at best, half of this appreciation can be ascribed to the B-S effect". That is why alternative explanations have been proposed, notably by Fischer (2004). He builds an amended BS framework which consists of three sectors and four inputs, where investment demand is explicitly taken into account. In his framework, a productivity shock affects both the supply and demand sides of the economy.[20] He shows that a productivity shock in any sector raises the equilibrium stock of physical capital, increasing investment demand which, in turn, rises the relative price of non-tradables.

Because of the role of currency misalignments as an Early Warning Indicator of currency crises, the BS effect has been examined as a long term driver of the equilibrium exchange rate in Latin American Countries (LAC). Over comparable samples of countries and over more or less long-time spans, panel estimations provide evidence in favor of the BS hypothesis within this area (Alberola et al., 2003; Drine and Rault, 2003; Garcià-Solanes and Torrejon-Flores, 2009; Iyke, 2017). These different investigations rely on various proxies: Consumer Price Index (CPI)-to-Producer Price Index (PPI) ratio (Alberola et al., 2003), GDP PC (Drine and Rault, 2003) and average labor productivity (Garcià-Solanes and Torrejon-Flores, 2009; Iyke, 2017).

Probably guided by data availability motivations, the first attempts of the BS investigations focused on OECD economies.[21] Earlier studies by Chinn and Johnston (1996) and Chinn (1997) make use of industry specific TFP series for these countries. Considering 14 OECD economies over the period 1970-1991, they established that TFP differential is a driver of the long-run RER movements. More recently, Ricci et al. (2013) fail to find evidence for this effect for 20 advanced economies over the period 1980-2004.[22] Using respectively TFP differential and productivity in the tradable sector, Lee and Tang (2007) and Gubler and Sax (2019) obtain results in contradiction with the

[17] According to the Organisation for Economic Co-operation and Development's (OECD) definition, CEECs includes: Albania, Bulgaria, Croatia, the Czech Republic, Hungary, Poland, Romania, the Slovak Republic, Slovenia, and the three Baltic States: Estonia, Latvia and Lithuania. Following the literature dealing with the BS effect for these countries, the use of CEECs term rather refers to the following countries: Czech Republic, Hungary, Poland, Slovakia and Slovenia.

[18] In the following, we only review the main papers investigating this issue, see the survey by Egert et al. (2006) for more details.

[19] According to De Broeck and Sløk (2006), the transition period spans from 1991 to 1998.

[20] It is obvious that this transmission mechanism is in contradiction with the traditional BS effect as the demand side is also affected.

[21] For example, the STructural ANalysis (STAN) database from OECD has been widely used at the end of the 1990s and the beginning of the 2000s to investigate the effect of productivity differential on the exchange rate.

[22] In their empirical investigation, Ricci et al. (2013) rely on six-sectors' value-added deflator.

traditional view, as a higher productivity depreciates the RER. Contrary to the previous authors who use a static classification of tradability, Dumrongrittikul (2012) implements a new approach allowing for country-specific and time varying tradability over each industry. With this new methodology, he finds results inconsistent with the BS effect as the traded productivity growth depreciates the REER. More recently, applying a new panel cointegration test, Wang et al. (2016) confirm the existence of a cointegrating relationship between GDP PC and REER for 20 advanced economies. Using the group mean estimator, they find an elasticity equals to 0.76 confirming thus the BS effect. The numerous studies previously discussed rely on various proxies, creating thus uncertainty on the robustness of this effect to the measures. This essential issue has been considered by Bénassy-Quéré et al. (2009) for the G20 countries over the period 1980-2005 through the use of four different proxies: the CPI-to-PPI ratio, three-sectors' value-added deflator, GDP PC and GDP PW. Estimated coefficients for the first two variables are shown to be close to unity, while the last two proxies present lower marginal effects. Besides these panel analyses, time-series investigations have also been performed for advanced economies. However, at best only mixed results are obtained (see Faria and Ledesma, 2003; Bahmani-Oskooee and Nasir, 2004; Drine and Rault, 2005; among others).

Concerning EMDEs, only scarce evidence is provided regarding the BS effect. Choudhri and Kahn (2007) propose a stimulating empirical investigation of the BS hypothesis for a sample of 16 developing economies over the 1976-1994 period. They find evidence of a positive effect of the relative price of non-traded to traded goods on the RER. Finally, they also confirm the internal version as a rise in the productivity differential is associated with higher relative price of non-tradables. Ricci et al. (2013) confirm these previous findings. They show a significant impact of relative productivity on the RER for emerging markets. While previous studies find support for the productivity bias hypothesis with sectoral measures, aggregate proxies seem to be clearly inadequate to corroborate the BS hypothesis. For example, using GDP PW, Bahmani-Oskooee and Nasir (2004) find that the major fail of the productivity bias hypothesis is due to the emerging economies in their time-series analysis of 44 advanced and emerging economies. Relying on a panel cointegration analysis, Wang et al. (2016) confirm the previous result. Indeed, the null hypothesis of no cointegration between GDP PC and RER is not rejected for a sample of 20 developing economies.

5 Data and methodology

5.1 Which proxies to measure the Balassa-Samuelson effect?

As we aim to investigate the internal and external versions of the BS effect, two "ingredients" are required. As a reminder, the internal version is a positive relationship between the productivity differential and the non-traded to traded relative price. Hence, a proper measurement of both variables is required.[23] To achieve this purpose, some of the existing databases can be useful. For example, the EU KLEMS database is of interest for who investigates the BS hypothesis in the European Union (EU). This database allows the construction of quite precise measures because of the availability of the following variables: sectoral value-added (at current and constant prices) and sectoral employment over the 1995-2015 period for 28 member states of the EU. A very narrow classification is also available: 34 industries and 8 aggregates. Consideration of a larger countries sample is possible if the World Input Output (WIO) database is used.[24] Although having the appealing feature to present 56 sectors, its time-series span is more limited covering the period 2000-2014. Furthermore, given our interest in EMDEs, the most interesting database for our purpose would be the GGDC database. Unfortunately, although being available from almost the year 1950, the end date is too early for us and only a few number of EMDEs are covered. Finally, we can also mention the database from Mano and Castillo (2015) which makes available real labor productivity in the non-traded and traded sectors for 56 countries.[25] All in all, although presenting very appealing features, these different databases are hardly compatible with our purposes. As we need to measure the relative price of non-tradables, we rely on sectoral value-added deflators. The use of relative prices can be difficult as an accurate separation between the tradable and non-tradable sectors is required.[26] We thus use three-sectors' value-added deflators to track the evolution of the relative price of non-tradables to tradables.[27] The three-sectors' value-added deflator is based on (i) agricultural, (ii) industrial, and (iii) services sectors.[28] Total services are used to represent the non-tradable sector, whereas manufacturing and agriculture are taken together to represent the tradable sector (Choudhri and Kahn, 2005; Bénassy-Quéré et al., 2009). This decomposition is usually considered as broadly consistent with the tradability measure (Betts and Kehoe, 2001).

For a country i, the price index of value-added at time t for each sector k ($PVA^k_{i,t}$) is calculated by dividing value-added at current prices by value-added at constant prices in the accounting period and the considered sector, using 2010 as base year. As the tradable sector is composed of 2 sectors, the latter have to be aggregated together to form a composite tradable sector. To this aim, country-specific weights for each sector have been used which are measured by the value-added share of the sector in total output (i.e the sum of the values added of agriculture and manufacturing sectors). Although commonly employed in the literature, it can be viewed as a rather strong assumption as it does not take into account the evolution of the productive base. However, it ensures that the only source behind the relative price of non-tradables' dynamic is variations in non-tradable or tradable goods' price. It is thus not driven by any variation in sectoral weights. As an alternative to fixed country-specific weights, we also

[23]Obviously the investigation of the external version of the BS hypothesis also requires a measure of non-traded to traded relative price.

[24]In addition to the 28 previous EU countries, 15 other major countries are available.

[25]Their database contains advanced and some emerging economies over the period 1989-2013.

[26]This issue is also widely discussed in the literature focusing on the role of the relative price of non-tradable goods in accounting for real exchange rate fluctuations (see the seminal and most cited paper of Engel (1999)).

[27]It is possible to rely on the CPI-to-PPI ratio to track the evolution of the relative price of non-tradables to tradables. In this configuration, PPI mainly concerns tradables' prices, whereas CPI covers essentially non-tradables' prices. However, PPI is unfortunately unavailable for most countries of our sample over a long time span.

[28]Agriculture corresponds to International Standard Industrial Classification (ISIC) divisions 1-5, and includes forestry, hunting, and fishing, cultivation of crops and livestock production. Industry corresponds to ISIC divisions 10-45, and services to ISIC divisions 50-99.

compute time-varying country-specific weights for each industry. We get for a country i at time t:

$$pva_{i,t,f} = pva_{i,t}^{serv} - \beta pva_{i,t}^{agr} - (1-\beta) pva_{i,t}^{ind} \tag{14}$$

$$pva_{i,t,tv} = pva_{i,t}^{serv} - \beta_t pva_{i,t}^{agr} - (1-\beta_t) pva_{i,t}^{ind} \tag{15}$$

where β and β_t for a country i are computed as follows:

$$\beta = \frac{\sum_{t=1}^{T} VA^{agr}}{\sum_{t=1}^{T} VA^{agr} + \sum_{t=1}^{T} VA^{ind}} \tag{16}$$

$$\beta_t = \frac{VA_t^{agr}}{VA_t^{agr} + VA_t^{ind}} \tag{17}$$

Computing the value-added deflators as the deviation from the main trading partners, we get:

$$def3\_fr_{i,t} = pva_{i,t,f} - \sum_{j=1}^{N} (w_{i,j,t} \times pva_{j,t,f}) \tag{18}$$

$$def3\_tvr_{i,t} = pva_{i,t,tv} - \sum_{j=1}^{N} (w_{i,j,t} \times pva_{j,t,tv}) \tag{19}$$

where $pva_{i,t,f}$ and $pva_{i,t,tv}$ are respectively the values added deflator based on fixed and time-varying weights, for country i expressed in logarithms. $serv$, agr and ind respectively denote services, agriculture and industry. $w_{i,j,t}$ is country i's trade-based weights for all its partners j, N denoting the number of trading partners. $def3\_fr_{i,t}$ and $def3\_tvr_{i,t}$ are respectively three-sectors value-added deflators computed with fixed and time-varying weights as a deviation from trading partners.

We also consider a six-sectors' value added deflator. In the six-sector disaggregation, the following sectors are distinguished: (i) agriculture, hunting, forestry, fishing; (ii) mining, manufacturing; (iii) construction; (iv) wholesale, retail trade, restaurants and hotels; (v) transport, storage and communications; and (vi) other activities.[29] Following De Gregorio and al.'s (1994) classification, construction, wholesale, retail trade, restaurants and hotels, and other services are classified in the non-tradable sector, while agriculture, manufacturing, mining, utilities and transport are treated as tradable goods.

To derive the six-sectors' deflator, we follow Lee and Tang (2007) and compute country-specific weights for each sector ($\omega_{i,k}$), measured by its value-added share in total output:

[29]The description of the divisions associated with each of these sectors is available in Table B.7 in the Appendix.

$$\omega_{i,k} = \frac{\sum_{t=1}^{T} VA_{i,k,t}}{\sum_{k\in h}(\sum_{t=1}^{T} VA_{i,k,t})} \tag{20}$$

where h denotes the nature of the sector k under consideration, i.e., tradable (T) or non-tradable (NT) sector. We also compute time-varying country-specific weights for each sector k ($\omega_{i,k,t}$), measured by its value-added share in total output at a year t:

$$\omega_{i,k,t} = \frac{VA_{i,k,t}}{\sum_{k\in h}(VA_{i,k,t})} \tag{21}$$

For each country i, the aggregated value added deflator of the non-tradable ($pva_{i,t}^{NT}$) and tradable ($pva_{i,t}^{T}$) sectors is then calculated as a weighting average of value-added deflators for respectively all non-tradable sectors and all tradable sectors:

$$pva_{i,t,f}^{NT} = \sum_{k\in NT}(\omega_{i,k} \times pva_{i,t}^{k}) \tag{22}$$

$$pva_{i,t,f}^{T} = \sum_{k\in T}(\omega_{i,k} \times pva_{i,t}^{k}) \tag{23}$$

$$pva_{i,t,tv}^{NT} = \sum_{k\in NT}(\omega_{i,k,t} \times pva_{i,t}^{k}) \tag{24}$$

$$pva_{i,t,tv}^{T} = \sum_{k\in T}(\omega_{i,k,t} \times pva_{i,t}^{k}) \tag{25}$$

where $pva_{i,t,f}^{NT}$, $pva_{i,t,f}^{T}$, $pva_{i,t,tv}^{NT}$ and $pva_{i,t,tv}^{T}$ are expressed in logarithmic terms. Denoting $def6$ the BS measure based on six-sectors' value-added deflators, we get for a country i at time t:

$$def6\_f_{i,t} = (pva_{i,t,f}^{NT} - pva_{i,t,f}^{T}) \tag{26}$$

$$def6\_tv_{i,t} = (pva_{i,t,tv}^{NT} - pva_{i,t,tv}^{T}) \tag{27}$$

As previously, variables are also expressed as a deviation from the main trading partners:

$$def6\_fr_{i,t} = def6\_f_{i,t} - \sum_{j=1}^{N} w_{i,j,t}(pva_{j,t,f}^{NT} - pva_{j,t,f}^{T}) \tag{28}$$

$$def6\_tvr_{i,t} = def6\_tv_{i,t} - \sum_{j=1}^{N} w_{i,j,t}(pva_{j,t,tv}^{NT} - pva_{j,t,tv}^{T}) \tag{29}$$

The internal version of the BS hypothesis also requires to measure the productivity differential between tradable and non-tradable sectors. To this aim, we rely on the sectoral value-added per worker of three sectors: agriculture, industry and services.[30] As previously, we assume that agriculture and industry are tradable sectors, and services are assumed to cover non-tradable goods. The productivity differential (in logarithm) is computed as follows:

$$prod\_diff\_tv_{i,t} = a^{T}_{i,t,tv} - a^{NT}_{i,t} \tag{30}$$

$$prod\_diff\_fixed_{i,t} = a^{T}_{i,t,f} - a^{NT}_{i,t} \tag{31}$$

where $a^{T}_{i,t,f}=\beta vapw^{agr}_{i,t} + (1-\beta)vapw^{ind}_{i,t}$, $a^{T}_{i,t,tv}=\beta_t vapw^{agr}_{i,t} + (1-\beta_t)vapw^{ind}_{i,t}$ and $a^{NT}_{i,t}=vapw^{serv}_{i,t}$. $vapw_{i,t}$ is the value-added per worker of country i expressed in logarithms. *prod_diff_tv* and *prod_diff_fixed* stand respectively for the productivity differential between the tradable and the non-tradable sectors computed with time-varying and fixed weights.

Another way to assess relative productivity is to rely on GDP PW. Investigations of the BS hypothesis have also frequently been carried out relying on relative labor productivity to capture total-economy productivity differentials (see Hsieh, 1982; Marston, 1986; Canzoneri et al., 1999; Schnatz, 2004; Bénassy-Quéré et al., 2009; among others). In this respect, we measure labor productivity as GDP per worker in constant 2011 PPP U.S. dollars.

5.2 Empirical strategy

In the following, we briefly outline our empirical strategy. The first step consists to select the RER determinants used in the examination of the external version of the BS effect. To tackle the uncertainty surrounding the selection of RER determinants, we rely on the Bayesian Model Averaging (BMA) methodology to have a parsimonious specification. The potential RER drivers examined are those reviewed in Section 3.3: trade openness, TOT, government consumption expenditures, fertility rate, government investment, FDI, NFA and investment. Specifically, the external version is apprehended through the following equation:

$$reer_{i,t} = \alpha_i + \beta_1 X_{i,t} + \gamma Z_{i,t} + \epsilon_{i,t} \tag{32}$$

where *reer* is the logarithm of the REER. $Z_{i,t}$ stands for the REER determinants obtained using the BMA approach. $X_{i,t}$ refers to the different measures used to proxy the relative price of non-tradables to tradables in deviation from the main trading partners ($def6\_tvr_{i,t}$, $def6\_fr_{i,t}$, $def3\_tvr_{i,t}$ and $def3\_fr_{i,t}$).

In a second step, we test the null hypothesis of cross-sectional dependence (CSD) to apply the appropriate panel unit root tests. To this aim, we rely on the Pesaran (2004)'s test (see Section C.A in Appendix). In the next step of

[30] Our first best will be to rely on a six sectoral values added per worker. However, due to data availability issues, this was impossible.

our empirical strategy, we implement the CIPS (Pesaran, 2007) and the Bai and Carrion-I-Silvestre (BCIS hereafter, 2009) panel unit root tests to assess the integration order of our series (see Section C.B in the Appendix). The CIPS and BCIS tests allow us to control for the presence of CSD. It is worth noting that the BCIS test is also robust to the existence of structural breaks in the series. Finally, our estimations are performed using the CS-DL approach proposed by Chudik et al.(2016) (see Section C.C in Appendix). This approach allows us to assess long-run effects in a panel where CSD is present, regardless if the regressors are I(0) or I(1), even if the common factors contain unit roots (Chudik et al., 2016). This estimator is interesting for our investigation as it presents better small sample performance compared to the CS-ARDL approach, for moderately large time-series ($30 \leq T \leq 50$) as shown by Monte Carlo simulations (Chudik et al., 2016). To check the robustness of our results, we follow the suggestion formulated by Chudik et al.(2017) which is to test different lags of the first-differenced regressors. We start our empirical investigation by documenting the non-traded to traded relative price determinants for EMDEs. To this aim, the two following equations are estimated:

$$\text{def6\_tv}_{i,t} = \alpha_i + \beta_1 \text{gov cons}_{i,t} + \beta_2 \text{public debt}_{i,t} + \beta_3 \text{tb}_{i,t} + \beta_4 \text{yc}_{i,t} + \beta_5 \text{public capital}_{i,t} + \epsilon_{i,t} \tag{33}$$

$$\text{def6\_tv}_{i,t} = \alpha_i + \beta_1 \text{gov cons}_{i,t} + \beta_2 \text{public debt}_{i,t} + \beta_3 \text{tb}_{i,t} + \beta_4 \text{yc}_{i,t} + \beta_5 \text{gov inves}_{i,t} + \epsilon_{i,t} \tag{34}$$

Where i=1,...,N and t=1,...,T respectively denote the cross-sectional and time-series dimensions of the panel. *gov cons* stands for the government consumption expenditures. *public debt* is the gross government public debt. *tb* is the trade balance, *yc* the logarithm of the GDP PC in constant PPP terms, *public capital* is the public capital stock and *gov inves* is the government investment. $\epsilon_{i,t}$ stands for a serially uncorrelated error term that, potentially, could be dependent across countries due to the presence of unobserved common factors. As can be seen from equations (33) and (34), two different specifications are examined. The first one considers the public capital stock, while the second one analyses the effect of government investment. As there is no consensus on the more adequate variable, we thus examine these two measures. Furthermore, we also investigate if the lagged values of both variables can help to explain the dynamic of the non-traded to traded relative price. It can be interesting as the productivity effect of these variables is likely to take time. Furthermore, we check the robustness of our results to the aggregation method. To do so, we also estimate equations (33) and (34) using the six-sectors value-added deflators with fixed weights as the dependent variable.

Our investigation of the robustness of the BS hypothesis begins with the consideration of its internal version. As suggested by the simple theoretical framework presented earlier, a rise in the productivity differential should be associated with a higher domestic relative price of non-tradables. To investigate this hypothesis, we examine two different measures of productivity: GDP PW and the productivity differential between the tradable and non-tradable sectors. For each proxy, we consider two different specifications. The first one excludes government investment from the regressors to properly disentangle the productivity's effect. This goal is difficult to achieve if government investment is also included as it is expected to affect the relative price of non-tradables through a BS effect (Galstyan and Lane, 2009). To see how government investment affects our previous findings, our second regression augments the previous one with the government investment.[31] All in all, our different specifications are the following:

$$\text{def6\_tv}_{i,t} = \alpha_i + \beta_1 \text{prod\_diff}_{i,t} + \beta_2 \text{gov cons}_{i,t} + \beta_3 \text{public debt}_{i,t} + \beta_4 \text{tb}_{i,t} + \beta_5 \text{yc}_{i,t} + \epsilon_{i,t} \tag{35}$$

$$\text{def6\_tv}_{i,t} = \alpha_i + \beta_1 \text{gdp pw}_{i,t} + \beta_2 \text{gov cons}_{i,t} + \beta_3 \text{public debt}_{i,t} + \beta_4 \text{tb}_{i,t} + \epsilon_{i,t} \tag{36}$$

$$\begin{aligned} \text{def6\_}tv_{i,t} = \alpha_i + \beta_2 \text{gdp pw}_{i,t} + \beta_2 \text{gov cons}_{i,t} + \beta_3 \text{public debt}_{i,t} + \beta_4 \text{tb}_{i,t} + \beta_5 \text{yc}_{i,t} \\ + \beta_6 \text{gov inves}_{i,t} + \epsilon_{i,t} \end{aligned} \tag{37}$$

$$\text{def6\_tv}_{i,t} = \alpha_i + \beta_2 \text{gdp pw}_{i,t} + \beta_2 \text{gov cons}_{i,t} + \beta_3 \text{public debt}_{i,t} + \beta_4 \text{tb}_{i,t} + \beta_5 \text{gov inves}_{i,t} + \epsilon_{i,t} \tag{38}$$

gdp pw stands for the logarithm of the gross domestic product per worker. *prod_diff* is the logarithm of the productivity differential between the tradable and non-tradable sectors. We use the *prod_diff_fixed* and *prod_diff_tv* variables. Unfortunately, the productivity differential only starts in 1991 while GDP PW begins in 1980. A comparison between the two proxies is hard because of the shorter span for the productivity differential. Moreover, due to collinearity between GDP PW and GDP PC, equations (36) and (38) do not include *yc*. As previously, we check the robustness of our results to the aggregation method using also the six-sectors value-added deflators with fixed weights as the dependent variable for equations (35) to (38).

5.3 Data sources

Our sample is composed of 38 developing and emerging economies, equally divided between both groups.[32] Our sample is representative of the EMDEs world as shown by the diversity of its geographical coverage: Central and South America, South-Eastern Asia and different regions of Africa (see Table A.1 in the Appendix). Our dataset spans from 1980 to 2016. A detailed description of the sources, as well as additional information on the variables, are available in Table A.2 in the Appendix. The trade balance and government consumption expenditures, expressed as a share of GDP, come from the World Development Indicators (WDI) database of the World Bank. General government final consumption expenditures include all government current expenditures for purchases of goods and services (including compensation of employees). The general government (i.e., central plus subnational governments) investment (gross fixed capital formation) is extracted from the International Monetary Fund (IMF, fiscal affairs department). The general government capital stock arises from the same source and is computed using the perpetual inventory method. Moreover, Central government debt (as a share of GDP) series are extracted from WDI. For

[31] As in the seminal paper of Galstyan and Lane (2009), we use government investment as share of GDP.

[32] The choice of our sample is guided by data availability issues to obtain a balanced panel. To classify our countries between developing and emerging economies, we rely on the classification of the International Monetary Fund from Ghosh et al. (2014).

the sake of comparison with previous studies, we express these three variables as a percent of GDP. GDP PC, expressed in constant PPP, and the majority of the REER's determinants (terms of trade, trade openness, fertility rate, investment and FDI) are extracted from WDI. The NFA position comes from the Lane and Milesi-Ferretti database (2007). The REER and the weights used in the computation of the variables expressed in relative terms (i.e. relatively to trading partners) are extracted from the EQCHANGE database (CEPII, Couharde et al., 2018). The GDP PC expressed as a deviation from the trading partners comes from the same source. The data used in the computation of the non-traded to traded goods' price arise from various sources. Current and constant values added for agriculture, industry and services are extracted from the WDI database. The current and constant values added required for the six-sectors' value-added deflators are taken from United Nations Conference on Trade and Development (UNCTAD). The value-added per worker series for the agriculture, industry and services sectors are extracted from the International Labour Organization (ILOSTAT) database. They are available only from 1991 to 2016. The GDP PW expressed in PPP is calculated using the Penn World Table 9.0 for the 1980-2014 period. The 2015 and 2016 years are filled up using the ILOSTAT database.

6 Econometric results

6.1 Selection of REER determinants: BMA results

Table 1 below displays the BMA results. Among the eight potential determinants, three of them present decisive evidence of having an effect as illustrated by their PIPs over 0.99. Indeed, in all specifications considered, the PIP is equal to one for TOT and the fertility rate. Our result for TOT is conform to previous findings (see Grekou, 2019). Moreover, the openness ratio and government consumption expenditures belong to the robust REER determinants. The BMA analysis confirms thus the crucial role attributed to the openness ratio for developing economies by earlier studies. The four other potential determinants can not be considered as robust variables as shown by their PIPs below 0.50.

6.2 CSD and panel unit root tests results

Once the REER determinants have been selected, we test for the presence of CSD in our series. Table B.1 in the Appendix displays the results of the Pesaran (2004)'s test. We conclude to the presence of CSD as the null hypothesis of cross-sectionally uncorrelated errors is strongly rejected for all the variables examined. Table B.2 in the Appendix displays the results of these panel unit root tests, a summary being provided in Table 2 below. For the CIPS test, two different specifications are examined. The first one includes a trend and a constant, while the second only assumes a constant in the regression. The BCIS test considers level shift as well as trend level shift. The majority of the test statistics points in favor of stationarity for the trade balance, government investment and public capital stock as the null hypothesis is almost unanimously rejected. Moreover, the last two determinants of the non-traded to traded relative price (i.e. government consumption expenditures and GDP PC) are integrated of order one as their first difference is stationary.

Table 1: BMA results: REER determinants

	Model prior	Random	Random	Fixed	Fixed	Uniform
	Parameter prior	UIP	RIC	UIP	RIC	UIP
tot		**1.000**[a]	**1.000**[a]	**1.000**[a]	**1.000**[a]	**1.000**[a]
fertility		**1.000**[a]	**1.000**[a]	**1.000**[a]	**1.000**[a]	**1.000**[a]
open		**0.996**[a]	**0.997**[a]	**0.994**[a]	**0.994**[a]	**0.997**[a]
gov cons		**0.983**[a]	**0.987**[a]	**0.972**[a]	**0.971**[a]	**0.984**[a]
fdi		0.480	0.479	0.255	0.266	0.478
gov inves		0.313	0.309	0.136	0.134	0.308
investment		0.248	0.248	0.094	0.095	0.245
nfa		0.246	0.247	0.093	0.091	0.244

Note: The results are based on 100.000 burn-ins and 200.000 draws. Simulations are made using birth-death MCMC sampler. "^{a}" denotes a PIP over 0.99 and decisive evidence of a regressor. Numbers in bold denote PIP over 0.50 and evidence of robustness. RIC=Risk Inflation Criterion. UIP= Unit Information Prior.

Our results provide mixed evidence regarding the presence of unit root for the two six-sectors' value-added deflators, particularly depending on the specification. Moving to the three and six sectors' value-added deflators in relative terms, we find that these series are integrated of order one as shown by the results of the CIPS and BCIS on their first differences. Moving to REER, the usual evidence of unit root is confirmed for this variable. A quite similar conclusion is shared by trade openness and TOT which are integrated or order 1. Examining the fertility rate, the conclusions are less clear-cut as some evidence of unit root is provided. Finally, our different productivity proxies (*gdp pw*, *prod_diff_tv*, *prod_diff_fixed* and *yc_r*) are first difference stationary.

6.3 Relative non-traded to traded price determinants: empirical results

Table 3 below displays the results regarding the relative non-traded to traded price determinants.

Columns (1) to (4) display the results for the six-sectors' value-added deflator based on a time-varying weights aggregation scheme, while columns (5) to (8) present the findings associated with the six-sectors' deflator with fixed weights. Our results show that whatever the specification examined, the government consumption expenditures variable is significant. This variable is thus an essential driver of the relative price of non-tradables for EMDEs. As expected, this determinant is positively signed as higher government consumption expenditures are associated with a rise in the non-traded goods' demand. Hence, it pushes up the non-traded price, increasing the relative price of non-traded to traded goods. The marginal effect of government consumption expenditures also seems to be higher if the sectors are aggregated with fixed weights, as shown by the comparison between columns (5) to (8) with columns (1) to (4). Although being a crucial determinant of the relative price of non-tradables for EMDEs, its marginal effect remains quite low with an average coefficient of about 2.083. However, it is four times higher than

Table 2: **Panel unit root tests: summary**

Variable	Abbreviation	Integration order
PNT_PT determinants		
External balance	tb	I(0)
Gross net debt	public debt	I(0)/I(1)
Government consumption expenditures	gov cons	I(1)
Government investment	gov inves	I(0)
Public capital stock	public capital	I(0)
GDP PC	yc	I(1)
Price deflators		
Deflator 6 sectors (TV weights)	def6_tv	I(0)/I(1)
Deflator 6 sectors (Fixed weights)	def6_f	I(0)/I(1)
Deflator 6 sectors relative (TV weights)	def6_tvr	I(1)
Deflator 6 sectors relative (Fixed weights)	def6_fr	I(1)
Deflator 3 sectors relative (TV weights)	def3_tvr	I(0)/I(1)
Deflator 3 sectors relative (Fixed weights)	def3_fr	I(1)
REER determinants		
Real Effective Exchange Rates	reer	I(1)
Openess	open	I(0)/I(1)
Fertility rate	fertility	I(1)
Terms Of Trade	tot	I(1)
Productivity proxies		
GDP PC relative	ycr	I(1)
Productivity differential (TV weights)	prod_diff_tv	I(1)
Productivity differential (Fixed weights)	prod_diff_fixed	I(1)
Gross Domestic Product Per Worker	gdp pw	I(1)

Source: Author's calculations

Table 3: **Relative non-tradable price determinants: whole sample**

	def6_tv	**def6_tv**	**def6_tv**	**def6_tv**	**def6_f**	**def6_f**	**def6_f**	**def6_f**
	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
gov cons	1.48**	2.11**	1.93***	2.15***	1.70**	2.44*	2.33**	2.53**
	(2)	(2.18)	(2.57)	(2.62)	(2.05)	(1.94)	(2.53)	(2.53)
public debt	-0.07	-0.07	-0.09	-0.07	-0.11	-0.09	-0.13	-0.08
	(-0.88)	(-0.78)	(-0.90)	(-0.50)	(-1.57)	(-1.00)	(-1.30)	(-0.67)
tb	-0.38***	-0.55***	-0.3	-0.14	-0.49***	-0.62***	-0.32	-0.17
	(-2.71)	(-3.24)	(-1.43)	(-0.50)	(-2.88)	(-3.26)	(-1.33)	(-0.59)
yc	-0.23	-0.25	-0.16	-0.15	-0.24	-0.23	-0.16	-0.07
	(-1.23)	(-0.92)	(-1.13)	(-0.70)	(-1.28)	(-0.81)	(-1.15)	(-0.39)
public capital	-0.10	-0.15			-0.11	-0.10		
	(-0.62)	(-0.64)			(-0.57)	(-0.38)		
gov inves			-1.02	-0.97			-1.53*	-1.53
			(-1.39)	(-0.87)			(-1.77)	(-1.37)
Observations	1406	1406	1406	1406	1406	1406	1406	1406
No of countries	38	38	38	38	38	38	38	38

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects. ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

the average coefficient obtained by Galstyan and Velic (2018). All in all, a 10% increase in government consumption expenditures is associated with a rise of about 0.21% for the non-tradable relative price (in mean).

Examining the gross public debt, we fail to find a significant effect on the relative price of non-traded to traded. Our findings are in contradiction with Galstyan and Velic (2018) who obtain a positive effect for this variable over a sample of advanced economies. In their theoretical framework, public debt influences the relative non-traded to traded goods price through a rise in labor taxation, which affects both the non-traded and traded labor supplies. Structural macroeconomic differences between industrialized and EMDEs help to explain our results. Indeed, the transmission channel is likely to be not working for EMDEs. The main financing source for higher debt for the least advanced economies is probably not labor taxation because of the existence of high costs associated with tax collections. Government budget deficits in these economies are more likely to be financed through money creation for some economies rather than higher labor taxation, explaining probably our findings.

Moving to the trade balance, this variable is significant at the 1% level in half of our specifications. Once government investment is not included in our specifications, we display the significant negative sign expected by the study from Lane and Milesi-Ferretti (2002). Our estimated coefficients range from a low of 0.38 to a high of 0.62. Moreover, our estimates are lower than the estimations provided by Lane and Milesi-Ferretti (2002) who obtain a coefficient equals to -1.17.

Although government consumption expenditures play a key role in the non-traded to traded goods price's dynamic, the contemporaneous government investment is not significant. This result is robust to the variable considered as both the public capital stock and government investment are not correlated with our endogenous variables. In the best case, government investment is significant at the 10% level (columns (7)) with a negative sign. Our results illustrate thus the ambiguous sign of this variable. Following Galstyan and Lane (2009), our findings correspond to the case in which the productivity rises in tradable and non-tradable sectors are equal. Due to this equality, we have a neutral effect on the relative price of non-tradables. As previously argued, we also examine the effect of the lagged values of government investment and public capital stock. The results are available in Table B.3 in the Appendix. Even if lagged values are included as regressors, we continue to fail to find evidence for these determinants. These results may arise from the fact that monetary measures can be inappropriate to catch the trends in public capital stock (Calderón et al., 2015) as the link between spending and physical capital can depend on inefficiency and corruption surrounding the projects.

Higher GDP PC is expected to positively influence the relative price of non-tradables to tradables through a positive wealth effect. It should rise the demand for non-traded goods, pushing thus upwards the relative price of non-traded goods. Over the whole sample, we do not find evidence of a positive effect of this determinant across our 8 specifications. Two elements can help to understand such results: (i) the existence of a non-linear relationship between income and relative price of non-tradables, (ii) the inclusion of economies setting at different economic development stages. In fact, our two explanations are very closely linked one to each other. Indeed, the existence of this non-linear relationship can arise from the pooling of countries setting at different development stages. This intuition is confirmed by Hassan (2016). He shows that the Penn-effect depends on countries' income: low-income countries exhibit a negative price-income relationship while middle and high income countries display a positive one. One way to address this non-linearity is to include the squared GDP PC in our regressions. However, it is at the cost of high multicollinearity between our regressors. To overcome this drawback, we control for such heterogeneity in the development stage, by performing sub-sample regressions by splitting our sample into developing and emerging economies.

Tables 4 and 5 respectively display the estimation results for developing and emerging economies. Our sub-sample regressions exhibit some striking differences between the two group of countries. We continue to find evidence of a robust effect of government consumption expenditures on the non-traded to traded relative price. Examining developing economies (Table 4), this determinant is significant across all the specifications, while it turns out to be non significant in our last two specifications (columns (7.2) and (8.2) of Table 5) for emerging economies. The average marginal effect for this driver is "virtually" higher for developing countries than for emerging ones. Overall, for both groups of countries, the effect is larger than for advanced economies, perhaps due to the increasing trend in the government consumption expenditures in these economies.

Table 4: **Relative non-tradable price determinants: developing economies**

	def6_tv	**def6_tv**	**def6_tv**	**def6_tv**	**def6_f**	**def6_f**	**def6_f**	**def6_f**
	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)
	(1.1)	(2.1)	(3.1)	(4.1)	(5.1)	(6.1)	(7.1)	(8.1)
gov cons	1.65**	2.43**	1.77***	2.20**	1.79*	2.65**	1.81**	2.13*
	(2.26)	(2.29)	(2.81)	(1.96)	(1.95)	(2.15)	(2.41)	(1.92)
public debt	0.02	0.03	0.03	0.06	-0.01	0.02	0.01	0.02
	(0.25)	(0.30)	(0.33)	(0.43)	(-0.13)	(0.20)	(0.11)	(0.17)
tb	0.01	-0.06	-0.01	0.01	-0.15	-0.18	-0.23	-0.24
	(0.04)	(-0.21)	(-0.04)	(0.03)	(-0.58)	(-0.62)	(-0.77)	(-0.56)
yc	-0.28	-0.23	-0.21	-0.28	-0.27	-0.21	-0.17	-0.24
	(-1.56)	(-0.89)	(-1.19)	(-1.18)	(-1.47)	(-0.83	(-0.97)	(-1)
public capital	-0.14	-0.16			-0.12	-0.12		
	(-0.58)	(-0.52)			(-0.52)	(-0.49)		
gov inves			-0.71	-0.52			-1.00	-0.92
			(-0.92)	(-0.35)			(-0.99)	(-0.55)
Observations	703	703	703	703	703	703	703	703
No of countries	19	19	19	19	19	19	19	19

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects. ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

Table 5: **Relative non-tradable prices determinants: emerging economies**

	def6_tv	**def6_tv**	**def6_tv**	**def6_tv**	**def6_f**	**def6_f**	**def6_f**	**def6_f**
	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)
	(1.2)	(2.2)	(3.2)	(4.2)	(5.2)	(6.2)	(7.2)	(8.2)
gov cons	1.55**	1.86**	1.81**	2.07*	1.62**	1.71**	1.68	1.95
	(2.35)	(2.33)	(2.01)	(1.85)	(2.13)	(2.06)	(1.62)	(1.50)
public debt	-0.22*	-0.18	-0.32**	-0.28	-0.25*	-0.27*	-0.36**	-0.35
	(-1.83)	(-1.13)	(-2.13)	(-1.27)	(-1.79)	(-1.93)	(-2.12)	(-1.52)
tb	-0.35	-0.19	-0.5***	-0.34	-0.37	-0.12	-0.45**	-0.24
	(-1.59)	(-0.61)	(-2.63)	(-1.21)	(-1.42)	(-0.33)	(-2.05)	(-0.96)
yc	0.21**	0.32	0.30**	0.32	0.16	0.31	0.300**	0.29
	(2.13)	(1.57)	(2.20)	(1.55)	(1.51)	(1.45)	(2.04)	(1.45)
public capital	-0.22	-0.16			-0.28	-0.14		
	(-1.17)	(-0.44)			(-1.33)	(-0.34)		
gov inves			-0.61	-0.96			-0.49	-0.92
			(-0.57)	(-0.52)			(-0.46)	(-0.49)
Observations	703	703	703	703	703	703	703	703
No of countries	19	19	19	19	19	19	19	19

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects. ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

Furthermore, the results obtained for developing countries are quite in line with those of the whole sample, except the trade balance which has no effect. All in all, our results show that the great majority of determinants evidenced for advanced economies are clearly inadequate for developing market economies. The analysis of emerging markets economies displays findings contrasting with the whole sample results. In more than half of our specifications more or less strong support is provided for the gross government public debt. Contrary to advanced economies where a rise in gross public debt is associated with higher non-traded to traded relative price (Galstyan and Velic, 2018), we observe opposite results. Indeed, columns (1.2), (3.2), (5.2), (6.2) and (7.2) display a negative coefficient for this determinant. Such results would be consistent with the verification of the assumption of a tradable sector output more labor intensive than the non-tradable one. Thus a rise in gross public debt leads to a reduction of labor supply higher in the tradable sector than in the non-tradable.[33] Hence, as non-tradable price decrease is lower than the one experienced by the tradable sector, the relative price of the non-traded to traded sector shrinks. The estimated coefficients range from a low of 0.22 to a high of 0.36. Our estimated marginal effects are quite higher than the ones obtained by Galstyan and Velic (2018). Furthermore, we get weakly significant negative coefficients for the trade balance across our different specifications. Moving to GDP PC, some evidence of a positive effect is provided by Table 5 (see columns (1.2), (3.2) and (7.2)). Considering columns (3.2) and (7.2), a 1% increase in GDP PC is associated with a rise in the relative price of non-tradables of about 0.30%. This elasticity is consistent with non-linearity in the price-income relationship as higher elasticity is obtained for advanced economies. Indeed, Lane and Milesi-Ferretti (2002) and Galstyan and Lane (2009) respectively obtain elasticities of 0.56 and 0.83. Finally, the potential sample heterogeneity does not affect the government investment and public capital stock effects. Indeed, Tables 4 and 5 show no significant effect of both the public capital stock and government investment. As previously, we address the absence of potential simultaneous effect of these variables by using their lagged values. Tables B.4 and B.5 in the Appendix display our estimation results. Even if we consider this issue, we continue to fail to find evidence of significant impact of government investment and public capital stock.

6.4 Internal and external versions of the BS hypothesis: empirical results

Table 6 displays the estimation results for the investigation of the internal version of the BS hypothesis, using GDP PW as a proxy for the productivity differential. As a reminder, following this hypothesis, a higher productivity differential between the traded and non-traded sector is expected to rise the relative price of non-traded to traded. Columns (5.3) to (8.3) show that GDP PW does not affect the dynamic of the relative price of non-traded to traded. We can expect this finding is due to the inclusion of government investment in the specification. Indeed, as discussed previously this variable is expected to work through a productivity channel, leading to difficulty in disentangling the "pure" productivity differential effect. However, the non inclusion of this variable does not affect our previous results as illustrated by columns (1.3) to (4.3). All in all, GDP PW does not allow us to conclude to the existence of an internal version of the BS hypothesis.

[33] This assumption is hardly falsifiable because of data availability. Indeed, sectoral labor decomposition statistics start in 1991, while our panel begins in 1980.

Table 6: **Internal version of the bs hypothesis: gdp pw**

	def6_tv	**def6_tv**	**def6_f**	**def6_f**	**def6_tv**	**def6_tv**	**def6_f**	**def6_f**
	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)
	(1.3)	(2.3)	(3.3)	(4.3)	(5.3)	(6.3)	(7.3)	(8.3)
gdp pw	-0.18	-0.17	-0.20	-0.17	-0.14	-0.07	-0.12	0.01
	(-1.30)	(-0.90)	(-1.33)	(-0.97)	(-0.90)	(-0.28)	(-0.85)	(0.06)
gov cons	1.91***	2.04**	2.35***	2.38**	1.98***	2.16***	2.36***	2.48***
	(2.81)	(2.49)	(2.68)	(2.36)	(2.79)	(3.01)	(2.78)	(2.89)
public debt	-0.04	-0.06	-0.12	-0.09	-0.08	-0.06	-0.13	-0.08
	(-0.80)	(-0.46)	(-1.20)	(-0.75)	(-0.73)	(-0.4)	(-1.30)	(-0.57)
tb	-0.30*	-0.31	-0.31	-0.38	-0.31	-0.20	-0.33	-0.30
	(-1.66)	(-1.63)	(-1.19)	(-1.65)	(-1.48)	(-0.87)	(-1.44)	(-1.15)
gov inves					-1.03	-0.90	-1.59*	-1.60
					(-1.40)	(-0.84)	(-1.90)	(-1.57)
Observations	1406	1406	1406	1406	1406	1406	1406	1406
No of countries	38	38	38	38	38	38	38	38

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects. ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

The consideration of the labor productivity differential between the traded and non-traded sectors is maybe more adequate to investigate this hypothesis. The results for the investigation of the internal version of the BS hypothesis performed with the labor productivity differential between tradable and non-tradable sector is available in Table 7 below.

Columns (1.4) to (4.4) display the results for the productivity differential with time-varying weights, while fixed weights' findings are available in columns (5.4) to (8.4). Contrasting with the results displayed in Table 6, our measure of productivity differential is significant across 6 of our 8 specifications. It is also robust to the aggregation scheme used to regroup the sectors between tradable and non-tradable, even if the elasticities exhibit some variations. Examining columns (1.4) to (4.4), our results show that a 1% rise in the productivity differential is associated with an increase of 0.15 to 0.26% of the relative price of the non-traded to traded goods. Columns (5.4) to (8.4) confirm the previous findings. However, the elasticities obtained are much higher if we consider specifications (5.4) and (7.4). They range from a low of 0.23 to a high of 0.35. Hence, a higher effect of productivity differential is evidenced with a fixed weighting scheme. Using time-varying weights mitigates thus the productivity effect as another source of variation is included in our productivity differential variable. Furthermore, the inclusion of government investment in the specification seems to be a clearly inadequate choice to study the internal version of the BS hypothesis. Indeed, once this variable is included, the productivity differential significance vanishes (see Table B.6). It is in line with Galstyan and Lane's (2009) findings as they show that government investment significantly affects the relative sectoral productivity either positively or negatively. Hence, once this variable is controlled for, the effect of productivity differential is no longer detectable. To sum up, our empirical investigation shows some

Table 7: Internal version: productivity differential

	def6__tv	**def6__tv**	**def6__f**	**def6__f**	**def6__tv**	**def6__tv**	**def6__f**	**def6__f**
	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)
	(1.4)	(2.4)	(3.4)	(4.4)	(5.4)	(6.4)	(7.4)	(8.4)
prod__diff__tv	0.16**	0.21	0.15*	0.26**				
	(1.98)	(1.41)	(1.65)	(2.21)				
prod__diff__fixed					0.32***	0.17	0.35***	0.23*
					(3.15)	(1.28)	(2.84)	(1.87)
gov cons	0.94	1.58	1.14	2.35*	0.95	1.16	1.22	1.40
	(1.52)	(1.10)	(1.63)	(1.68)	(1.30)	(0.60)	(1.34)	(0.80)
public debt	-0.02	-0.02	-0.01	-0.05	0.03	0.23	0.05	0.09
	(-0.03)	(-0.04)	(-0.05)	(-0.12)	(0.37)	(1.00)	(0.21)	(0.53)
tb	-0.59	-0.99	-0.54	-0.60	0.76**	-0.51	-0.75*	-0.48
	(-1.55)	(-1.25)	(-1.54)	(-1.30)	(-2.30)	(-1.09)	(-1.83)	(-1.04)
yc	0.08	0.06	0.13	-0.27	-0.03	-0.03	0.02	-0.32
	(0.30)	(0.06)	(0.46)	(-0.22)	(-0.09)	(-0.09)	(0.06)	(-0.91)
Observations	988	988	988	988	988	988	988	988
No of countries	38	38	38	38	38	38	38	38

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects. ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

support in favor of the internal version of the BS hypothesis. Examining GDP PW does not allow us to conclude to the existence of a positive relationship between productivity and the relative price of non-traded to traded goods for EMDEs. The use of the labor productivity differential allows us to confirm this internal version although it is sensitive to the specification. Investigation of this hypothesis for EMDEs is a hard task suffering from two major caveats. Although productivity differential between the traded and non-traded sector is available, it is perhaps difficult to detect a long-run relationship because of rather short time-series span (26 years, here) and availability of few sectors.

Our empirical investigation also aims to examine the robustness of the external version of the BS effect to the proxies. Table 8 below displays the corresponding estimation results. Examining the REER determinants, we find a significantly negative effect of the trade openness ratio on the REER in the majority of our specifications. The estimated marginal effect ranges from -0.27 to -0.43. As expected, higher trade openness is associated with a more depreciated REER as it lowers the domestic price level. The relevance of this determinant for the REER dynamic of EMDEs is thus confirmed by our results. Examining the Terms Of Trade, we fail to find evidence of a robust effect for this determinant. It is only weakly significant in one of our ten specifications (column (3.5)). The positive sign for this specification suggests that the income effect dominates the substitution effect. From a theoretical perspective, the non-significance of TOT can be explained by the non-dominance of either the income or substitution effects. Our estimations confirm the relevance of government consumption expenditures for EMDEs. Once the six-sector value-added is not examined, we find evidence of a significant effect of the government consumption expenditures (columns (5.5) to (10.5)), confirming the Froot-Rogoff effect. Considering column (5.5), a 10% increase in government consumption expenditures raises the REER by 0.145%. Our estimated marginal effects are lower than the ones obtained by Ricci et al. (2013). Indeed, for a sample of 27 emerging markets economies, they get a marginal effect of about 0.386%. Finally, we observe strong significance and a positive sign for the fertility rate, in line with Rose et al. (2009) who show that lower fertility rate depreciates the exchange rate as the young consumption is biased towards non-traded goods.

According to the external version of the BS hypothesis, higher relative price of the non-traded to traded goods should appreciate the REER. To investigate this hypothesis, we rely on three and six sectors' value-added deflators. As can be seen from Table 8, the external version of the BS hypothesis is strongly verified for EMDEs as illustrated by columns (1.5) to (8.5). This finding is not affected by the difference in the price deflator and by the weighting system used as our proxies are always significant. At least a 1% increase in this proxy appreciates the REER of about 0.18% to 0.32%. Moreover, the marginal effects of the relative price of non-traded to traded goods are lower for the three-sectors' value-added deflators. Indeed, the estimated coefficients for these variables range from 0.180 to 0.257 (columns (5.5) to (8.5)), while the elasticities go from 0.275 to 0.319 for the six-sectors' value-added deflators. Another interesting result is that our findings are not affected by the weighting system considered for the computation. For the sake of comparison, we also examine the effect of GDP PC. Our results show no significant effect of this variable on the REER. We thus confirm previous findings, even if CSD is taken into account, GDP PC appears to be a misleading proxy for the BS effect in EMDEs. All in all, our results provide very strong support for the BS hypothesis. We show that this finding is robust to measurement issues using four alternative proxies for the relative price of non-traded to traded goods.

Table 8: External version of the BS hypothesis

	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)	CS-DL(1)	CS-DL(2)
	(1.5)	(2.5)	(3.5)	(4.5)	(5.5)	(6.5)	(7.5)	(8.5)	(9.5)	(10.5)
def6__tvr	0.298***	0.296*								
	(2.886)	(1.895)								
def6__fr			0.275**	0.319***						
			(2.274)	(2.698)						
def3__tvr					0.248***	0.180*				
					(3.604)	(1.761)				
def3__fr							0.257***	0.199**		
							(3.891)	(1.973)		
yc__r									0.086	0.411
									(0.549)	(0.873)
open	-0.42***	-0.30***	-0.42***	-0.30***	-0.17	-0.14	-0.16	-0.11	-0.43***	-0.27**
	(-5.25)	(-2.73)	(-4.67)	(-3.75)	(-1.31)	(-0.82)	(-1.23)	(-0.61)	(-4.30)	(-1.93)
tot	0.08	0.08	0.09***	0.07	-0.02	-0.08	-0.01	-0.08	0.06	-0.02
	(1.61)	(0.97)	(1.82)	(0.98)	(-0.31)	(-1.59)	(-0.16)	(-1.53)	(1.10)	(-0.16)
gov cons	0.57	0.89	0.57	0.88	1.45***	2.37***	1.52***	2.33***	1.32***	2.18***
	(1.43)	(1.39)	(1.19)	(1.35)	(4.53)	(5.04)	(4.61)	(5.18)	(3.22)	(2.73)
fertility	0.31**	0.34*	0.33**	0.36**	0.36*	0.49***	0.36*	0.54**	0.34**	0.21
	(2.47)	(1.89)	(2.10)	(2.02)	(1.91)	(2.07)	(1.93)	(2.11)	(2.19)	(0.88)
Observations	1406	1406	1406	1406	1406	1406	1406	1406	1406	1406
Number of countries	38	38	38	38	38	38	38	38	38	38

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects. ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

7 Conclusion

This paper has investigated the robustness of the BS effect on a sample of 38 EMDEs over the period 1980-2016 using five alternative measures. We address two crucial issues that have not been examined so far in the existing literature: (i) the measurement issues associated with the productivity differential and the relative price of non-traded to traded goods, (ii) accounting for the first time for CSD, using the CS-DL approach (Chudik et al., 2016).

We start by identifying the determinants of the relative price of non-tradables for EMDEs. Our findings strongly support the key role played by government consumption expenditures in the dynamic of the relative price of non-traded goods for both developing and emerging market economies over the period 1980-2016: higher government consumption expenditures are indeed associated with a rise in the relative price of non-tradables to tradables. The trade balance exerts a significant effect, but to a lesser extent. Finally, we show that higher wealth positively affects the non-traded to traded price for emerging market economies.

Investigating both the internal and external versions of the BS hypothesis, we show that while the internal version is not confirmed if we use GDP PW as a proxy for productivity differential, quite strong support is obtained with the labor productivity differential. A rise in this latter is associated with an increase in the relative price of non-traded to traded goods of 0.15% to 0.35%. The verification of the internal version thus depends crucially on the proxy chosen. Moving to the external version, we find strong support for the BS effect. An increase in the non-traded to traded price is associated with an appreciation of the REER. Although the estimated elasticity is more or less affected by the proxies, we find evidence of a significant positive effect if three-sectors or six-sectors deflators are used. An increase in the relative price of the non-tradables to tradables is associated with a REER appreciation of 0.180% to 0.319%. We also show that GDP PC is not relevant to test for the external version of the BS hypothesis.

Our results show that the proxy choice is of great importance in the validation of the BS effect as it may affects the conclusion reached. Hence, equilibrium exchange rate estimation for EMDEs must be performed using various proxies to ensure that the value obtained is not drove by an inadequate choice. As illustrated by our findings, the Balassa-Samuelson effect works through the productivity differential between the traded and non-traded sectors. However, identifying the potential source of this increase has not yet been a concern in the literature. Although Du et al. (2013) constitute a first attempt in this direction by considering the impact of transport infrastructure on the RER, transport infrastructure represents only one dimension of an economy network. Accounting for the multidimensional characteristic (including transportation, communication, and energy and utilities dimension) of economies' infrastructure would be a promising extension of the present paper. Another interesting avenue for future research would be to examine the transmission channel of infrastructure on the RER through productivity in the traded and non-traded sectors.

*Appendix

A Sample and data sources

Table A.1: Sample of countries

Emerging	Geographical zone	Developing	Geographical zone
Chile	South America	Bangladesh	Southern Asia
Colombia	South America	Benin	Western Africa
Costa Rica	Central America	Bolivia	South America
Dominican republic	Caribbean	Botswana	Southern Africa
Ecuador	South America	Burkina Faso	Western Africa
El Salvador	Central America	Cameroon	Middle Africa
Guatemala	Central America	Central african republic	Middle Africa
India	Southern Asia	Comoros	Middle Africa
Korea republic	Eastern Asia	Congo	Middle Africa
Malaysia	South-Eastern Asia	Gabon	Middle Africa
Mexico	Central America	Honduras	Central America
Morocco	Northern Africa	Kenya	Middle Africa
Pakistan	Southern Asia	Malawi	Eastern Africa
Panama	Central America	Mali	Western Africa
Philippines	South-Eastern Asia	Mauritius	Eastern Africa
South Africa	Southern Africa	Rwanda	Eastern Africa
Sri Lanka	Southern Asia	Senegal	Western Africa
Thailand	South-Eastern Asia	Togo	Western Africa
Uruguay	South America	Uganda	Eastern Africa

Table A.2: Data sources

Primary	Abbreviation	Data sources	Comments
External balance on goods and services	tb	WDI	Expressed as share of GDP
Gross public debt	public debt		Expressed as share of GDP
Government consumption expenditures	gov cons	WDI	Expressed as share of GDP
Public capital stock	public capital	IMF, fiscal affair department	Expressed as share of GDP
Government investment	gov inves	IMF, fiscal affair department	Expressed as share of GDP
Gross Domestic Product per Capita	yc	WDI	Expressed in constant PPP
Terms of Trade	tot	WDI	Ratio of export prices to import prices
Net Foreign asset	nfa	Lane and Milesi-Ferretti (2007)	Expressed as share of GDP
Trade openess	open	WDI	Sum of exports and imports as share of GDP
Fertility rate	fertility	WDI	Births per woman
Investment	investment	WDI	Expressed as share of GDP
Foreign Direct Investment	fdi	WDI	Net inflows, Expressed as share of GDP
Real Effective Exchange Rate	reer	EQCHANGE	REER based on 186 trade partners using the 2008-2012 trading weights
GDP per workers		PWT 9.0, ILOSTAT	GDP in constant 2011 PPP U.S divided by total employment (number of engaged people)
Value added per worker (agriculture)	vw_agr	ILOSTAT	Value added per unit of input measured in constant 2010 us dollars
Value added per worker (industry)	vw_indus	ILOSTAT	Value added per unit of input measured in constant 2010 us dollars
Value added per worker (services)	vw_serv	ILOSTAT	Value added per unit of input measured in constant 2010 us dollars
Values added (agriculture)		WDI	Value added of agriculture input measured in constant/current 2010 us dollars
Values added (industry)		WDI	Value added of industry input measured in constant/current 2010 us dollars
Values added (services)		WDI	Value added of services measured in constant/current 2010 us dollars
Values added of division 01-05		UNCTAD	Value added of the sector 01-05 in constant/current 2010 us dollars
Values added of division 10-41		UNCTAD	Value added of of the sector 10-41 measured in constant/current 2010 us dollars
Values added of division 45		UNCTAD	Value added of the sector 45 measured in constant/current 2010 us dollars
Values added of division 50-55		UNCTAD	Value added of the sector 50-55 measured in constant/current 2010 us dollars
Values added of division 60-64		UNCTAD	Value added of the sector 60-64 measured in constant/current 2010 us dollars
Values added of division 65-99		UNCTAD	Value added of the sector 65-99 measured in constant/current 2010 us dollars
GDP PC F1	yc_r	CEPII EQCHANGE	Deviation of countries GDP PC from trading partners.
Commercial weights	w_ijt	CEPII EQCHANGE	Time-invariant weighting scheme representative of foreign trade between 2008-2012

Note: CEPII=Centre d'études prospectives et d'informations internationales; WDI=World Development Indicator (World Bank); PWT=Penn World Table; ILOSTAT=International Labour Organization Statistics; UNCTAD=United Nations Conference on Trade and Development; IMF=International Monetary Fund; GDP= Gross Domestic Product; PC= Per capita; PPP= Purchasing Power Parity.

B Additional results

Table B.1: Cross section dependence tests: results

def6__tv	**def6__f**	**tb**	**public debt**	**gov cons**	**gov inves**
29.100	29.346	32.044	27.990	35.530	26.176
(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
public capital	**yc**	**prod__diff__fixed**	**prod__diff__tv**	**reer**	**def6__fr**
18.860	5.391	16.598	18.046	32.550	8.009
(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
def6__tvr	**def3__fr**	**def3__tvr**	**open**	**fertility**	**ycr**
19.048	19.717	19.048	14.155	7.686	18.571
(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
tot	**gdp pw**				
16.757	10.308				
(0.00)	(0.00)				

Note: Pesaran (2004)'s test of cross-section independence. Null hypothesis of cross-section independence. P-values are given in parentheses.

Table B.2: Panel unit root tests: supplementary results

Specification	Test	Stat	tb	public debt	Δ public debt	gov cons	Δ gov cons	gov inves
Trend and constant	CIPS		-2.975***	-2.554*	-4.964***	-2.609**		-3.521***
Constant	CIPS		-2.607***	-2.002	-4.929***	-2.119**		-2.54***
	BCIS	Z	-2.036**	-2.515***	-4.35***	-0.396	-4.61***	-1.066
Level mean shift	BCIS	Pm	4.074***	3.293***	24.22***	-0.746	31.243***	2.637***
	BCIS	P	126.23***	116.602***	374.664***	66.794	461.196***	108.516***
	BCIS	Z	-1.39	0.414	-3.79***	2.009**	-4.938***	5.446***
Trend level shift	BCIS	Pm	2.364**	0.969	16.45***	-0.96	23.383***	0.447
	BCIS	P	105.149**	87.949	278.84***	64.044	364.296***	81.511

			public capital	GDP PC	Δ GDP PC	def6_tv	Δ def6_tv	
Trend and constant	CIPS		-2.600*	-2.306	-5.282***	-2.556*		
Constant	CIPS		-2.239***	-1.813	-5.021***	-2.362***		
	BCIS	Z	1.972**	4.44***	-4.963***	2.618***	-4.586***	
Level mean shift	BCIS	Pm	-2.276**	-2.407***	20.080***	-1.866***	28.511***	
	BCIS	P	47.928	46.319	323.574***	52.99	427.18***	
	BCIS	Z	2.528***	2.34	-4.381***	-1.319	-4.938***	
Trend level shift	BCIS	Pm	-1.674**	0.284	20.038***	-1.478	23.383***	
	BCIS	P	55.361	79.503	323.057***	57.773	364.296***	

			def6_f	Δ def6_f	def6_tvr	Δ def6_tvr	def6_fr	Δ def6_fr
Trend and constant	CIPS		-2.766***		-2.488	-5.373***	-2.288	-5.531***
Constant	CIPS		-2.490***		-2.011	-5.292***	-2.226	-5.534***
	BCIS	Z	4.709***	-4.631***	2.700***	2.062**	1.101	-4.667***
Level mean shift	BCIS	Pm	-2.530***	29.472***	0.322	-0.365	-0.819	29.813***
	BCIS	P	44.799	439.364***	79.975	71.498	65.898	443.561***
	BCIS	Z	-1.035	-4.886***	-0.606	-5.032***	0.159	-5.241***
Trend level shift	BCIS	Pm	0.21	24.47***	0.7.2	24.274***	-0.918	23.995***

Table B.2: (continued)

	BCIS	P	78.6	377.695***	85.036	375.272***	64.678	371.831***
			def3__tvr	**Δdef3__tvr**	**def3__fr**	**Δdef3__fr**	**reer**	**Δreer**
Trend and constant	CIPS		-2.488		-2.499	-5.308***	-2.467	-5.416***
Constant	CIPS		-2.228***		-2.154**	-5.265***	-2.037	-5.192***
	BCIS	Z	2.062**	-4.686***	1.148	-4.631	5.242***	-4.486***
Level mean shift	BCIS	Pm	-0.365	27.633***	0.519	29.517***	-1.649	25.28***
	BCIS	P	71.498	416.687***	82.406	439.916***	55.66	387.682***
	BCIS	Z	4.069***	-5.346***	-0.698	-5.354***	1.699	-5.191***
Trend level shift	BCIS	Pm	-2.052**	24.872***	-1.602	25.703***	-0.811	24.825***
	BCIS	P	50.699	382.654***	56.238	392.9***	65.99	382.063***
			open	**Δopen**	**fertility**	**Δfertility**	**tot**	**Δtot**
Trend and constant	CIPS		-2.101	-5.603***	-2.473	-3.494***	-2.21	-5.670***
Constant	CIPS		-1.982	-5.495***	-2.183**	-3.553***	-1.725	-5.497***
	BCIS	Z	-1.452	-4.408***	-2.933***	29***	-2.236**	-4.397***
Level mean shift	BCIS	Pm	2.442***	26.713***	0.112	-2.89***	-1.906**	23.718***
	BCIS	P	106.110**	405.351***	77.39	15.7	52.49	368.420***
	BCIS	Z	-1.033	-4.858***	43.5***	68.13***	-0.332	-5.046***
Trend level shift	BCIS	Pm	1.338	22.865***	-6.022***	-5.575***	-1.572	23.410***
	BCIS	P	92.496	357.898***	1.753	7.261	95.386	364.627***
			prod diff__tv	**Δprod diff__tv**	**prod diff__fixed**	**Δprod diff__fixed**	**gdp pw**	**Δ gdp pw**
Trend and constant	CIPS		-2.67**	-4.483***	-2.498	-4.768***	-2.341	-5.162***
Constant	CIPS		-1.922	-4.338***	-1.865	-4.616***	-1.841	-5.087***
	BCIS	Z	5.191***	-4.467***	6.07***	-4.49***	3.950***	-4.581***
Level mean shift	BCIS	Pm	-2.191**	15.919***	-1.45	20.77***	-0.589	25.34***
	BCIS	P	48.982	272.270***	58.06	332.15***	68.736	388.418***
	BCIS	Z	1.634	-4..650***	1.32	-4.62***	3.314***	-5.260***
Trend level shift	BCIS	Pm	-1.362	12.179***	-1.44	14.84***	1.268	24.826***
	BCIS	P	59.199	226.155***	58.2	258.911***	91.637	382.086***

Table B.2: (continued)

			yc_r	Δ yc_r
Trend and constant	CIPS		-2.438	-5.013***
Constant	CIPS		-1.652	-4.783***
	BCIS	Z	2.447***	-4.710***
Level mean shift	BCIS	Pm	-1.781**	28.275***
	BCIS	P	54.034	424.601***
	BCIS	Z	10.042	-1.441
Trend level shift	BCIS	Pm	-0.611	19.210***
	BCIS	P	68.458	312.846***

Note: Critical values for the trend and constant (resp. constant) model at 10%, 5% and 1% are respectively -2,55, -2,60 and -2,72 (resp. -2,05, -2,11 and -2,23). Critical values for the Z and Pm statistics at the 1% and 5% level are respectively 2,33 and 1.645. Critical values for P at the 1% level and 5% level are 97.35 and 107.58. Maximum number of factors allowed equals to 5. Bayesian information criterion in Bai and Ng (2002) used to estimate the number of factors.

Table B.3: Whole sample: alternative specifications

	def6_tv CS-DL(1)	**def6_tv** CS-DL(2)	**def6_tv** CS-DL(1)	**def6_tv** CS-DL(2)	**def6_f** CS-DL(1)	**def6_f** CS-DL(2)	**def6_f** CS-DL(1)	**def6_f** CS-DL(2)
gov cons	0.95	1.38	1.78**	1.63	1.44**	1.52	2.19**	2.12*
	(0.84)	(1.33)	(2.34)	(1.60)	(2.03)	(1.49)	(2.46)	(1.74)
public debt	-0.23	-0.09	-0.09	-0.03	-0.11	-0.10	-0.11	-0.03
	(-1.35)	(-0.75)	(-0.90)	(-0.20)	(-1.22)	(-0.71)	(-1.10)	(-0.20)
tb	-0.47	-0.3*	-0.31	-0.27	-0.41***	-0.41**	-0.26	-0.29
	(-0.65)	(-1.77)	(-1.48)	(-1.13)	(-2.56)	(-2.05)	(-0.96)	(-1.04)
yc	-0.24	-0.24	-0.16	-0.15	-0.18	-0.23	-0.15	-0.09
	(-0.48)	(-0.66)	(-1.05)	(-0.60)	(-1.05)	(-0.65)	(-0.99)	(-0.42)
lagged public capital	0.12	-0.007			-0.10	0.03		
	(0.24)	(-0.02)			(-0.61)	(0.12)		
lagged gov inves			-0.44	0.04			-0.78	-0.22
			(-0.89)	(0.04)			(-1.51)	(-0.29)
Observations	1370	1370	1370	1370	1370	1370	1370	1370
Number of countries	38	38	38	38	38	38	38	38

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

Table B.4: Developing economies: alternative specifications

	def6_tv CS-DL(1)	**def6_tv** CS-DL(2)	**def6_tv** CS-DL(1)	**def6_tv** CS-DL(2)	**def6_f** CS-DL(1)	**def6_f** CS-DL(2)	**def6_f** CS-DL(1)	**def6_f** CS-DL(2)
gov cons	1.46**	2.45**	1.68**	2.12**	1.7**	2.52**	1.91**	2.3**
	(2.25)	(2.01)	(2.51)	(2.04)	(1.98	(2.05)	(2.42)	(2.32)
public debt	-0.03	-0.03	0.01	0.04	-0.03	-0.01	-0.01	0.04
	(-0.33)	(-0.33)	(0.11)	(0.24)	(-0.33)	(-0.13)	(-0.13)	(0.24)
tb	-0.19	0.13	-0.10	0.10	-0.34	-0.08	-0.24	-0.15
	(-0.66)	(0.45)	(-0.40)	(0.24)	(-1.10)	(-0.17)	(-0.86)	(-0.27)
yc	-0.27	-0.26	-0.25	-0.31	-0.25	-0.25	-0.18	-0.24
	(-1.43)	(-0.77)	(-1.38)	(-1.25)	(-0.09)	(-0.75)	(-1.10)	(-1.01)
lagged public capital	-0.08	-0.12			-0.32	-0.13		
	(-0.28)	(-0.20)			(-1.33)	(-0.27)		
lagged gov inves			-0.35	0.24			-0.56	0.24
			(-0.416)	(0.150)			(-0.60)	(0.15)
Observations	684	684	684	684	684	684	684	684
Number of countries	19	19	19	19	19	19	19	19

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

Table B.5: Emerging economies: alternative specifications

	def6_tv CS-DL(1)	**def6_tv** CS-DL(2)	**def6_tv** CS-DL(1)	**def6_tv** CS-DL(2)	**def6_f** CS-DL(1)	**def6_f** CS-DL(2)	**def6_f** CS-DL(1)	**def6_f** CS-DL(2)
gov cons	1.65*	2.03*	1.43*	2.02***	1.00	0.48	0.84	-0.07
	(1.81)	(1.74)	(1.64)	(2.66)	(0.75)	(0.29)	(0.64)	(-0.04)
public debt	-0.36**	-0.31	-0.24*	-0.19	-0.19	-0.26	-0.19	-0.29
	(-2.12)	(-1.63)	(-1.71)	(-1.36)	(-1.58)	(-1.63)	(-1.36)	-(1.32)
tb	-0.48**	-0.44	-0.27	-0.29	-0.60	-0.07	-0.31	-0.08
	(-2.18)	(-1.16)	(-1.13	(-1.16)	(-1.94)	(-0.16)	(-0.94)	(-0.14)
yc	0.29*	0.33	0.18	0.35***	0.18***	0.11	0.25***	0.26***
	(1.87)	(1.55)	(1.30)	(3.06)	(3.11)	(1.37)	(4.54)	(3.47)
lagged public capital	-0.10	-0.10			-0.03	-0.02		
	(-0.36)	(-0.29)			(-0.10)	(-0.05)		
lagged gov inves			-1.06	-1.73			0.23	-0.13
			(-0.79)	(-0.75)			(0.31)	(-0.09)
Observations	684	684	684	684	684	684	684	684
Number of countries	19	19	19	19	19	19	19	19

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects. ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

Table B.6: Internal version of the BS hypothesis: alternative specification

	def6__tv CS-DL(1)	**def6__tv** CS-DL(2)	**def6__tv** CS-DL(1)	**def6__tv** CS-DL(2)	**def6__f** CS-DL(1)	**def6__f** CS-DL(2)	**def6__f** CS-DL(1)	**def6__f** CS-DL(2)
gov cons	0.60	1.09	1.22	-1.46	1.69	2.20	1.02	0.83
	(0.43)	(0.01)	(0.97)	(-0.10)	(0.85)	(0.14)	(0.52)	(0.01)
gross debt	-0.06	0.20	0.03	-0.03	0.04	0.13	-0.06	0.05
	(-0.46)	0.02)	(0.04)	(-0.01)	(0.15)	(0.06)	(-0.20)	(0.13)
tb	-0.59	0.28	-0.78	-1.59	-0.70	0.72	-0.41	-0.21
	(-1.59)	(0.01)	(-1.56)	(-0.30)	(-1.01)	(0.11)	(-0.55)	(-0.01)
yc	-0.03	-0.03	-0.13	0.03	-0.07	-0.04	0.06	-0.02
	(-0.09)	(-0.01)	(-0.64)	(0.10)	(-0.25)	(-0.12)	(0.03)	(-0.01)
gov inves	-1.25	-1.00	-1.13	-1.30	-1.89	-1.85	-1.71	-1.41
	(-0.58)	(-0.01)	(-0.73)	(-0.03)	(-0.66)	(-0.05)	(-0.47)	(-0.01)
prod__diff__fixed			0.08	0.17	0.06	0.05		
			(0.67)	(0.13)	(0.31)	(0.03)		
prod__diff__tv	0.11	0.09					0.05	0.12
	(0.65)	(0.02)					(0.15)	(0.03)
Observations	1408	1408	1408	1408	1408	1408	1408	1408
Number of countries	38	38	38	38	38	38	38	38

Note: CS-DL(1) and CS-DL(2) indicate cross-sectionally distributed lag model with respectively one year and two years lags of the first-differenced regressors and contemporaneous cross-sectional mean value of the dependent variable and regressors (in level). Each regression includes country-specific fixed effects ***, **, and * denote the levels of statistical significance at 1, 5, and 10%, respectively. t-stat are reported in parentheses.

Table B.7: Six sectors' classification

Sector	Divisions
Agriculture, hunting, forestry, fishing	01-05
Mining, manufacturing, utilities	10-41
Construction	45
Wholesale, retail trade, restaurants and hotels	50-55
Transport, storage and communications	60-64
Other activities	65-99

Source: International Standard Industrial Classification of All Economic Activities (ISIC), Revision 3.1, United Nations.

C Methodology

Appendix C.A Cross-sectional dependence in panel data: issues and modelisation

Pesaran (2015) argues that: "ignoring cross-sectional dependence of errors can have serious consequences, and the presence of some form of cross-section correlation of errors in panel data applications in economics is likely to be the rule rather than the exception". This cross-sectional correlation of errors arises from omitted common effects such as aggregate demand and/or supply shocks. In our case, the existence of CSD is particularly important from the estimation and unit root testing perspectives. Indeed, if conventional estimators are employed, misleading inference[34] and even inconsistent estimations are likely to occur if the presence of CSD is not taken into account (Pesaran, 2015; Reese and Westerlund, 2016). In the context of panel unit root tests, the first generation of test is no longer suitable as they assume cross-sectionaly uncorrelated errors.[35] Indeed, Chang (2002) points out that in the case where such a hypothesis does not hold, the test distributions are no longer valid. To test for the presence of CSD,[36] Pesaran (2004) has proposed a simple test based on average of pairwise correlation of the residuals. We consider the following panel data model:

$$y_{i,t} = a_i + \beta_i' x_{i,t} + u_{i,t} \tag{C.1}$$

where $x_{i,t}$ is a k-dimensional vector of regressors, and a_i and β_i' are assumed to be fixed unknown coefficients. The null of cross-sectionally uncorrelated errors is defined by:

$$Ho : E(u_{i,t}u_{j,t}) = 0, \text{for all } t \text{ and } i \neq j \tag{C.2}$$

Consider the panel model described by equation (2.C.1), and let $\widehat{u_{it}}$ be the OLS estimator of u_{it} defined by:

$$\widehat{u_{i,t}} = y_{i,t} - \hat{a}_i - \hat{\beta}_i' x_{i,t} \tag{C.3}$$

$\widehat{a_i}$ and $\widehat{\beta_i'}$ being the OLS estimates of a_i and β_i', based on the T sample observations. Using the previous results, we estimate the pair wise correlation of the residuals, $\widehat{u_{i,t}}$ and $\widehat{u_{j,t}}$, for $i \neq j$:

$$\widehat{\rho_{i,j}} = \widehat{\rho_{j,i}} = \frac{\sum_{t=1}^{T} \widehat{u_{i,t}}\widehat{u_{j,t}}}{(\sum_{t=1}^{T} \widehat{u_{i,t}}^2)^{1/2} (\sum_{t=1}^{T} \widehat{u_{j,t}}^2)^{1/2}} \tag{C.4}$$

Denoting CD_p the cross-sectional dependence test statistics, computed as the average of the pairwise correlation of the residuals:

$$CD_p = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \widehat{\rho_{ij}} \right) \tag{C.5}$$

[34]One way to get ride of this problem is to rely on standard errors robust to CSD, as the one proposed by Driscoll and Kraay (1998).

[35]The first generation of panel unit root tests encompasses the following tests: Im et al. (2003); Choi (2001) and Maddala and Wu (1999). See Hurlin and Mignon (2005) for a survey on the panel unit root tests.

[36]To test for the presence of CSD in the context of ordered individuals, the Moran test (1948) is applicable.

Under the null of cross-sectionally uncorrelated errors, we get:

$$CD_p \rightarrow \mathcal{N}(0,1) \tag{C.6}$$

Once the presence of CSD has been tested for, it has to be control for. To provide a better understanding of the tests implemented in this paper, we present a convenient way used to modelise the CSD. To this aim, we set up a common factor models:

$$y_{i,t} = \gamma_{i,1} f_{1,t} + \gamma_{i,2} f_{2,t} + ... + \gamma_{i,k} f_{m,t} + e_{i,t} \tag{C.7}$$

which can also be written:

$$y_t = \Gamma f_t + e_t \tag{C.8}$$

where f_t=$(f_{1,t}, f_{2,t}, ..., f_{m,t})'$, e_t=$(e_{1,t}, e_{2,t}, ..., e_{N,T})'$, and Γ=$(\gamma_{i,j})$, for i=1,2,...,N, j=1,2,...,m is an N*m matrix of factor loadings. k=1,2,...,K.

The observed variable $y_{i,t}$ is thus explained in terms of m unobserved common factors, $f_t = (f_{1,t}, f_{2,t}, ...f_{m,t})$. These factors influence the dynamic of y_{it} through factor loadings $(\gamma_{i,1}, \gamma_{i,2}, ...\gamma_{i,k})$. As these factor loadings are heterogeneous, the different cross sections units are differently impacted by the common factors. Two main difficulties arise once a factor model is considered:

- determination of the "true" number of common factors, k
- estimation of these k unobserved common factors

The first issue is important for the empirical and theoretical validity of factor models (Bai and Ng, 2002). That is why, Bai and Ng (2002) propose different criteria to consistently estimate the number of factors. Their method consists to specify a penalty function which depends on both the cross sectional and time-series dimensions of the panel data. Prior to the specification of the penalty function is the estimation of the k unobserved factors and their heterogeneous factor loadings. The unobserved common factors are estimated relying on the asymptotic principal component estimator. Then, heterogeneous factor loadings are obtained by solving the following optimization problem:

$$V(k, \widehat{f_{m,t}}) = \underset{\gamma}{\text{Min}} (NT)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} (y_t - \gamma_{i,k} \widehat{f_{m,t}}) \tag{C.9}$$

where $\widehat{f_{m,t}}$ is the estimation of the unobserved common factors. As argued by Bai and Ng (2002), as the model is linear and the factors are observed, γ_{ik} can be estimated by applying ordinary least squares to each equation. Then, a loss function denoted PC(k) can be used to determine the k number of factors:

$$PC(k) = V(k, \widehat{f_{m,t}}) + kg(N,T) \tag{C.10}$$

where g(N,T) is the penalty function for overfitting.

However, the procedure proposed by Bai and Ng (2002) is only applicable in the presence of stationary data. To overcome this drawback, Bai and Ng (2004) propose the Panel Analysis of Nonstationarity in Idiosyncratic and Common components (PANIC) methodology. This approach consists to decompose the series into two unobserved components: idiosyncratic and common components. To deal with potential non stationary variables, the first difference is applied to the series. Then, the principal component method is applied to estimated the first-differenced common and idiosyncratic components. The level of the common and idiosyncratic components is then recovers through a difference reaccumulation approach.

Pesaran (2006) proposes an easier way to get ride of the two issues previously mentioned. He shows that an approximation of the linear unobserved common factors is consistently estimated by a cross-section average of the dependent variables and its first lagged difference. Moreover, this approach does not require the knowledge of the number of unobserved factors. It must only have a fixed number of factors. Rather than being mutually exclusive, the cross-section average (CA) augmentation and PANIC approaches can be combined in an efficient way (Reese and Westerlund, 2016).[37] Reese and Westerlund (2016) propose the PANICCA methodology which is PANIC applied on CA. This methodology exploits advantages from both approaches. The CA of Pesaran (2006) proposes an easy way to obtain consistent estimation of the unobserved factors, while the PANIC methodology is interesting as it allows inference based on normal distribution. That is why Reese and Westerlund (2016) develop CA based tests that support asymptotically normal inference.

Appendix C.B Panel unit root tests

Our presentation of the CIPS test starts by outlining a benchmark specification for the unit root hypothesis in a heterogeneous panel:

$$y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + \varepsilon_{it} \tag{C.11}$$

where ρ_i is the heterogeneous autoregressive parameter. We subtract each side of the equation by $y_{i,t-1}$, we get:

$$\Delta y_{it} = \alpha_i + \phi_i y_{i,t-1} + \varepsilon_{it} \tag{C.12}$$

where $\phi_i = (\rho_i - 1)$. Pesaran (2007) shows that cross-section dependency can be controled for through the add of the contemporaneous and lagged average of the endogenous variable to the regression. The test is named the

[37] See Urbain and Westerlund (2015) for a comparison between CA and PANIC methodologies.

cross-section augmented Dickey-Fuller CADF. We get the following equation:

$$\Delta y_{i,t} = \alpha_i + \phi_i y_{i,t-1} + b_i \overline{y_{t-1}} + c_i \Delta \overline{y_t} + \varepsilon_{it} \tag{C.13}$$

where $\overline{y_t} = N^{-1} \sum_{i=1}^{N} y_{it}$ and $\overline{\Delta y_t} = N^{-1} \sum_{i=1}^{N} \Delta y_{it}$. The set of hypotheses for the CIPS test is the following:

$$\begin{cases} Ho : \rho_i = 1 \, for \; i = 1, ..., N \\ H1 : \rho_i = 1 \, for \; i = 1, ..., N_1 \\ H1 : \rho_i < 1 \, for \; i = N_1, ..., N \end{cases} \tag{C.14}$$

The test statistic is obtained as the average of the N individual CADF t-statistics:

$$CIPS = \frac{1}{N} \sum_{i=1}^{N} CADF_i \tag{C.15}$$

where $CADF_i$ is the ADF statistics obtained for a cross section :i from the individual regression.[38] As shown by Pesaran (2007), the joint asymptotic limit distribution of the CIPS statistic is nonstandard and critical values for each values of N and T are provided. Although controlling for CSD, the CIPS test does not include the existence of structural breaks in its procedure. Failure to properly control for the existence of structural breaks in the testing procedure can lead to misleading conclusions regarding the order of integration of a time-series (Bai and Carrion-I-Silvestre, 2009). To circumvent this drawback, we thus use the BCIS test. Assuming that we test the null hypothesis of unit root for a variable X_t, we have the following set of hypotheses:

$$\begin{cases} \text{Ho} : \rho_i = 1 \text{ for } \; i = 1, ..., N \\ \text{H1} : |\rho_i| = 1 \text{ for some } \; i \end{cases} \tag{C.16}$$

Following Bai and Carrion-I-Silvestre (2009), we consider the following general panel model:

$$X_{i,t} = D_{i,t} + F_t' \pi_i + e_{i,t} \tag{C.17}$$

$$(I - L) F_t = C(L) u_t \tag{C.18}$$

$$(1 - \rho_i L) e_{i,t} = H_i(L) \epsilon_{i,t} \tag{C.19}$$

where $C(L) = \sum_{j=0}^{\infty} C_j L^j$ and $H_i(L) = \sum_{j=0}^{\infty} H_{i,j} L^j$. D_{it} stands for the deterministic part of the model and $e_{i,t}$ is the idiosyncratic error term. F_t is a vector that accounts for the common factors of the panel. Regarding the deterministic component $D_{i,t}$, two deterministic models are considered:

[38] It is worth mentioning that Pesaran et al. (2013) extend this approach to a multifactor error term structure. They show that a way to control for the cross-section dependence is to use the information contained in additional variables, which are assume to share the same common factor of the variable of interest. Then, the trick is to include in the ADF regression test the average of these variables.

$$D_{i,t} = \mu_i + \sum_{j=1}^{l_i} \theta_{i,j} DU_{i,j,t} \tag{C.20}$$

$$D_{i,t} = \mu_i + \beta_i t + \sum_{j=1}^{l_i} \theta_{i,j} DU_{i,j,t} + \sum_{k=1}^{m_i} \gamma_{i,k} DT_{i,k,t} \tag{C.21}$$

Bai and Carrion-I-Silvestre (2009) allow for l_i structural breaks affecting the mean and m_i structural breaks affecting the trend of the time-series.[39] The dummy variables are defined as $DU_{i,j,t}$=1 for t> $T^i_{a,j}$ and 0 elsewhere, and $DT_{i,K,t}$=(t-$T^i_{b,k}$) for t> $T^i_{b,k}$ and 0 elsewhere, where $T^i_{a,j}$ and $T^i_{b,k}$ denote the j-th and k-th dates of the breaks in level and the trend, respectively for the i-th individual, j=1,...,l_i, k=1,...,m_i. The first specification of the deterministic component consists of breaks affecting the level of the series. In the second specification, breaks in level and trend are allowed. As can be seen from the two previous specifications, BCIS is a very general test which allows a significant degree of individual heterogeneity. Indeed, in both models structural breaks can be positioned at different dates for different individuals, different magnitudes of shift, different numbers of structural breaks by individuals, different locations for the breaks for the level and the slope. As the number and location of the different structural breaks are unknown, they have to be estimated. Following, Bai and Carrion-I-Silvestre (2009)'s suggestion, we rely on the Bai and Perron's (1998) procedure to estimate the number and the location of the breaks. Three different test statistics have been proposed by Bai and Carrion-I-Silvestre (2009):

$$Z = \sqrt{N} \frac{MSB(\lambda - \bar{\xi})}{\bar{\sigma}} \rightarrow \mathcal{N}(0,1) \tag{C.22}$$

where MSB stands for the modified Sargan-Bhargava tests. $\overline{MSB(\lambda)} = N^{-1} \sum_{i=1}^{N} MSB_i(\lambda_i)$, $\bar{\xi} = N^{-1} \sum_{i=1}^{N} \xi_i$, and $\bar{\sigma}^2 = N^{-1} \sum_{i=1}^{N} (\sigma_i)^2$, where ξ_i and σ_i^2 respectively denote the mean and the variance of the individual $MSB_i(\lambda_i)$ statistics respectively.

$$P = -2 \sum_{i=1}^{N} ln(P_i) \rightarrow \chi^2(2N) \tag{C.23}$$

$$P_m = \frac{-2 \sum_{i=1}^{N} lnP_i - 2N}{\sqrt{4N}} \rightarrow \mathcal{N}(0,1) \tag{C.24}$$

where P_i are the P-value associated with the individual MSB tests. According to Bai and Carrion-I-Silvestre (2009), the P-stat is designed for fixed N, while the P_m-stat is suitable for large N panel.

Appendix C.C Cross-sectional augmented distributed lags model

The goal of the Cross Sectional-Distributed Lag (CS-DL) estimator (Chudik et al., 2016) is to estimate a long-run relationship as follows:

[39] Bai and Carrion-I-Silvestre do not constraint the number of break(s) in mean to be equal to the number of break(s) in trend.

$$y_{i,t} = \alpha_i + \gamma_i X_{i,t} + \epsilon_{i,t} \tag{C.25}$$

Where i denotes the country and t the year of observation. α_i stands for the country fixed effects. $X_{i,t}$ is the matrix of exogenous variables of interest, γ is the vector of the corresponding long-run parameters. $\epsilon_{i,t}$ is an error term which can be cross-sectionally correlated across countries. To explain how the CS-DL approach works, we write equation (C.25) as an ARDL model assuming one lag of the exogenous variables:[40]

$$y_{i,t} = \phi y_{i,t-1} + \beta_1 X_{i,t} + \beta_2 X_{i,t-1} + \epsilon_{i,t} \tag{C.26}$$

$$\epsilon_{i,t} = \gamma_i' f_t + u_{i,t} \tag{C.27}$$

$\epsilon_{i,t}$ is assumed to be a serially uncorrelated error term that, potentially, could be dependent across countries due to the presence of unobserved common factors as shown by (C.27). In this framework, the long-run coefficients are usually inferred from the short-run coefficients of the explanatory variables (i.e, ϕ, β_1 and β_2). The approach proposed by Chudik et al. (2016) consists to rewrite equation (C.26):

$$y_{i,t} = \gamma X_{i,t} + \alpha \Delta X_{i,t} + \tilde{\epsilon_{i,t}} \tag{C.28}$$

where $\tilde{\epsilon_{it}} = \dfrac{\epsilon}{1-\phi}$ and $\alpha = \beta_1 + \beta_2$. The unobserved common factors, f_t, can be captured by adding cross-sectional averages of the observables in our regressions (Pesaran, 2006; Chudik and Pesaran, 2015). Eq (C.25) can be estimated using the MG (Pesaran and Smith, 1995) and Pooled Mean Group (PMG; Pesaran et al., 1999) estimators. The MG estimator is obtained as follows:

$$MG = N^{-1} \sum_{i=1}^{N} \gamma_i \tag{C.29}$$

Where the γ_i is the individual coefficients obtained from a country OLS regression. To obtain the PMG estimator, the individual long-run coefficients are restricted to be the same across countries. The PMG estimator uses a maximum likelihood approach to estimate the model based on the Newton-Raphson algorithm. The MG estimator allows us to take into account full heterogeneity of the individuals in the panel, but this estimator produces consistent estimates of the average of the parameters only for sufficiently long time-series dimensions (Pesaran and Smith, 1995). Given the moderate time dimensions of our panel data, we thus prefer to rely on the PMG estimator.

[40] For the simplicity of presentation, we omit deterministic elements.

References

Aguirre, A. & Calderón, C. (2005), 'Real exchange rate misalignments and economic performance', *Documentos de Trabajo (Banco Central de Chile)* (315), 1–49.

Alberola, E. (2003), Misalignment, liabilities dollarization and exchange rate adjustment in Latin America, Working Papers 0309, Banco de España.

Bahmani-Oskooee, M. & Nasir, A. (2004), 'Ardl approach to test the productivity bias hypothesis', *Review of development Economics* **8**(3), 483–488.

Bahmani-Oskooee, M. & Rhee, H.-J. (1996), 'Time-series support for balassa's productivity-bias hypothesis: Evidence from korea', *Review of International Economics* **4**(3), 364–370.

Bai, J. & Carrion-I-Silvestre, J. L. (2009), 'Structural changes, common stochastic trends, and unit roots in panel data', *The Review of Economic Studies* **76**(2), 471–501.

Bai, J. & Ng, S. (2002), 'Determining the number of factors in approximate factor models', *Econometrica* **70**(1), 191–221.

Bai, J. & Ng, S. (2004), 'A panic attack on unit roots and cointegration', *Econometrica* **72**(4), 1127–1177.

Bai, J. & Perron, P. (1998), 'Estimating and testing linear models with multiple structural changes', *Econometrica* pp. 47–78.

Balassa, B. (1964), 'The purchasing-power parity doctrine: A reappraisal', *Journal of Political Economy* **72**(6), 584–596.

Bems, R. (2008), 'Aggregate investment expenditures on tradable and nontradable goods', *Review of Economic Dynamics* **11**(4), 852–883.

Bénassy-Quéré, A., Béreau, S. & Mignon, V. (2009), 'Robust estimations of equilibrium exchange rates within the g20: A panel beer approach', *Scottish Journal of Political Economy* **56**(5), 608–633.

Bénassy-Quéré, A. & Coulibaly, D. (2014), 'The impact of market regulations on intra-european real exchange rates', *Review of World Economics* **150**(3), 529–556.

Bergin, P. R., Glick, R. & Taylor, A. M. (2006), 'Productivity, tradability, and the long-run price puzzle', *Journal of Monetary Economics* **53**(8), 2041–2066.

Berka, M. & Steenkamp, D. (2018), Deviations in real exchange rate levels in the OECD countries and their structural determinants, Working Papers 2018-16, CEPII research center.

Berka, M., Steenkamp, D. et al. (2018), *Deviations in real exchange rate levels in the OECD countries and their structural determinants*, Reserve Bank of New Zealand, Te Pūtea Matua.

Betts, C. M. & Kehoe, T. J. (2001), 'Tradability of goods and real exchange rate fluctuations', *Federal Reserve Bank of Minneapolis Staff Report* .

Bordo, M. D., Choudhri, E. U., Fazio, G. & MacDonald, R. (2017), 'The real exchange rate in the long run: Balassa-samuelson effects reconsidered', *Journal of International Money and Finance* **75**, 69–92.

Burstein, A., Eichenbaum, M. & Rebelo, S. (2006), 'The importance of nontradable goods' prices in cyclical real exchange rate fluctuations', *Japan and the World Economy* **18**(3), 247–253.

Calderón, C., Moral-Benito, E. & Servén, L. (2015), 'Is infrastructure capital productive? a dynamic heterogeneous approach', *Journal of Applied Econometrics* **30**(2), 177–198.

Canzoneri, M. B., Cumby, R. E. & Diba, B. (1999), 'Relative labor productivity and the real exchange rate in the long run: evidence for a panel of oecd countries', *Journal of International Economics* **47**(2), 245 – 266.

Caputo, R. (2018), 'Real exchange rate appreciation after the financial crisis of 2008-2009: Misalignment or fundamental correction?', *International Finance* **21**(3), 253–272.

Cardi, O. & Restout, R. (2015), 'Imperfect mobility of labor across sectors: a reappraisal of the balassa–samuelson effect', *Journal of International Economics* **97**(2), 249–265.

Chang, Y. (2002), 'Nonlinear iv unit root tests in panels with cross-sectional dependency', *Journal of econometrics* **110**(2), 261–292.

Chinn, M. D. (1997), Sectoral productivity, government spending and real exchange rates: empirical evidence for oecd countries, Technical report, National Bureau of Economic Research.

Chinn, M. D. (2000), 'The usual suspects? productivity and demand shocks and asia–pacific real exchange rates', *Review of International Economics* **8**(1), 20–43.

Chinn, M. & Johnston, L. (1996), Real exchange rate levels, productivity and demand shocks: evidence from a panel of 14 countries, Technical report, National Bureau of Economic Research.

Choi, I. (2001), 'Unit root tests for panel data', *Journal of international money and Finance* **20**(2), 249–272.

Chong, Y., Jordà, Ò. & Taylor, A. M. (2012), 'The harrod–balassa–samuelson hypothesis: Real exchange rates and their long-run equilibrium', *International Economic Review* **53**(2), 609–634.

Choudhri, E. U. & Khan, M. S. (2005), 'Real exchange rates in developing countries: Are balassa-samuelson effects present?', *IMF Staff Papers* **52**(3), 387–409.

Choudhri, E. U. & Schembri, L. L. (2010), 'Productivity, the terms of trade, and the real exchange rate: Balassa–samuelson hypothesis revisited', *Review of International Economics* **18**(5), 924–936.

Christopoulos, D. K., Gente, K. & León-Ledesma, M. A. (2012), 'Net foreign assets, productivity and real exchange rates in constrained economies', *European Economic Review* **56**(3), 295–316.

Chudik, A., Mohaddes, K., Pesaran, M. H. & Raissi, M. (2016), Long-run effects in large heterogeneous panel data models with cross-sectionally correlated errors, *in* 'Essays in Honor of man Ullah', Emerald Group Publishing Limited, pp. 85–135.

Coto-Martinez, J. & Reboredo, J. C. (2014), 'The relative price of non-traded goods under imperfect competition', *Oxford Bulletin of Economics and Statistics* **76**(1), 24–40.

Couharde, C., Delatte, A.-L., Grekou, C., Mignon, V. & Morvillier, F. (2017), 'Eqchange: A world database on actual and equilibrium effective exchange rates'.

De Broeck, M. & Sløk, T. (2006), 'Interpreting real exchange rate movements in transition countries', *Journal of International Economics* **68**(2), 368–383.

De Gregorio, J., Giovannini, A. & Wolf, H. C. (1994), 'International evidence on tradables and nontradables inflation', *European Economic Review* **38**(6), 1225–1244.

De Gregorio, J. & Wolf, H. C. (1994), Terms of trade, productivity, and the real exchange rate, Technical report, National Bureau of Economic Research.

Devereux, M. B. (1999), 'Real exchange rate trends and growth: a model of east asia', *Review of International Economics* **7**(3), 509–521.

Doan, T. H. T. & Gente, K. (2013), 'Real exchange rate and productivity in an olg model', *Annals of Economics and Statistics/ANNALES D'ÉCONOMIE ET DE STATISTIQUE* pp. 259–281.

Doan, T. H. T. & Gente, K. (2014), 'Real exchange rate and productivity in a specific-factor model with skilled and unskilled labour', *Journal of Macroeconomics* **40**, 1–15.

Drine, I. & Rault, C. (2003*a*), 'Do panel data permit the rescue of the balassa-samuelson hypothesis for latin american countries?', *Applied Economics* **35**(3), 351–359.

Drine, I. & Rault, C. (2003*b*), 'A re-examination of the balassa–samuelson hypothesis using recent panel data unit-root and cointegration tests: Evidence from mena countries', *African Development Review* **15**(2-3), 106–125.

Drine, I. & Rault, C. (2004), 'Does the balassa-samuelson hypothesis hold for asian countries? an empirical analysis using panel data and cointegration tests', *An Empirical Analysis Using Panel Data and Cointegration Tests (August, 17 2008). Applied Econometrics and International Development* **4**(4).

Drine, I. & Rault*, C. (2005), 'Can the balassa-samuelson theory explain long-run real exchange rate movements in oecd countries?', *Applied Financial Economics* **15**(8), 519–530.

Driscoll, J. C. & Kraay, A. C. (1998), 'Consistent covariance matrix estimation with spatially dependent panel data', *Review of economics and statistics* **80**(4), 549–560.

Dumrongrittikul, T. (2012), 'Real exchange rate movements in developed and developing economies: A reinterpretation of the balassa-samuelson hypothesis', *Economic Record* **88**(283), 537–553.

Edwards, S. (1988), 'Real and monetary determinants of real exchange rate behavior: Theory and evidence from developing countries', *Journal of development economics* **29**(3), 311–341.

Edwards, S. (1989), *Real exchange rates, devaluation, and adjustment: exchange rate policy in developing countries*, MIT press Cambridge, MA.

Égert, B. (2002), 'Estimating the impact of the balassa–samuelson effect on inflation and the real exchange rate during the transition', *Economic Systems* **26**(1), 1–16.

Égert, B., Drine, I., Lommatzsch, K. & Rault, C. (2003), 'The balassa–samuelson effect in central and eastern europe: myth or reality?', *Journal of comparative Economics* **31**(3), 552–572.

Égert, B., Lommatzsch, K. & Lahrèche-Révil, A. (2006), 'Real exchange rates in small open oecd and transition economies: Comparing apples with oranges?', *Journal of Banking & Finance* **30**(12), 3393–3406.

Engel, C. (1999), 'Accounting for us real exchange rate changes', *Journal of Political Economy* **107**(3), 507–538.

EQCHANGE: A world database on actual and equilibrium effective exchange rates (2018), *International Economics* **156**, 206 – 230.

Faria, J. R. & Leon-Ledesma, M. (2003), 'Testing the balassa–samuelson effect: Implications for growth and the ppp', *Journal of Macroeconomics* **25**(2), 241–253.

Fernandez, C., Ley, E. & Steel, M. F. (2001a), 'Benchmark priors for bayesian model averaging', *Journal of Econometrics* **100**(2), 381–427.

Fernandez, C., Ley, E. & Steel, M. F. (2001b), 'Model uncertainty in cross-country growth regressions', *Journal of applied Econometrics* **16**(5), 563–576.

Fischer, C. (2004), 'Real currency appreciation in accession countries: Balassa-samuelson and investment demand', *Review of World Economics* **140**(2), 179–210.

Gala, P. (2007), 'Real exchange rate levels and economic development: theoretical analysis and econometric evidence', *Cambridge Journal of economics* **32**(2), 273–288.

Galstyan, V. & Lane, P. R. (2009), 'The composition of government spending and the real exchange rate', *Journal of Money, Credit and Banking* **41**(6), 1233–1249.

Galstyan, V. & Velic, A. (2018), 'Public debt and relative prices in a cross-section of countries', *Review of World Economics* **154**(2), 229–245.

García-Solanes, J., Sancho-Portero, F. I. & Torrejón-Flores, F. (2008), 'Beyond the balassa–samuelson effect in some new member states of the european union', *Economic Systems* **32**(1), 17–32.

García-Solanes, J. & Torrejón-Flores, F. (2009), 'The balassa-samuelson hypothesis in developed countries and emerging market economies: different outcomes explained', *Economics: The Open-Access, Open-Assessment E-Journal* **3**, 2.

Gente, K. (2006), 'The balassa–samuelson effect in a developing country', *Review of Development Economics* **10**(4), 683–699.

Grekou, C. (2019), 'From nominal devaluations to real depreciations', *International Economics* **157**, 68–81.

Gubler, M. & Sax, C. (2019), 'The balassa-samuelson effect reversed: new evidence from oecd countries', *Swiss Journal of Economics and Statistics* **155**(1), 3.

Halpern, L. & Wyplosz, C. (2001), Economic Transformation and Real Exchange Rates in the 2000s: The Balassa-Samuelson Connection, ECE Discussion Papers Series 2001-1, UNECE.

Hassan, A., Salim, R. & Bloch, H. (2011), 'Population age structure, saving, capital flows and the real exchange rate: A survey of the literature', *Journal of Economic Surveys* **25**(4), 708–736.

Hassan, F. (2016), 'The price of development: The Penn-Balassa-Samuelson effect revisited', *Journal of International Economics* **102**, 291–309.

Higgins, M. (1998), 'Demography, national savings, and international capital flows', *International Economic Review* pp. 343–369.

Hsieh, D. A. (1982), 'The determination of the real exchange rate: The productivity approach', *Journal of International Economics* **12**(3-4), 355–362.

Hurlin, C. & Mignon, V. (2005), 'Une synthèse des tests de racine unitaire sur données de panel', *Economie prevision* (3), 253–294.

Im, K. S., Pesaran, M. H. & Shin, Y. (2003), 'Testing for unit roots in heterogeneous panels', *Journal of econometrics* **115**(1), 53–74.

Imai, H. (2018), 'China's rapid growth and real exchange rate appreciation: Measuring the balassa-samuelson effect', *Journal of Asian Economics* **54**, 39–52.

Inklaar, R. & Timmer, M. P. (2008), GGDC Productivity Level Database: International Comparisons of Output, Inputs and Productivity at the Industry Level, Technical report.

Ito, T., Isard, P. & Symansky, S. (1999), Economic growth and real exchange rate: an overview of the balassa-samuelson hypothesis in asia, *in* 'Changes in exchange rates in rapidly developing countries: Theory, practice, and policy issues (NBER-EASE volume 7)', University of Chicago Press, pp. 109–132.

Iyke, B. N. & Odhiambo, N. M. (2017), 'An empirical test of the balassa-samuelson hypothesis: Evidence from eight middle-income countries in africa', *Economic Systems* **41**(2), 297–304.

Kakkar, V. & Yan, I. (2012), 'Real exchange rates and productivity: evidence from asia', *Journal of Money, Credit and Banking* **44**(2-3), 301–322.

Lane, P. R. & Milesi-Ferretti, G. M. (2002), 'External wealth, the trade balance, and the real exchange rate', *European Economic Review* **46**(6), 1049–1071.

Lane, P. R. & Milesi-Ferretti, G. M. (2007), 'The external wealth of nations mark ii: Revised and extended estimates of foreign assets and liabilities, 1970–2004', *Journal of international Economics* **73**(2), 223–250.

Lee, J. & Tang, M.-K. (2007), 'Does productivity growth appreciate the real exchange rate?', *Review of International Economics* **15**(1), 164–187.

Ley, E. & Steel, M. F. (2009), 'On the effect of prior assumptions in bayesian model averaging with applications to growth regression', *Journal of applied econometrics* **24**(4), 651–674.

Lombardo, G. & Ravenna, F. (2012), 'The size of the tradable and non-tradable sectors: Evidence from input–output tables for 25 countries', *Economics letters* **116**(3), 558–561.

Lothian, J. R. & Taylor, M. P. (2008), 'Real exchange rates over the past two centuries: how important is the harrod-balassa-samuelson effect?', *The Economic Journal* **118**(532), 1742–1763.

Loungani, M. P., Mishra, M. S., Papageorgiou, M. C. & Wang, K. (2017), *World trade in services: Evidence from a new dataset*, International Monetary Fund.

MacDonald, R. & Ricci, L. A. (2005*a*), 'The real exchange rate and the balassa–samuelson effect: The role of the distribution sector', *Pacific Economic Review* **10**(1), 29–48.

MacDonald, R. & Ricci, L. A. (2005*b*), 'The real exchange rate and the balassa–samuelson effect: the role of the distribution sector', *Pacific Economic Review* **10**(1), 29–48.

Maddala, G. S. & Wu, S. (1999), 'A comparative study of unit root tests with panel data and a new simple test', *Oxford Bulletin of Economics and statistics* **61**(S1), 631–652.

Mano, R. & Castillo, M. (2015), *The level of productivity in traded and non-traded sectors for a large panel of countries*, number 15-48, International Monetary Fund.

Marston, R. C. (1986), Real exchange rates and productivity growth in the united states and japan, Working Paper 1922, National Bureau of Economic Research.

Méjean, I. (2008), 'Can firms' location decisions counteract the balassa–samuelson effect?', *Journal of International Economics* **76**(2), 139–154.

Mihaljek, D. & Klau, M. (2004), 'The balassa–samuelson effect in central europe: a disaggregated analysis', *Comparative Economic Studies* **46**(1), 63–94.

Mihaljek, D. & Klau, M. (2008), 'Catching-up and inflation in transition economies: the balassa-samuelson effect revisited'.

Moral-Benito, E. (2015), 'Model averaging in economics: An overview', *Journal of Economic Surveys* **29**(1), 46–75.

Moral-Benito, E. & Roehn, O. (2016), 'The impact of financial regulation on current account balances', *European Economic Review* **81**, 148–166.

Moran, P. A. (1948), 'The interpretation of statistical maps', *Journal of the Royal Statistical Society. Series B (Methodological)* **10**(2), 243–251.

Njindan Iyke, B. (2017), 'The penn effect revisited: new evidence from latin america', *Review of Development Economics* **21**(4), 1364–1379.

Obstfeld, M. & Rogoff, K. (1996), 'Foundations of international macroeconomics'.

Ouyang, A. Y. & Rajan, R. S. (2013), 'Real exchange rate fluctuations and the relative importance of nontradables', *Journal of International Money and Finance* **32**, 844–855.

Pesaran, M. H. (2004), 'General diagnostic tests for cross section dependence in panels'.

Pesaran, M. H. (2006), 'Estimation and inference in large heterogeneous panels with a multifactor error structure', *Econometrica* **74**(4), 967–1012.

Pesaran, M. H. (2007), 'A simple panel unit root test in the presence of cross-section dependence', *Journal of applied econometrics* **22**(2), 265–312.

Pesaran, M. H. (2015), *Time series and panel data econometrics*, Oxford University Press.

Pesaran, M. H., Shin, Y. & Smith, R. P. (1999), 'Pooled mean group estimation of dynamic heterogeneous panels', *Journal of the American Statistical Association* **94**(446), 621–634.

Pesaran, M. H., Smith, L. V. & Yamagata, T. (2013), 'Panel unit root tests in the presence of a multifactor error structure', *Journal of Econometrics* **175**(2), 94–115.

Pesaran, M. H. & Smith, R. (1995), 'Estimating long-run relationships from dynamic heterogeneous panels', *Journal of econometrics* **68**(1), 79–113.

Piton, S. (2019), 'Staff working paper no. 799 do unit labour costs matter? a decomposition exercise on european data'.

Raftery, A. E. (1995), ‘Bayesian model selection in social research’, *Sociological methodology* **25**, 111–164.

Reese, S. & Westerlund, J. (2016), ‘Panicca: Panic on cross-section averages’, *Journal of Applied Econometrics* **31**(6), 961–981.

Ricci, L. A., Milesi-Ferretti, G. M. & Lee, J. (2013), ‘Real exchange rates and fundamentals: a cross-country perspective’, *Journal of Money, Credit and Banking* **45**(5), 845–865.

Rogoff, K. (1996), ‘The purchasing power parity puzzle’, *Journal of Economic literature* **34**(2), 647–668.

Rose, A. K., Supaat, S. & Braude, J. (2009), ‘Fertility and the real exchange rate’, *Canadian Journal of Economics/Revue canadienne d’économique* **42**(2), 496–518.

Samuelson, P. A. (1964), ‘Theoretical notes on trade problems’, *The Review of Economics and Statistics* **46**(2), 145–154.

Schnatz, B., Vijsellaar, F. & Osbat, C. (2004), ‘Productivity and the euro-dollar exchange rate’, *Review of World Economics* **140**(1), 1–30.

Sheng, Y. & Xu, X. (2011), ‘Real exchange rate, productivity and labor market frictions’, *Journal of International Money and Finance* **30**(3), 587–603.

Solow, R. M. (1957), ‘Technical change and the aggregate production function’, *The review of Economics and Statistics* pp. 312–320.

Thomas, A. & King, A. (2008), ‘The balassa–samuelson hypothesis in the asia-pacific region revisited’, *Review of International Economics* **16**(1), 127–141.

Tsen, W. H. (2011), ‘The real exchange rate determination: An empirical investigation’, *International Review of Economics & Finance* **20**(4), 800–811.

Wang, W., Xue, J. & Du, C. (2016), ‘The balassa–samuelson hypothesis in the developed and developing countries revisited’, *Economics Letters* **146**, 33–38.

Westerlund, J. & Urbain, J.-P. (2015), ‘Cross-sectional averages versus principal components’, *Journal of Econometrics* **185**(2), 372–377.

Williamson, J. (1994), *Estimating equilibrium exchange rates*, Peterson Institute.

Yehoue, M. E. B. & Dufrénot, G. J. (2005), *Real exchange rate misalignment: A panel co-integration and common factor analysis*, number 5-164, International Monetary Fund.