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**Working Paper**

## Balassa-Samuelson, product differentiation and transition

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Arbeiten aus dem



OSTEUROPA-INSTITUT  
MÜNCHEN

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**Balassa-Samuelson, Product Differentiation  
and Transition**

Richard FRENCH



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# Contents

Tabellenverzeichnis.....	iv
Abbildungsverzeichnis.....	iv
<b>Comparative prices and transition: the EU-accession countries in international perspective .....</b>	<b>v</b>
Abstract .....	v
1 Introduction.....	1
2 A review of Balassa-Samuelson .....	2
2.1 Purchasing power parity and deviations .....	2
2.2 The empirical BS-literature .....	4
3 Comparative prices and per capita incomes: evidence from PWT data.....	6
3.1 Comparative prices and real effective exchange rates.....	6
3.2 The price-income benchmark relationship.....	7
4 Transition countries in the cross-section price-income relationship.....	11
5 Comparative price changes and per capita income growth .....	14
5.1 The simple BS view .....	14
5.2 An extended view.....	15
5.2.1 Equilibrium forces versus adjustment to equilibrium .....	15
5.2.2 "Real" sources of real exchange rate development during transition and beyond.....	17
5.2.3 Data and measurement.....	19
5.2.4 Specification and estimation results .....	20
6 Conclusions.....	24
References.....	25
Appendix A: Tables and Figures .....	27
Appendix B. PWT data construction and economic relationships.....	34
<b>Balassa-Samuelson effects in the presence of product differentiation and trade barriers. Implications for transition and convergence.....</b>	<b>35</b>
Abstract .....	35
1.Introduction.....	36
2 Prices and productivities: The simple arbitrage view .....	37
3 A monopolistic competition model of trade with product differentiation .....	39
3.1 Production.....	39
3.2 Demand .....	39
3.3 Barriers to trade .....	40
3.4 International equilibrium.....	41
3.5 Prices and productivities with trade in differentiated goods .....	42
4 Prices and productivities: An extended arbitrage view for transition economies .....	44
5 Some conclusions .....	48
References.....	49

## List of Tables

Table 1: OLS results for equation (6), 2000 .....	10
Table 2: Panel least squares results for equation (7), 1988–2000 .....	14
Table 3a: Panel EGLS estimation results for equation (22), 1992–2000, in levels .....	22
Table 3b: Panel EGLS estimation results for equation (22), 1992–2000, in yearly changes .....	23
Table A-1: Country list .....	30
Table A-2: Extrema of the distributions of residuals from regression (6), 1991–2000 .....	31
Table A-2 contd.: Extrema of the distributions of residuals from regression (6), 1991–2000 .....	32

## List of Figures

Figure 1: Comparative prices, $p_j$ , versus PPP-adjusted income per capita, $y_j$ , in logs for a large panel of countries, 1950–2000 .....	7
Figure 2: Comparative prices versus PPP-adjusted income per capita at various points in time .....	8
Figure 3: Standardised residuals from equation (6) versus reform indices of transition economies .....	12
Figure 4: Standardised residuals from (6) for selected transition countries, 1991–2000 .....	13
Figure A-1: Growth differentials between trade weighted CPI-deflated real effective exchange rates, $reer_j$ , and comparable price levels, $p_j$ , versus PPP-adjusted income per capita, $y_j$ , 1990–2000 .....	27
Figure A-2a: Comparative price level versus PPP-adjusted income per capita, in logs, 2000 .....	28
Figure A-2b: Residuals from $\ln p_j = c(1) + c(2) \cdot \ln y_j$ , for the year 2000 .....	28
Figure A-3a: Comparative price level versus PPP-adjusted income per capita, in logs, 2000. Partial relationship, controlling for OECD $\cdot \ln y_j$ .....	29
Figure A-3b: Residuals from $Resid_j = c(1) + c(2) \cdot \ln y_j$ , for the year 2000 .....	29
Figure A-4: Standardised residuals from estimating equation 7. Table 2, column (2) .....	33

## **Comparative prices and transition: the EU-accession countries in international perspective**

### **Abstract**

Recent panel studies have found relatively high estimates for the elasticity of real exchange rates with respect to productivity measures in transition economies within Balassa-Samuelson frameworks. This contrasts with other findings reporting cross-section price-income elasticity estimates to depend positively on average income in the sample. This paper aims to reconcile both results by putting real exchange rate developments of transition economies in an international perspective. We illustrate the special status of these economies in a simple world-wide Balassa-Samuelson-type price-income benchmark relationship between a real exchange rate measure (Penn World Table comparative prices, i.e., exchange rate gaps) and PPP-adjusted per capita income. A pronounced undervaluation at the start of transition, followed by a strong appreciation results in normalisation towards the benchmark for Central and East European economies (CEEC) but not for the CIS. We then make an attempt at extending the simple price-income relationship to incorporate other real factors as well as reforms related to price deregulation. Our results imply that, when accounting for demand shifts, external liberalisation, and especially for reform effort, the price-income-elasticity for CEEC economies was not different from that of non-transition economies during the nineties.

*JEL*-Classification: F40, F43

Keywords: Balassa-Samuelson, transition

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## **1 Introduction**

Recent panel data studies (e.g., Dobrinsky, 2003; de Broeck and Sløk, 2006) have found comparatively high estimates for the elasticity of real exchange rates with respect to productivity measures for transition economies, especially in Central and Eastern Europe (CEEC), within Balassa-Samuelson frameworks. While this feature is usually explained by extraordinary reform efforts in these countries spurring productivity growth, it seems to contrast with other findings reporting that price-income elasticity estimates from cross-section regression analyses vary greatly depending on sample composition, where the inclusion of poor countries tends to generate lower elasticities (Maeso-Fernandez et al., 2005).

The idea of the paper is to investigate both features by putting the real exchange rate behaviour of transition economies into an international perspective. For this purpose, we first motivate and introduce a simple international benchmark relationship between comparative prices and per capita income, based on the Balassa-Samuelson effect (sections 2 and 3). Section 4 illustrates the special status of transition economies in the cross-section version of this relationship: a pronounced undervaluation at the start of transition is followed by a transition-specific pattern of strong appreciation during the nineties, and results in some sort of “normalisation” for Central and East European economies (CEEC) but not for the CIS, i.e., at the beginning of the decade, CEEC economies were not part of the international price-income benchmark relationship, while by the end of the decade they were. To study what moved them there we make an attempt in section 5 to extend the simple price-income relationship to incorporate demand shifts, liberalisation, and especially reform efforts related to price deregulation. Results of estimating the extended approach with panel data both in levels and in yearly changes imply that, when accounting for demand shifts, external liberalisation, and especially for price deregulation effort, the price-income-elasticity for CEEC economies is not different from that of non-transition economies.



## 2 A review of Balassa-Samuelson

### 2.1 Purchasing power parity and deviations

Purchasing power parity is linked to the tradability of goods and services. Arbitrage ensures that the price of an internationally traded good be the same everywhere in the world when expressed in a common currency at the going exchange rate. If all goods are tradable ( $T$ ) – and enter each country’s basket used to construct the aggregate price level with the same weight – aggregate price levels,  $P_j^T$ , are identical for each pair of countries if expressed in a common currency at the going exchange rate,

$$P_1^T e_{12} = P_2^T, \quad (1)$$

where country indices are 1 and 2, and  $e_{12}$  is the going nominal exchange rate, which will express purchasing power parity (PPP): one unit of home currency buys the same basket of goods at home as the equivalent amount of foreign currency. Defining the deviation of the ratio of two countries’ aggregate price levels from their nominal exchange rate as the real exchange rate between countries 2 and 1,  $RER_{21}$ , absolute PPP is equivalent to

$$RER_{21}^T = \frac{P_2^T / P_1^T}{e_{12}} = 1. \quad (2)$$

Transactions costs may imply a violation of PPP. In addition, different countries tend to produce goods that are similar rather than perfectly substitutable, and the weights attached to similar goods in aggregate price indices may differ across countries. But most importantly, not all goods are traded. PPP may still hold for tradables but not for non-tradables, e.g. for many services. Consequently, even with identical baskets used to construct the aggregate price level with the same weights across countries, in the presence of non-tradables the nominal current market exchange rate cannot be expected to express PPP.

In fact, as noticed about forty years ago by Balassa and Samuelson,<sup>1</sup> what one observes empirically is not just a deviation, but a systematic deviation of current exchange rates from PPP levels: at the going exchange rate – even if expressing PPP for tradables – non-tradables prices and thus aggregate price levels are higher in richer than in poorer economies. Both Balassa (1964) and Samuelson (1964 and again 1994) rationalised this pattern in a chain of arguments building on (a): purchasing power holds for tradables, (b): relative prices reflect relative labour productivities, (c): national labour markets are homogenous across sectors of production, and (d): the biggest differences in (labour) productivity across countries are in tradable rather than non-tradable production. Leav-

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<sup>1</sup> As noted in Bergin et al. (2004), this observation can be traced back to Ricardo and from there, via Taussig, Ohlin, Viner, Harrod, and Rothschild, to Balassa (1964) and Samuelson (1964).

ing (d) aside leaves the productivity gap version of the Balassa-Samuelson-hypothesis (BS): for each pair of countries, country 2's real exchange rate is higher than that of country 1 if country 2's productivity in the tradables sector, relative to the non-tradables sector, is higher than in country 1.

The simplest productivity gap version of the BS-hypothesis is stated within a framework of two countries, two homogenous goods (one tradable, one non-tradable), and one factor of production (labour). Steps (a) – (c) then ensure,<sup>2</sup>

$$RER_{21} = \frac{P_2 / P_1}{e_{12}} = \frac{(A_2^T A_1^N)^{1-\theta}}{(A_1^T A_2^N)^{1-\theta}}, \quad (3)$$

where  $P_j$  is the national price level when not all goods are tradable.  $A_j^T$  and  $A_j^N$  are (labour) productivity in country  $j$ 's tradable and non-tradable sectors, and equal preferences across countries are described by constant and equal consumption expenditure shares for tradables and non-tradables,  $\theta$  and  $1-\theta$ , respectively. Adding step (d), i.e., that cross-country productivity differences occur only in the tradable goods sector immediately implies the original statement of the BS-hypothesis,

$$RER_{21} = \frac{P_2 / P_1}{e_{12}} = \left( \frac{y_2}{y_1} \right)^{1-\theta}, \quad (4)$$

with  $y_j$  as per capita income. Specifically, the consumption expenditure share for non-tradables,  $1-\theta$ , corresponds to the elasticity of the real exchange rate with respect to relative per capita income, which we will in this paper refer to as the *price-income elasticity* within the BS-framework.

There is considerable opposition to both the theoretical validity and the empirical content of steps (a) – (d): as a most obvious point of criticism, it is challenging to determine whether or not a particular product is tradable (Parsley and Wei, 2004; Stein, 2005). Equally fundamental, even slight deviations from the simplifying assumptions of the two-countries-two-goods-one-factor set-up illustrate the knife-edge role of these assumptions for the resulting systematic relationship between prices and productivities (Podkaminer, 2003).

More specifically on the evidence for and against (a) – (d), the idea that purchasing power parity holds for tradables must certainly be restricted to take account of the existence of quality differences. In addition, while the notion that relative prices reflect relative productivities is in general little contested, Strauss (1997) does not find support for the assumption of wage equalisation across traded and nontraded sectors, at least not for the short and medium term, defined as up to four years, in fourteen OECD countries under consideration. I.e., at least in the short and medium term, wage differentials may have an impact on the real exchange rate.

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<sup>2</sup> For a simple derivation see, e.g., Frensch (2005).

For the U.S., Slifman and Corrado (1996) concluded that “the published figures for business and manufacturing labor productivity suggest that since the beginning of the 1990s output per hour in the nonmanufacturing sector of the economy has been disappointing” (p. 1). Gullickson (1992) enhance this finding by observing that “all of the growth in private business multifactor productivity in the US during the 1980s could be attributed to manufacturing.” But most importantly, based on the 1996 benchmark study of the Penn World Tables, Herrendorf and Valentinyi (2005) find “that, indeed, the cross-country TFP differences are by far larger in the tradable than in the nontradable sectors. Moreover, since this translates into differences in labor productivities, our evidence suggests that Balassa and Samuelson were right!” (pp. 18f).

Summing up, while there is considerable opposition to both the original BS-hypothesis and its productivity gap version, the criticism holds equally for both versions of the proposition. Thus, step (d) constitutes perhaps the least contested element of the BS-proposition: the biggest differences in (labour) productivity across countries are indeed in tradable rather than non-tradable production. One might therefore argue that testing either the original BS-hypothesis or its productivity gap version is by and large equivalent: if the one holds (or not), so does the other. In this paper, we will stick to the original BS-hypothesis, assuming that results hold for the productivity gap version as well.

## **2.2 The empirical BS-literature**

The literature tests for two basic questions: is there a – persistent rather than transitory – systematic deviation of nominal exchange rates from PPP, and if yes, is this deviation related to sectoral productivity gaps or per capita income differences? On the first question, Froot and Rogoff (1995) argue that producer price indices lend more support to PPP since they contain a higher proportion of tradable goods than consumer price indices. This has recently been confirmed in Coakley et al. (2005).

According to the recent overview by Bahmani-Oskooee and Nasir (2005), the empirical support for the productivity bias hypothesis is mixed and depends mainly on the technique and the nature of the data employed. Time series and panel data approaches have lately provided more support to the hypothesis than earlier cross-sectional studies. In addition, most recent studies “have included other factors in their model in addition to productivity differentials” (p. 692).

The BS literature on transition countries has, with the notable exception of de Broeck and Sløk (2006), so far been mostly confined to identifying BS-effects within this country group’s data, without putting them into an international perspective. Early results in this vein had been used as a basis for arguing that real appreciation in the region is to a large extent due to BS. Dobrinsky (2003) confirms results by Halpern and Wyplosz (2001) and “suggests roughly a one-to-one proportion between the differential in productivity growth and the consumer price-based real exchange rate appreciation in the second half of the 1990s” (p. 329).

However, Égert and Halpern (2005) in their meta-regression analysis of studies on CEEC real exchange rate behaviour fail to find a significant influence of a simple BS-driven behaviour on real exchange rate developments in the region, i.e., there seems more at work than BS. Recent work has supported this on the ground that PPP does not necessarily hold even for tradables, e.g. due a quality adjustment bias, referred to in Cincibuch and Podbiera (2004). This seems to imply that BS does perhaps not add that much to inflation differences vis-à-vis the euro area for many countries in the region such that specifically BS will not eventually collide with the Maastricht conditions concerning the new EU member states' readiness to join EMU.

Summing up the issue, Égert et al. (2004) stress three stylised facts of real exchange rate behaviour in transition:

1. until around the mid-nineties, real exchange rates in transition countries were substantially undervalued in terms of PPP;
2. although different in extent across countries, the region has witnessed strong appreciation from the outset of transition.
3. different from the BS pattern of explanation, all types of goods, not only non-tradable services, were or still are undervalued in terms of PPP.

### 3 Comparative prices and per capita incomes: evidence from PWT data

#### 3.1 Comparative prices and real effective exchange rates

The Penn World Table (PWT) price and income data used in this paper are derived from the International Comparison Project (ICP),<sup>3</sup> which is about establishing purchasing power parities over goods and services, combinations of goods and services, and finally over GDPs. This in turn enables one to find deviations between purchasing power parities and nominal exchange rates, i.e. *comparative price levels* in PWT terminology or *exchange rate gaps* in much of the literature. With only two countries, the comparative price level is identical to the definition of a real exchange rate in section 2. However, any attempt to measure deviations between purchasing power parities and nominal exchange rates between a country and a group of other countries or even the rest of the world necessarily involves a weighting scheme, and there are many weighting schemes conceivable. Each country's PWT comparative price level,  $p_j$ , is by construction a weighted real exchange rate against the international dollar, where the weighting scheme is based on the relative prices that underlie the derivation of the international dollar (see Appendix B).

The most popular measure for a country's weighted real exchange rate in a multilateral world is the trade-weighted real effective exchange rate index,  $reer_j$ , which is a weighted sum of each country's bilateral nominal exchange rates deflated by consumer price indices with weights corresponding to the relative importance of partner countries in trade. The IMF's *International Financial Statistics* (IFS) provide series for a number of countries on this measure, usually starting with 1975, in form of country-specific indices which – in contrast to comparative price levels – cannot be compared in levels across countries in an economically meaningful way.

In order to compare both measures over the period we are most interested in, we perform a simple OLS regressing yearly changes of all available IFS  $reer_j$  data for the decade between 1990 and 2000 on yearly changes of PWT  $p_j$ . The estimated slope coefficient of 0.40 is significant at the 1 per cent level, the intercept is insignificant at the 10 per cent level ( $R^2 = 0.29$ ; sample size = 864 observations between 1990 and 2000). Specifying country and/or period fixed effects does not qualify the results. Increasing the time horizon and thus eliminating nominal disturbances even strengthens the link between the two measures: the slope coefficient from an OLS regressing non-overlapping five-(ten-)year changes of all available IFS  $reer_j$  data for the decade between 1990 and 2000 on five-(ten-)year changes of  $p_j$  is 0.71 (0.93).

This supports the view that the differently constructed  $reer_j$  and  $p_j$  are well correlated measures of the deviation of a country's multilateral exchange rate from PPP. Also, as Figure A-1 in the appendix illustrates, differentials between rates of change of the two

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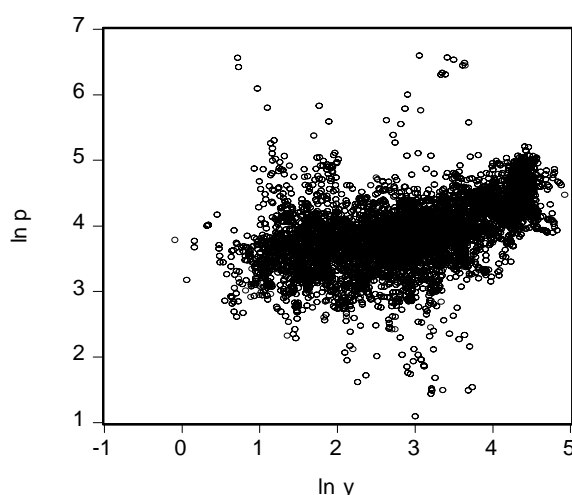
<sup>3</sup> Descriptions of the ICP and of the PWT dataset derived from the ICP can be found in Summers and Heston (1991), Heston and Summers (1996) and the PWT site at [http://pwt.econ.upenn.edu/php\\_site/pwt\\_index.php](http://pwt.econ.upenn.edu/php_site/pwt_index.php)

measures are not systematically related to PPP-adjusted income per capita,  $y_j$ . This is important, as in what follows we are mainly interested in the relationship between  $p_j$  and  $y_j$ . However, the comparative price level has the enormous advantages of being more widely available and being internationally comparable in level terms, which is why we use this real exchange rate measure in the rest of the paper.

### 3.2 The price-income benchmark relationship

Figure 1 gives a very rough account of the relationship between comparative price levels and PPP-adjusted per capita incomes in indiscriminately using all available PWT data between

Figure 1: **Comparative prices,  $p_j$ , versus PPP-adjusted income per capita,  $y_j$ , in logs for a large panel of countries, 1950–2000**



Notes: Both  $y_j$  and  $p_j$  are relative to the U.S., i.e.,  $y_{US} = 100$  and  $p_{US} = 100$ . Data are from Heston, Summers and Aten, Penn World Table Version 6.1, Center for International Comparisons at the University of Pennsylvania, October 2002. Sample size: 5,847 observations between 1950 and 2000.

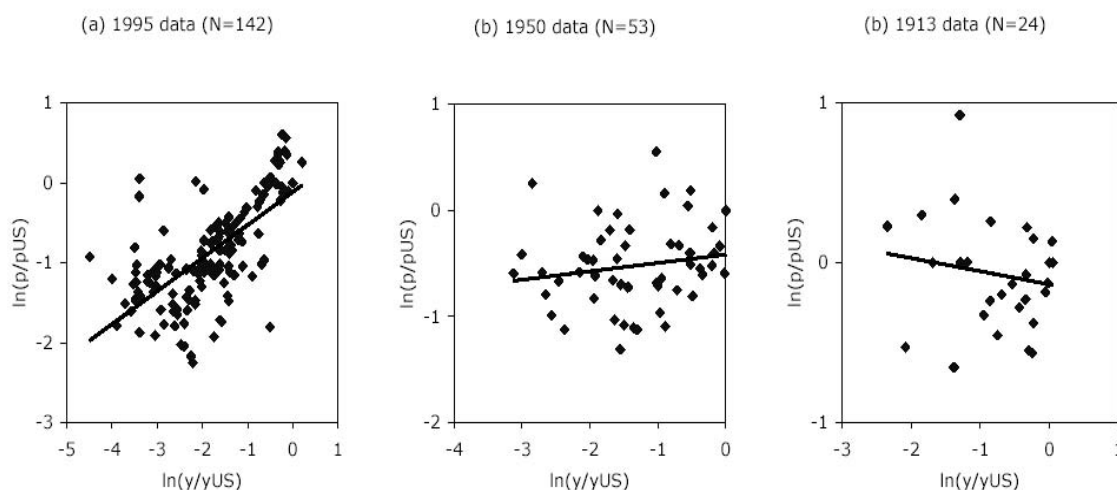
1950 and 2000. While this simple scatter plot does not seem to contradict the existence of a log-linear relationship between comparative price levels and per capita incomes to hold universally, a more detailed inspection of the panel data along both their time and cross-section dimensions suggests questioning this universality.

In terms of the *time dimension* of the data, Bergin et al. (2004), examining post-war data in detail and finding evidence even going back for centuries, conclude that “the price-income correlation was not really very strong until the last three or four decades” (p. 1; see also Figure 2 below). They perform simple cross-section OLS on PWT data, specified as,

$$\ln p_j = \alpha_0 + \alpha_1 \ln y_j + \varepsilon_j, \quad (5)$$

where both  $p_j$  and  $y_j$  are relative to U.S. levels (see Appendix B for details of constructing PWT data). As (5) is in logs, the estimated slope coefficient,  $\alpha_1$ , can be interpreted as the elasticity of the comparative price level with respect to per capita income, in short, the price-income elasticity. Bergin et al. (2004) conclude that “in a sequence of PWT cross sections every 5 years from 1950 to 1995, the BS effect has gradually strengthened, with the slope estimate roughly quadrupling in size over half a century” (p. 4). Interestingly, the null hypothesis of a zero slope can be rejected only since the early 1960s, which is when Balassa and Samuelson wrote their seminal papers.

Figure 2: **Comparative prices versus PPP-adjusted income per capita at various points in time**



Notes: N refers to the number of countries in respective samples. Note the difference in scale to Figure 1: here,  $y_{US} = 1$  and  $p_{US} = 1$ . Most countries have incomes and price levels lower than the United States, so the ratios are less than one and the logs are negative.

Source: Bergin et al. (2004) as presented in Taylor and Taylor (2004).

Why the BS-effect has altered over time remains a question of active research. According to one straightforward explanation (see equation 4) the non-tradable share in income has increased over time: after all, if (5) were to directly test the hypothesis (4),  $\alpha_1$  should correspond to an estimator of the consumption expenditure share for non-tradables,  $1-\theta$ . However, neither does such an effect seem to have enough magnitude to match the changes that have occurred, nor does it meet the timing of the changes: in fact, in 1950 traded shares of output were lower both than in 1913 or in 2000 (Taylor

and Taylor, 2004).<sup>4</sup> In consequence, either (5) is simply misspecified – or the BS-effect does not hold always and everywhere.

While the correlation described in (5) seems indeed present in today’s data, a closer look at the *cross-section dimension* of the  $p$ - $y$  relationship is also revealing. Maeso-Fernandez et al. (2005) report that price-income elasticity estimates from cross-section regression analyses vary greatly depending on sample composition. “(T)he inclusion of poor countries – particularly, African countries – tends to generate lower elasticities” (p. 139). Summers and Heston (1991, p. 336) express on the  $p$ - $y$  relationship for 1980 data that “the distinct heteroskedasticity apparent in the graph is not properly taken into account by the log-linear functional form that was used.” However, Figure A-2 in the appendix suggests that the residuals from a log-linear cross-country regression such as (5) exhibit serial correlation rather than heteroskedasticity when observations are in increasing order of  $y_j$ . This becomes especially visible when inspected in  $y$ - $p$  (rather than in  $\ln y - \ln p$ ) space, as is done in the right-hand panel of Figure A-2a. This evidence suggests different strengths of the  $p$ - $y$  relationship in sub-samples of countries at each point in time. However, when attempting to divide the sample one has to keep in mind that homogenous sub-samples – in term of income intervals – need not be identical for each year of observation. With this in mind, we formulate a simple variation of (5) that corrects for the correlation of residuals from (5) with  $y_j$ ,

$$\ln p_j = \beta_0 + \beta_1 \cdot \ln y_j + \beta_2 \cdot OECD \cdot \ln y_j + \varepsilon_j, \quad (6)$$

where again  $p_j$  and  $y_j$  are relative to U.S. levels and  $OECD = 1$  for OECD member countries.<sup>5</sup>

We now perform simple OLS cross-section regressions based on (6) with PWT data for each year between 1991 and 2000, where Figures A-3a and b illustrate that the specification in (6) is not plagued by heteroskedasticity or serial correlation when observations are in ascending order of  $y_j$ . Table 1 shows the result of the year 2000 regression. Especially, the elasticity of the price level with respect to PPP-adjusted per capita income now appears significantly higher for OECD countries than for others, a result in line with our earlier observation that the residuals from the basic log-linear cross-country regression (5) are correlated with income per capita.

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<sup>4</sup> Recent approaches to endogenise BS-effects (see especially Bergin et al., 2004) start out with the hypothesis that trade costs determine tradability.

<sup>5</sup> A “1” is assigned in case of full-year membership except for CEEC countries that joined the organisation during the 1990s, i.e. the Czech Republic in 1995, Hungary and Poland in 1996, and Slovakia in 2000. These four countries are consistently treated as non-OECD. For a full country list, see Table A-1 in the appendix.



Table 1: **OLS results for equation (6), 2000**

Dependent variable: $\ln p_j$			
Explanatory variable	Coefficient	<i>t</i> -statistic	<i>P</i> -value
<i>Constant</i>	2.77	26.34	0.00
$\ln y_j$	0.28	6.80	0.00
<i>OECD</i> • $\ln y_j$	0.13	4.70	0.00
Adjusted $R^2$ : 0.56			
Sample: 135 PWT countries			

## 4 Transition countries in the cross-section price-income relationship

Our next step consists of a “fishing expedition” à la Suhrcke (2001), i.e., our approach of putting transition countries’ comparative prices in an international perspective differs from de Broeck and Sløk (2006): while they estimate a  $p$ - $y$  relationship without transition economies and confront this benchmark with the developments of comparative prices in transition economies, we explicitly include transition economies in the sample to estimate ten cross-section regressions of equation (6) for each year between 1991 and 2000. We then inspect the distribution of the residuals from each of the ten regressions and identify outlier countries from the upper and lower 5 (10) per cent of these distributions.

At the very lower end of these distributions one should find countries with comparative price levels significantly lower than suggested by their PPP-adjusted per capita income. As Table A-2 shows, the lowest percentile of this relationship is between 1991 and 2000 almost exclusively made up of transition economies, which in our sample (see Table A-1) consist of the CEEC, the CIS, China, and Vietnam. Of the altogether 71 observations in the lowest 5 per cent of the distributions of the residuals from the ten regressions of equation (6) for the period 1991–2000, only 4 are non-transition economies (Nepal in 1991, Indonesia, Mauritius and Zimbabwe in 1998). Out of the 17 transition economies, which we have data for in 1992 (15 CEEC and the former CIS plus China and Vietnam), 11 are present in the group of 14 countries constituting the lowest 10 per cent of the residual distribution from the 1992 regression of equation (6). Almost needless to say, Table A-2 confirms that no transition economy features among the upper outliers of these distributions any time between 1991 and 2000.

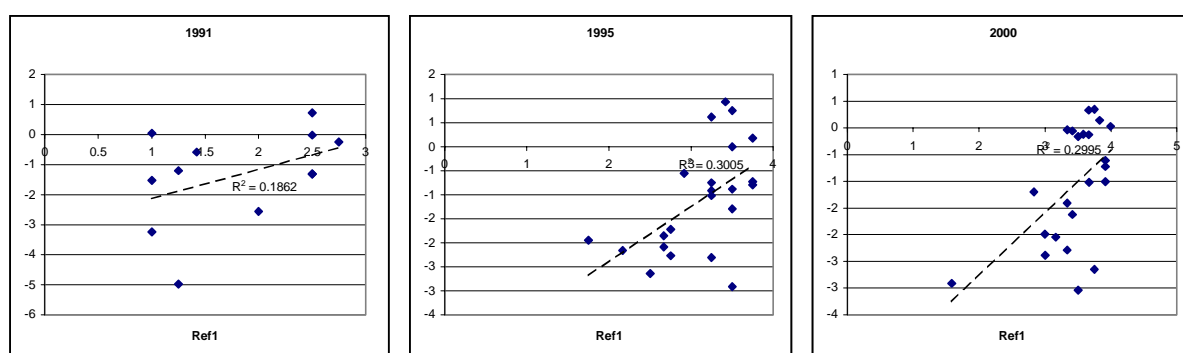
To conclude, at least until the mid-1990s it is possible to isolate transition economies without any prior knowledge as the ones “below” the international benchmark relationship between comparative prices and per capita income. In 2000, this is no longer true for the CEEC but still for the CIS economies (see Table A-2): all the ten CIS economies we have data for in 2000 are in the lower 10 per cent of the residual distribution from the 2000 regression of equation (6). But why are transition economies so special in the price-income relationship? Potential answers might include:

- incomplete price liberalisation: the output of a formerly centrally planned economy (CPE) is not yet fully priced on the market, subsidisation drives a wedge between prices and costs. This holds especially for services, i.e., non-tradables.
- Output quality is systematically lower in a CPE than in a market economy (Frensch, 2004), and this may *cet. par.* bias comparative price levels downwards for a CPE; in fact, this refers to potential measurement problems in the ICP.
- A part of stabilisation packages, most transition economies devalued their nominal exchange rates considerably in the early nineties, well below PPP rates.

Obviously, the first two issues above are immediately associated with reform effort during transition. It is therefore interesting to note that, as Figure 3 suggests, reform

effort during transition is related to the size of the residual from equation (6), i.e., the more successful a transition country has been in terms of policy reform, the more it conforms to the world-wide benchmark relationship between comparative prices and per capita income. Notably, this link between reform effort and distance to the benchmark relationship between comparative prices and per capita income becomes stronger over time.

Figure 3: **Standardised residuals from equation (6) versus reform indices of transition economies**



Notes: Ref1 is the EBRD aggregate first phase reforms index on liberalisation and privatisation.

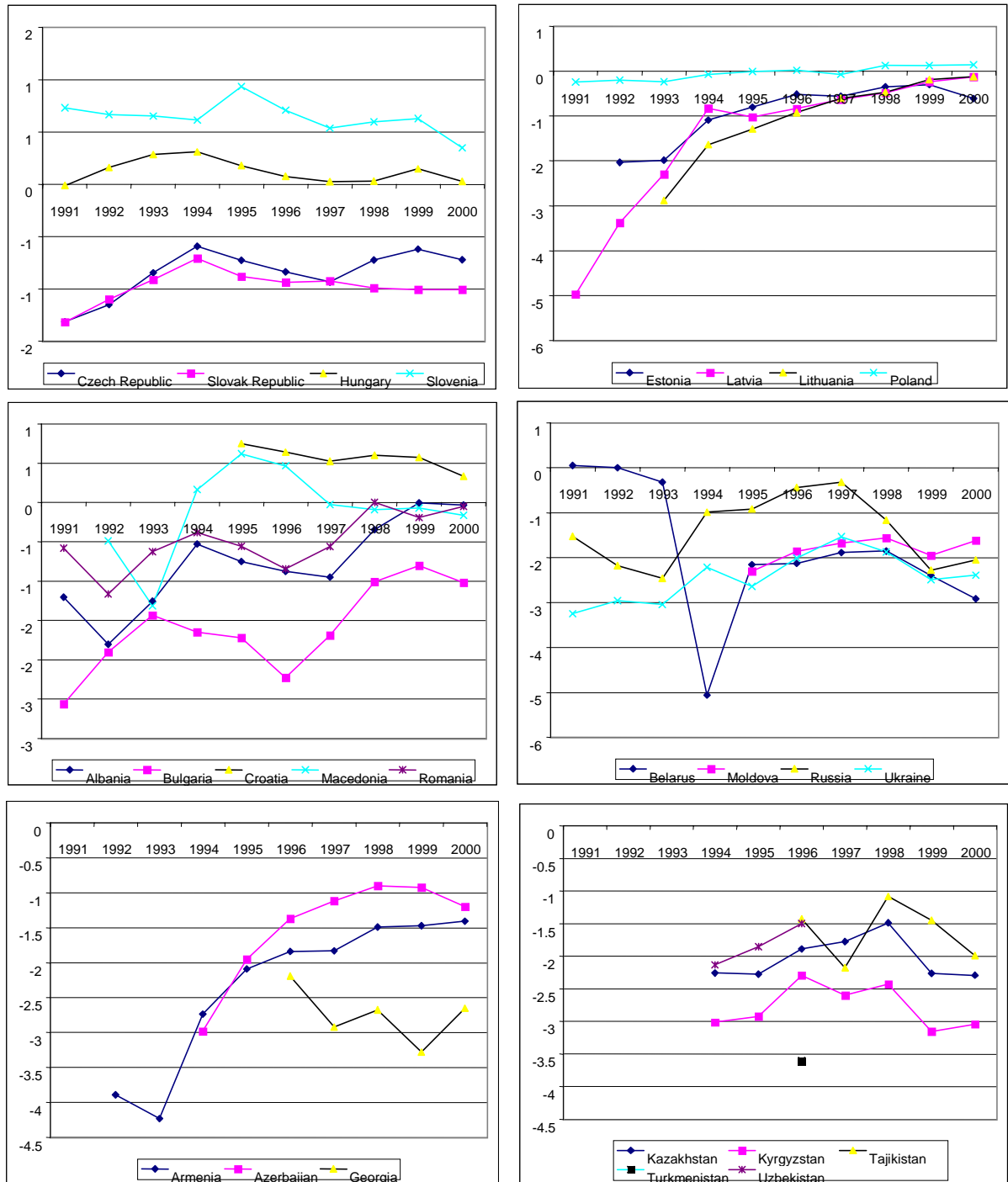
Figure 4 illustrates the “comparative statics” of transition countries’ distance to the world-wide benchmark relationship between comparative prices and per capita income in terms of the size of the residuals from equation (6) for the ten cross-section regressions between 1991 and 2000. The difference between country groups is evident: CEEC countries either appreciated in real terms to overcome an initial significant undervaluation or were never significantly outside the benchmark relationship to begin with (Slovenia, Hungary and Poland). CIS countries, on the other hand, either did not show significant real appreciation or when they did (Armenia and Azerbaijan) this was not sufficient to render the distance to the world-wide benchmark relationship between comparative prices and per capita income insignificant. Summing up the evidence so far:

- real exchange rates in transition countries at the beginning of the nineties were not just substantially undervalued in terms of PPP but these countries formed a group significantly *below* the world-wide benchmark relationship between comparative prices and per capita income (see also de Broeck and Sløk, 2006);
- at the end of the decade, this is no longer true for CEEC economies but still for the CIS;
- individual country progress in terms of overcoming the distance to the  $p$ - $y$  benchmark relationship appears to be related to reform effort.

We know that transition reforms had both direct effects on comparative prices, via initial liberalisation and further price deregulation, and indirect ones via productivity enhancing reforms. Thus, simple regressions such as (6) between real appreciation and

productivity growth during transition may be spurious, as both are due to reform efforts at least in part. We will return to this question in more detail in section 5.2.

Figure 4: Standardised residuals from (6) for selected transition countries, 1991–2000



## 5 Comparative price changes and per capita income growth

### 5.1 The simple BS view

The simplest *dynamic* version of a BS-based price-income relationship is equation (6) in growth terms over  $\tau$  periods, i.e.,

$$\ln p_{j,t+\tau} - \ln p_{jt} = \gamma_0 + \gamma_1 \cdot (\ln y_{j,t+\tau} - \ln y_{jt}) + \gamma_2 \cdot \text{OECD} \cdot (\ln y_{j,t+\tau} - \ln y_{jt}) + \varepsilon_{jt}. \quad (7)$$

Again both  $p$  and  $y$  are relative to U.S. levels. In estimating (7), we will, differently from the previous section, make use of the panel characteristics of our data. As our interest is mainly in transition economies, we estimate (7) over three four-year periods, 1988–92, 1992–96 and 1996–2000 with panel least squares, period fixed effects and White period-robust coefficient variance estimation to accommodate for arbitrary serial correlation and time-varying variances in the disturbances.

Table 2: **Panel least squares results for equation (7), 1988–2000**

Dependent variable: $\ln p_{j,t+4} - \ln p_{jt}$		
Explanatory variable	(1)	(2)
<i>Constant</i>	−0.05*** (−4.29)	−0.05*** (−4.22)
$\ln y_{j,t+4} - \ln y_{jt}$	0.02 (0.23)	
$\text{OECD} \cdot (\ln y_{j,t+4} - \ln y_{jt})$	0.78*** (2.76)	
$\ln \text{rgdp}_{j,t+4} - \ln \text{rgdp}_{jt}$		0.26* (1.75)
$\text{OECD} \cdot (\ln \text{rgdp}_{j,t+4} - \ln \text{rgdp}_{jt})$		0.39 (1.21)
Adjusted $R^2$	0.10	0.11
Observations	403	397

*Notes:* *rgdp* is an index of real GDP in local currency units, sourced from the World Bank's *World Development Indicators 2005*; here,  $\ln \text{rgdp}_{j,t+4} - \ln \text{rgdp}_{jt}$  is relative to the U.S. Estimation over three four-year periods, 1988–92, 1992–96 and 1996–2000 uses panel least squares with period fixed effects and White period standard errors and covariance;  $t$ -statistics in parentheses. \*, (\*\*), (\*\*\*) indicates significance at the 10, (5), (1) per cent level.

Column (1) in Table 2 reveals that there is no general support for a simple dynamic  $p$ - $y$ -relationship during the 1990s, except for OECD countries if per capita income is

measured in current year PPP-adjusted international prices.<sup>6</sup> One explanation for the poor explanatory power of equation (7), as opposed to the static version in (6), might be seen in too much nominal disturbance, which weighs more on rates of change than on level data.

Another explanation for this result, however, may be found in the way internationally comparable income data are constructed (see Appendix B). The essence of the discussion there is that growth rates derived from PWT do differ from those derived from national SNAs and that the recommendation is to use PPP-adjusted per capita income *level* data from PWT and real *growth* data from national accounts. Applying this to estimating (7) gives rise to the results reproduced in column (2) of Table 2: based on constant price data from national SNAs, there is general support for the existence of a dynamic  $p$ - $y$  relationship during the 1990s – and even in the same order of magnitude as in the static version (see Table 1) – but not for a distinct relationship for OECD countries.

Figure A-4 in the appendix reproduces the standardised residuals from this estimation for transition countries' real appreciation. For 15 out of 23 transition countries in Figure A-4 we have data on more than one four-year interval. In addition to the strong appreciation during the early nineties already noticed in section 4, the majority of these cases features a pattern of normalisation of the behaviour of residuals: a “significant overshooting” in the early nineties is followed by a deviation of less than one standard deviation from the dynamic benchmark relationship later in the decade.

## 5.2 An extended view

### 5.2.1 Equilibrium forces versus adjustment to equilibrium

So far we have relied on a simple BS-based price-income relationship with productivity as the only determinant of the real exchange rate, as stated in equation (4). The next logical step must consist of increasing the explanatory power of the approach by incorporating other sources that contribute to explaining transition and non-transition countries' real exchange rate behaviour. In fact, this is very much in line with other recent approaches. E.g., de Broeck and Sløk (2006) provide evidence that CEEC countries' real exchange rate appreciation during the early transition phase can be well explained in an international perspective within a behavioural equilibrium exchange rate approach.

Maeso-Fernandez et al. (2005) notice that many empirical approaches do not describe the transition process properly, where transition means adjustment to rather than fluctuating around a long-run equilibrium relationship. Ignoring this difference may imply many pitfalls. To avoid those, Maeso-Fernandez et al. (2005) offer two rules to be followed, one general, and one transition-specific. In general they require: “Following the

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<sup>6</sup> There is no qualitative change to this result when applying constant price measures of PPP-adjusted per capita income. To obtain such measures, see the PWT site at [http://pwt.econ.upenn.edu/php\\_site/pwt\\_index.php](http://pwt.econ.upenn.edu/php_site/pwt_index.php)

Balassa–Samuelson arguments, a higher level of productivity in industrialised countries relative to developing economies (proxied by GDP per capita) should be reflected in higher prices for nontraded goods relative to traded goods. In addition, demand-side factors – possibly related to non-homothetic preferences in the demand for nontraded services with luxury goods characteristics – as well as price regulation and tax policies are likely to influence the relative price of non-traded and traded goods in an economy” (p. 138). More specifically transition-related, they recommend always placing real exchange rate development of transition economies in an international context. We argue that both requirements together are not met so far in the literature.

The previous sections have clearly shown that transition economies were not part of the simple cross-section  $p$ - $y$  relationship in the early nineties; by the end of the decade, CEEC-countries have become part of this relationship. Therefore, when estimating a two-variable  $p$ - $y$  regression exclusively within the CEEC country group over the decade, and especially when adding the time-series dimension, the estimated price-income elasticity must necessarily be higher than that found for the benchmark relationship. However, even multi-variable approaches indicate that price-income elasticities in transition economies are either very high or explicitly higher than elsewhere. Following Dobrinsky (2003), the elasticity of CPI-based real exchange rate indices with respect to productivity is about 1. De Broeck and Sløk (2006) find that price-productivity elasticities in transition economies are explicitly higher than in OECD economies. Against the background of the discussion in section 3, this must come as a major surprise: if anything, price-income elasticities are higher for high income than for low income countries.

Given that reform is what differentiates transition countries, one would suspect that this phenomenon must have something to do with reform effort. In fact, de Brock and Sløk (2006) relate their result to reform efforts spurring productivity gains. This, of course, cannot be a sufficient explanation: whatever the effect of reforms on output or growth may be, why is the price-income elasticity so high in transition economies? In our view, as already indicated in Figure 3 above, high CEEC  $p$ - $y$  elasticity coefficients in the literature signal an adjustment to the benchmark relationship between comparative prices and per capita income due to transition reforms, as transition reforms have – over and above indirect effects on comparative prices via productivity gains – also direct effects on comparative prices via initial price liberalisation and subsequent deregulation. We thus maintain that high price-income elasticity estimates for CEEC countries in the literature are upward biased due to misspecification by omitting an important explanatory variable: some of this elasticity should be attributed to direct reform impetus on deregulating prices.

Our approach is therefore close in spirit to Coricelli and Jazbec (2001) in attempting to highlight the role of reforms, especially of price deregulation, for real exchange rate developments during transition.<sup>7</sup> However, we extend Coricelli and Jazbec (2001) by putting our approach in an international context.

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<sup>7</sup> Coricelli and Jazbec (2001) relate reform to real exchange rates during transition mostly by using reform indicators as instruments for transition induced reallocation.

## 5.2.2 “Real” sources of real exchange rate development during transition and beyond

Also much like Coricelli and Jazbec (2001), in addition to price deregulation efforts we focus on the role of real factors in this process. With Blanchard (1997) we define transition as resource *reallocation*, corporate *restructuring*, and *liberalisation*. In the simple set-up of equations (3) and (4), the only alternative to a deepening productivity gap to imply a more pronounced BS-type relationship was by a rise in the share of non-traded goods in GDP, which seems heavily at odds with recent empirical developments. The extended argument in Frensch (2005), on which we build here, however, allows to separate tradability from *reallocation* in terms of changes in income shares spent on services and industrial goods (see also Frensch, 2000). For further analysis, we return to the arbitrage view of the classic BS set-up in section 2, however extending the framework to incorporate transition. For ease of exposition, we change to logarithmic notation and omit time for the moment. Then following the notation in section 2,

$$\ln RER_{21} = \ln P_2 - \ln P_1 - \ln e_{12}. \quad (8)$$

Rather than differentiating only between tradables and non-tradables, we assume economies to have two sectors, industry ( $I$ ) and services ( $S$ ), with products entering national price levels with potentially different weights such that,

$$\ln P_j = \phi_j \ln P_j^I + (1 - \phi_j) \ln P_j^S. \quad (9)$$

We make a few simplifying but well-grounded assumptions to modify the set-up of section 2:

(A1) While all services are non-tradable, only some part of industrial goods,  $\tau_j$ , is tradable due to the existence of barriers to trade, i.e.,

$$\ln P_j^I = \tau_j \ln P_j^{I,T} + (1 - \tau_j) \ln P_j^{I,NT}. \quad (10)$$

(A2) Prices are proportional to unit labour costs,

$$\ln P_j^h = \lambda^h + \ln w_j - \ln A_j^h, \quad (11)$$

where  $h = S; I, T; I, NT$ ,  $w$  is the wage rate and  $A$  is labour productivity, which is the same in all of industry.

(A3) Exposure to international trade increases the intensity of competition, i.e.,

$$\lambda^S = \lambda^{I,NT} = \lambda^{NT} > \lambda^I = \lambda^{I,T}. \quad (12)$$

(A4) Purchasing power parity, as usually, does not hold for non-tradables; for tradables, PPP is restricted by quality differentials according to



$$\ln P_2^{I,T} = \ln P_1^{I,T} + \ln e_{12} + \kappa_{21}^{I,T}, \quad (13)$$

where country 2 product quality of tradables,  $\kappa_{21}^{I,T}$ , is defined relative to country 1.

From equations (8) and (13),

$$\ln RER_{21} = (\ln P_2 - \ln P_2^{I,T}) - (\ln P_1 - \ln P_1^{I,T}) + \kappa_{21}^{I,T}, \quad (14)$$

where (9) implies that

$$\ln P_j - \ln P_j^I = (1 - \phi_j)(\ln P_j^S - \ln P_j^I), \quad (15)$$

and from (10)

$$\ln P_j^I - \ln P_j^{I,T} = (1 - \tau_j)(\ln P_j^{I,NT} - \ln P_j^{I,T}). \quad (16)$$

From (15) and (16),

$$\begin{aligned} \ln P_j - \ln P_j^{I,T} &= (1 - \phi_j)(\ln P_j^S - \ln P_j^I) + (1 - \tau_j)(\ln P_j^{I,NT} - \ln P_j^{I,T}) \\ &= (1 - \phi_j)\ln P_j^S - \ln P_j^I + \phi_j \ln P_j^I + (1 - \tau_j)(\ln P_j^{I,NT} - \ln P_j^{I,T}). \end{aligned} \quad (17)$$

Substituting from (16),

$$\ln P_j - \ln P_j^{I,T} = (1 - \phi_j)\ln P_j^S - \ln P_j^{I,T} + \phi_j \ln P_j^I,$$

and from (10),

$$\begin{aligned} \ln P_j - \ln P_j^{I,T} &= (1 - \phi_j)\ln P_j^S - \ln P_j^{I,T} + \phi_j \ln P_j^{I,T} + \phi_j(1 - \tau_j)(\ln P_j^{I,NT} - \ln P_j^{I,T}) \\ &= (1 - \phi_j)(\ln P_j^S - \ln P_j^{I,T}) + \phi_j(1 - \tau_j)(\ln P_j^{I,NT} - \ln P_j^{I,T}). \end{aligned} \quad (18)$$

Substituting for prices according to (11) and collecting terms yields

$$\ln P_j - \ln P_j^{I,T} = (1 - \phi_j)(\ln A_j^I - \ln A_j^S) + (1 - \tau_j \phi_j)(\lambda^{NT} - \lambda^T). \quad (19)$$

Then, equation (14) implies,

$$\ln RER_{21} = [(1 - \phi_2)(\ln aA_2^I - \ln A_2^S) - (1 - \phi_1)(\ln A_1^I - \ln A_1^S)] + \kappa_{21}^{I,T} + (\tau_1 \phi_1 - \tau_2 \phi_2)(\lambda^{NT} - \lambda^T). \quad (20)$$

After total differentiation and again collecting terms, we decompose the rate of change of the real exchange rate of country 2 relative to country 1 into four separate effects (where a  $\Delta$  of a logarithmic value indicates a growth rate),

$$\Delta \ln RER_{21} = \tag{21}$$

$$(1 - \phi_2)(\Delta \ln A_2^I - \Delta \ln A_2^S) - (1 - \phi_1)(\Delta \ln A_1^I - \Delta \ln A_1^S)$$

(a) Productivity gap between industry and services

$$+ \Delta \kappa_{21}^{I,T}$$

(b) Quality improvement of tradables

$$+ \Delta \phi_1 [(\ln A_1^I - \ln A_1^S) + \tau_1 (\lambda^{NT} - \lambda^T)] - \Delta \phi_2 [(\ln A_2^I - \ln A_2^S) + \tau_2 (\lambda^{NT} - \lambda^T)]$$

(b) Sectoral reallocation between industry and services

$$+ (\phi_1 \Delta \tau_1 - \phi_2 \Delta \tau_2) (\lambda^{NT} - \lambda^T).$$

(d) Trade liberalisation

Compared to section 2, the slight twist in sectoral decomposition, by adding economic activity categories to the tradable-non-tradable dichotomy, is quite fruitful: remember that in the simple set-up, the only alternative to BS for  $P_2/P_1$  to increase was by a rise in the share of non-traded goods in GDP, which seems heavily at odds with empirical developments. The argument here, however, allows to separate tradability from income shares spent on services and industrial goods. This allows to show that, in addition to the productivity gap effect, reallocation from industry towards services in country 2, relative to country 1 ( $\Delta \phi_2 < 0$ ), also implies a real exchange rate appreciation assuming that productivity in industry is higher than in services. A unilateral reduction in country 2 versus country 1 foreign barriers to trade in industrial products ( $\Delta \tau_2 > 0$  and  $\Delta \tau_1 = 0$ ) implies a real depreciation for country 2. Symmetric reduction in barriers to trade ( $\Delta \tau_1 = \Delta \tau_2 > 0$ ) implies a depreciation for country 2 as long as the share of this country's services sector in total production is smaller than in country 1. While all of these three phenomena are supposed to be specifically pronounced during transition, in fact they occur elsewhere and at other times as well, allowing us to study real exchange rates of non-transition countries within the same framework.

### 5.2.3 Data and measurement

Price deregulation is a continuous process during transition, goes well beyond the initial price liberalisation, and also includes reform efforts that have price deregulation implications, such as privatisation. In terms of measuring price deregulation, the EBRD average indicator of stage 1 reforms, i.e., the *liberalisation plus privatisation* index (*Ref1*) therefore seems appropriate. For non-transition economies, we state that during the 1990s

there was no reform effort comparable in order of magnitude to what happened at the same time in transition economies.

We measure *reallocation* by the change in the nominal GDP share of services, drawn from the World Bank's *WDI 2005*. Regrettably, we do not have an appropriate independent measure for quality improvement, so that corporate *restructuring* and ensuing quality improvements affects per capita income over and above productivity growth. Given the notorious difficulties in measuring barriers to trade, we proxy *trade liberalisation* by the result, i.e., by the change in a country's openness to trade over the period under consideration, as measured by the PPP-adjusted GDP share of total trade provided in the PWT dataset.

Country coverage is driven by the aim of the exercise, i.e., putting CEEC real exchange rate behaviour in an international perspective. In order to reduce country heterogeneity in the following panel approach, we include only economies with  $10 < y_j < 110$ , with  $y_{US} = 100$ ; for the composition of CEEC country group data, this has the slight consequence that Albania is part of the panel with only 3 rather than with all 6 observations available during the 1990s.

#### 5.2.4 Specification and estimation results

The extensions of the simple BS set-up discussed so far give way to hypothesising our measure of the real exchange rate as,

$$\ln p_{jt} = \gamma_0 + \gamma_1 \cdot \ln y_{jt} + \gamma_2 \cdot \text{OECD} \cdot \ln y_{jt} + \gamma_3 \cdot \text{Services}_{jt} + \gamma_4 \cdot \text{Open}_{jt} + \varepsilon_{jt}, \quad (22a)$$

for non-transition economies, and

$$\ln p_{jt} = \gamma_5 + \gamma_6 \cdot \ln y_{jt} + \gamma_7 \cdot \text{Services}_{jt} + \gamma_8 \cdot \text{Open}_{jt} + \gamma_9 \cdot \text{Ref1}_{jt} + v_{jt}, \quad (22b)$$

for CEEC economies, with the *a priori* expectation that  $\gamma_1, \gamma_2, \gamma_3, \gamma_6, \gamma_7, \gamma_9 > 0$  and  $\gamma_4, \gamma_8 < 0$ .

While the cross-section analysis of section 4 provided a first indication of the *p-y* relationship, it ignored the time-series information in the data. In order to incorporate this information, we need to adopt a panel data approach. We estimate equation (22) in two specifications: in levels (logs) and in differenced logs, i.e. in yearly rates of change, both against two unbalanced panels of countries between 1992 and 2000: the first panel consists of non-transition PWT countries, the second is made up of the CEEC economies.

Choosing an appropriate panel specification for estimating (22) is crucial. The problem lies in adding a time series dimension to data on a hypothesis, which is originally a statement on cross-sections. Both the level and the dynamic specifications may involve potential violations of the assumptions of the classical linear regression model along time and cross-section dimensions, which may be especially serious given our additional constraint that the limited length of the time series prevents the use of cointegration techniques.

In *level* specification, country effects might restrict the variation in the data along the time-series dimension but if most of the variation were indeed along the cross-section dimension, country effects might bias against finding a BS-effect. On the other hand, “identifying the BS-effect from a time series correlation could be misleading, since high-frequency business-cycle correlations of the real exchange rate with output fluctuations (arising from quite different mechanisms) might cloud the picture” (Bergin et al., 2004, p. 6). Under these circumstances, Bergin et al. (2004) recommend estimating with a common intercept and a common AR(1) term for all countries, where the estimate of the AR(1) parameter can then be read as the half-life of deviations from equilibrium real exchange rates. Accordingly, we test in levels by generalised least squares with cross-section weights,<sup>8</sup> which takes into account and corrects for the presence of cross-section heteroskedasticity, as suggested by Dobrinsky (2003), and include an auto-regressive term.

When including a transition reform variable as an explanatory variable, level estimation involves potentially serious problems of common trends and collinearity among regressors: during transition, virtually everything is interdependent with reform effort, and *Ref1*, i.e. the EBRD reform indicator in levels, obeys a time trend. We correct for this by detrending the reform variable against a linear time trend; this detrended version of *Ref1* enters the estimation as *Ref1\_d*, minimising both the time trend problem and collinearity with other explanatory variables.

In the *dynamic* specification, auto-regressive terms do not appear feasible, a generalised least squares approach therefore has to stress the assumption of either cross-section or period heteroskedasticity and correlation. As cross-sectional weighting leaves any correlation other than contemporaneous unaddressed, we prefer period SUR weights correcting both for period heteroskedasticity and general correlation of observations within any cross-section.

As the level equation is in logs, estimated coefficients can be interpreted as long-run elasticities. Likewise, in the dynamic specification, the estimated coefficients can be interpreted as short-run elasticities, where in both specifications our main attention is for the price-income elasticity. We follow the previously introduced recommendation to use PPP-adjusted per capita income *level* data from PWT and real *growth* data from national accounts. Finally, in both the level and the dynamic specifications, we estimate (22b) with and without the reform variable.

Tables 3a and b present the results of estimating (22) in level and in dynamic specification, respectively. For the panel of non-transition economies (columns 1 and 4), different from earlier sections the OECD variable is no longer significant in this extended approach in either specification. All other coefficients have the expected signs and are significant. Very noticeably, the point estimates of the long-run and the short run price-income elasticities in the OECD sample are the same. Level and dynamic estimates of the coefficients of the services and openness variables are in the same order of magnitude.

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<sup>8</sup> The more preferred cross-section SUR approach is not feasible, as in our panel the number of cross-sections exceeds the number of periods.

Table 3a: **Panel EGLS estimation results for equation (22), 1992–2000, in levels**

Dependent variable: $\ln p_{jt}$			
Explanatory variable	(1)	(2)	(3)
<i>Constant</i>	2.95 <sup>***</sup> (15.65)	2.79 <sup>***</sup> (8.19)	3.29 <sup>***</sup> (8.03)
$\ln y_{jt}$	0.21 <sup>***</sup> (4.64)	0.28 <sup>***</sup> (3.11)	0.16 (1.42)
<i>OECD</i> • $\ln y_{jt}$	0.001 (0.08)		
<i>Services</i> <sub><i>jt</i></sub>	0.007 <sup>***</sup> (5.94)	0.003 <sup>*</sup> (1.79)	0.002 (0.94)
<i>Open</i> <sub><i>jt</i></sub>	−0.0009 <sup>***</sup> (−9.57)	−0.0007 <sup>***</sup> (−4.33)	−0.0006 <sup>***</sup> (−4.29)
<i>Ref1_d</i> <sub><i>jt</i></sub>			0.58 <sup>**</sup> (2.46)
AR(1)	0.95 <sup>***</sup> (117.17)	0.68 <sup>***</sup> (21.97)	0.71 <sup>***</sup> (21.32)
Adjusted $R^2$	0.99	0.91	0.92
Sample (total observations)	82 non-transition PWT countries (657)	13 CEEC countries (103)	13 CEEC countries (103)

*Notes:*  $p$ ,  $y$ , *Services* and *Open* and *Ref1* are all relative to the U.S. *Ref1\_d* is the detrended version of *Ref1*. Estimation is by panel EGLS with cross-section weights over unbalanced samples of countries with  $10 < y_{jt} < 110$ , 1992–2000;  $t$ -statistics in parentheses. \*, (\*\*), (\*\*\*) indicates significance at the 10, (5), (1) per cent level. CEEC countries in the sample (columns 2, 3); see Table A-1.

Tables 3a and b present the results of estimating (22) in level and in dynamic specification, respectively. For the panel of non-transition economies (columns 1 and 4), different from earlier sections the OECD variable is no longer significant in this extended approach in either specification. All other coefficients have the expected signs and are significant. Very noticeably, the point estimates of the long-run and the short run price-income elasticities in the OECD sample are the same. Level and dynamic estimates of the coefficients of the services and openness variables are in the same order of magnitude.

Table 3b: Panel EGLS estimation results for equation (22), 1992–2000, in yearly changes

Dependent variable : $\ln p_{jt} - \ln p_{j,t-1}$			
Explanatory variable	(4)	(5)	(6)
<i>Constant</i>	4.24 <sup>***</sup> (7.98)	8.46 <sup>***</sup> (4.20)	3.73 <sup>*</sup> (1.92)
$\ln rgdp_{jt} - \ln rgdp_{j,t-1}$	0.21 <sup>***</sup> (7.99)	0.41 <sup>***</sup> (4.21)	0.18 <sup>*</sup> (1.94)
<i>OECD</i> •( $\ln rgdp_{jt} - \ln rgdp_{j,t-1}$ )	0.0001 (0.44)		
$Services_{jt} - Services_{j,t-1}$	0.005 <sup>***</sup> (4.27)	0.007 <sup>***</sup> (4.00)	0.006 <sup>***</sup> (3.17)
$Open_{jt} - Open_{j,t-1}$	-0.001 <sup>***</sup> (-10.76)	-0.0004 <sup>**</sup> (-2.27)	-0.0008 <sup>***</sup> (-5.26)
$Ref1_{jt} - Ref1_{j,t-1}$			0.58 <sup>***</sup> (5.72)
Adjusted $R^2$	0.28	0.31	0.50
Sample (total observations)	81 non-transition PWT countries (650)	13 CEEC countries (103)	13 CEEC countries (103)

Notes:  $p$ ,  $y$ , *Services* and *Open* and *Ref1* are all relative to the U.S. Estimation is by panel EGLS with cross-section weights over unbalanced samples of countries with  $10 < y_{jt} < 110$ , 1992–2000;  $t$ -statistics in parentheses. \*, (\*\*), (\*\*\*) indicates significance at the 10, (5), (1) per cent level. CEEC countries in the sample (columns 5 and 6): see Table A-1.

In the CEEC sample, reform effort is a significant variable in both the level and the dynamic specification (columns 3 and 6). Accordingly, introducing reform effort always implies a reduction of the point-estimates of the price-income elasticities in the CEEC sample down to the benchmark order of magnitude in columns 1 and 4, respectively. This is especially noteworthy in the dynamic version, where on the basis of simple Wald-test, we can reject the hypothesis that the price-income-elasticity for transition economies estimated in column 5 is equal to the benchmark estimate in column 4.

These results lead us to conclude that during the 1990s, when accounting for the direct influence of reform effort on comparative prices, the price-income-elasticity for transition economies is not different from non-transition economies in the same income range. However, as indicated by the auto-regressive terms in column 3, the speed of adjustment to equilibrium appears slower in transition economies than elsewhere.

These conclusions are based on the notion that direct reform effort is independent: however, Campos and Coricelli (2002, p. 828) note the “issue of correlation between initial conditions and liberalization measures. One can argue that the extent of liberali-

zation and the speed of reform are not independent of initial conditions.” This debate is far from over, but at least for the first decade of transition it seems that if anything it is indeed initial conditions – rather than current or most recent economic performance over and above that induced by initial conditions – that shape progress in reform (see Falcetti et al., 2002 and Godoy and Stiglitz, 2006). While resolving this issue cannot be a subject of this paper, it may perhaps not really matter whether it is initial conditions or independent reform effort that subtract from the price-income elasticity when properly accounting for either.

## **6 Conclusions**

The paper puts real exchange rate developments of transition economies into an international perspective. To this end, we first illustrate the special status of transition economies in the world-wide benchmark relationship between comparative prices and per capita income: a pronounced undervaluation at the start of transition, followed by a transition-specific pattern of strong appreciation during the early nineties, results in “normalisation” for CEEC economies but not for the CIS. We then make an attempt at extending the BS framework. The results of this exercise imply that, when accounting for demand shifts, external liberalisation, and especially for reform effort, the price-income-elasticity for CEEC economies is not different from that of non-transition economies during the nineties.

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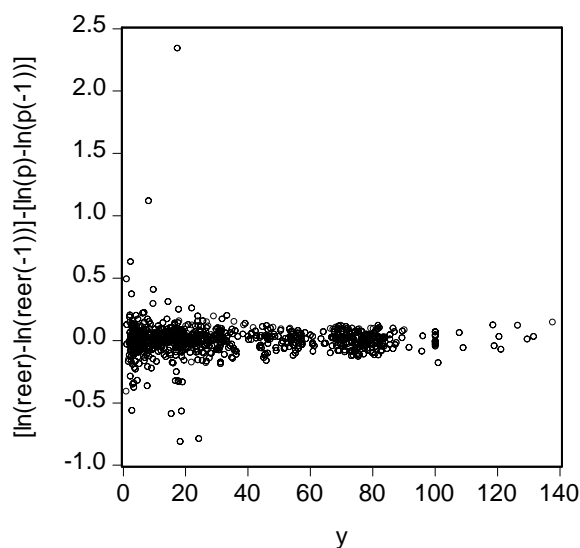


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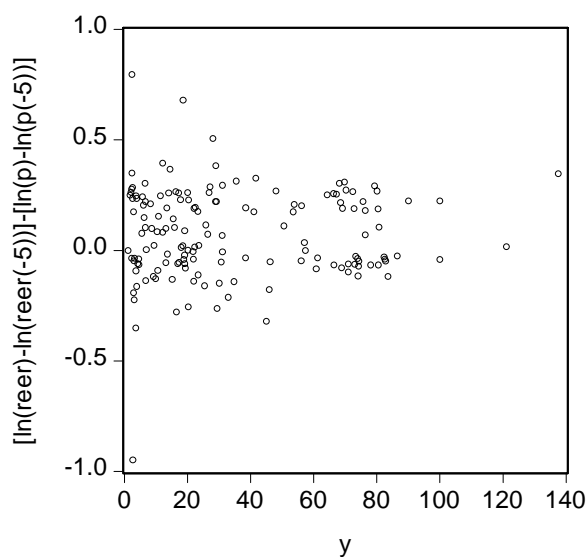
## Appendix A: Tables and Figures

Figure A-1: **Growth differentials between trade weighted CPI-deflated real effective exchange rates,  $reer_j$  and comparable price levels,  $p_j$  versus PPP-adjusted income per capita,  $y_j$ , 1990–2000**

a) one-year differentials

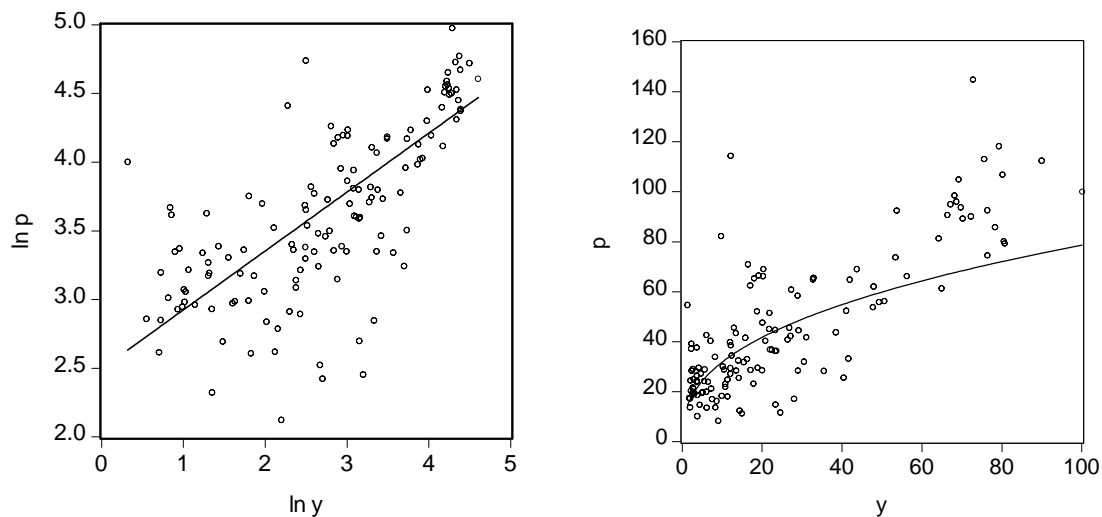


b) five-year differentials



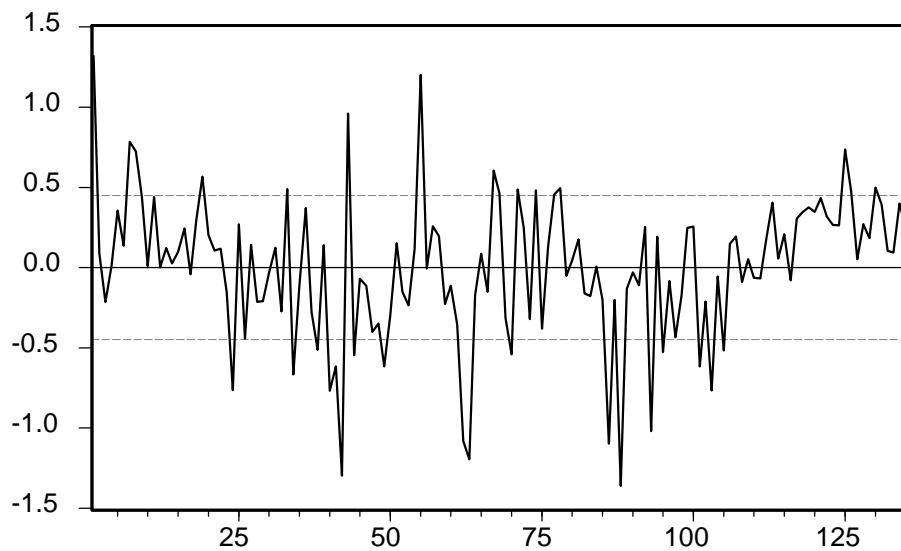
Notes:  $reer_j$  data are from IFS, for data on  $p_j$  and  $y_j$  see Figure 1. Sample size: 864 observations in a), and 153 in b). Simple OLS regressions between growth rate differentials of both measures and  $y_j$  produce insignificant slope coefficients.

Figure A-2a: **Comparative price level versus PPP-adjusted income per capita, in logs, 2000**



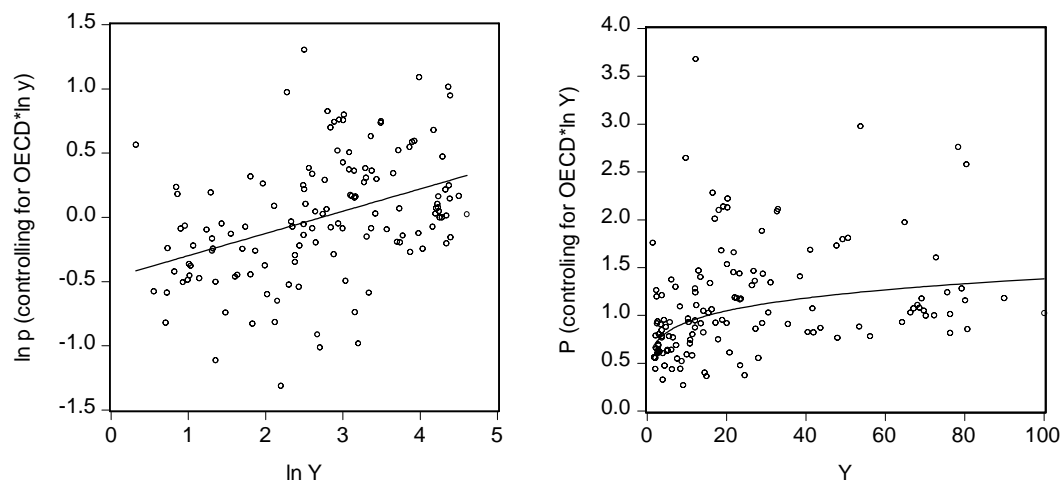
Notes: See Appendix A, Figure A-1. Sample size: 135 ( $y \leq 100$ ).

Figure A-2b: **Residuals from  $\ln p_j = c(1) + c(2) \cdot \ln y_j$ , for the year 2000**



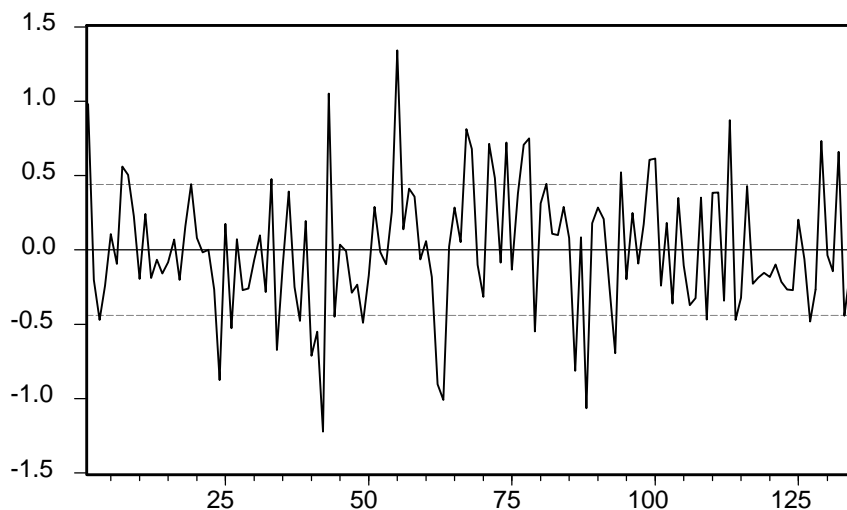
Notes: See Appendix A, Figure A-1. 135 observations in ascending order of  $y_j (\leq 100)$ . Breusch-Godfrey Serial Correlation LM Test (6 lags included) indicates serial correlation at 5 per cent level of significance. No significant White statistic on heteroskedasticity.

Figure A-3a: Comparative price level versus PPP-adjusted income per capita, in logs, 2000. Partial relationship, controlling for OECD•ln y<sub>j</sub>



Notes: See Appendix A, Figure A-11. Sample size: 135 ( $y \leq 100$ ).

Figure A-3b: Residuals from  $Resid_j = c(1) + c(2) \cdot \ln y_j$ , for the year 2000



Notes:  $Resid_j$  are the residuals from  $\ln p_j = c(3) + c(4) \cdot OECD \cdot \ln y_j$  for the year 2000, where  $OECD$  is a dummy variable (see text for explanation). 135 observations in ascending order of  $y_j (\leq 100)$ . Breusch-Godfrey Serial Correlation LM Test (6 lags included) does not indicate serial correlation. No significant White statistic on heteroskedasticity.

Table A-1: **Country list**

OECD members	Transition countries	Other countries
	<i>CEEC</i>	
Australia, Austria, Belgium, Canada, Switzerland, Denmark, Spain, Finland, France, United Kingdom, Germany, Greece, Ireland, Iceland, Italy, Japan, Luxembourg, Mexico, Netherlands, Norway, New Zealand, Portugal, Sweden, Turkey, United States	Albania, Bulgaria, Czech Republic, Estonia, Croatia, Hungary, Lithuania, Latvia, Macedonia, Poland, Romania, Slovak Republic, Slovenia	All other PWT countries. For full country composition, see the PWT site at <a href="http://pwt.econ.upenn.edu/php_site/pwt_index.php">http://pwt.econ.upenn.edu/php_site/pwt_index.php</a>
	<i>CIS</i>	
	Armenia, Azerbaijan, Belarus, Georgia, Kazakhstan, Kyrgyzstan, Moldova, Russia, Tajikistan, Turkmenistan, Ukraine, Uzbekistan	
	<i>Other</i>	
	China, Vietnam	

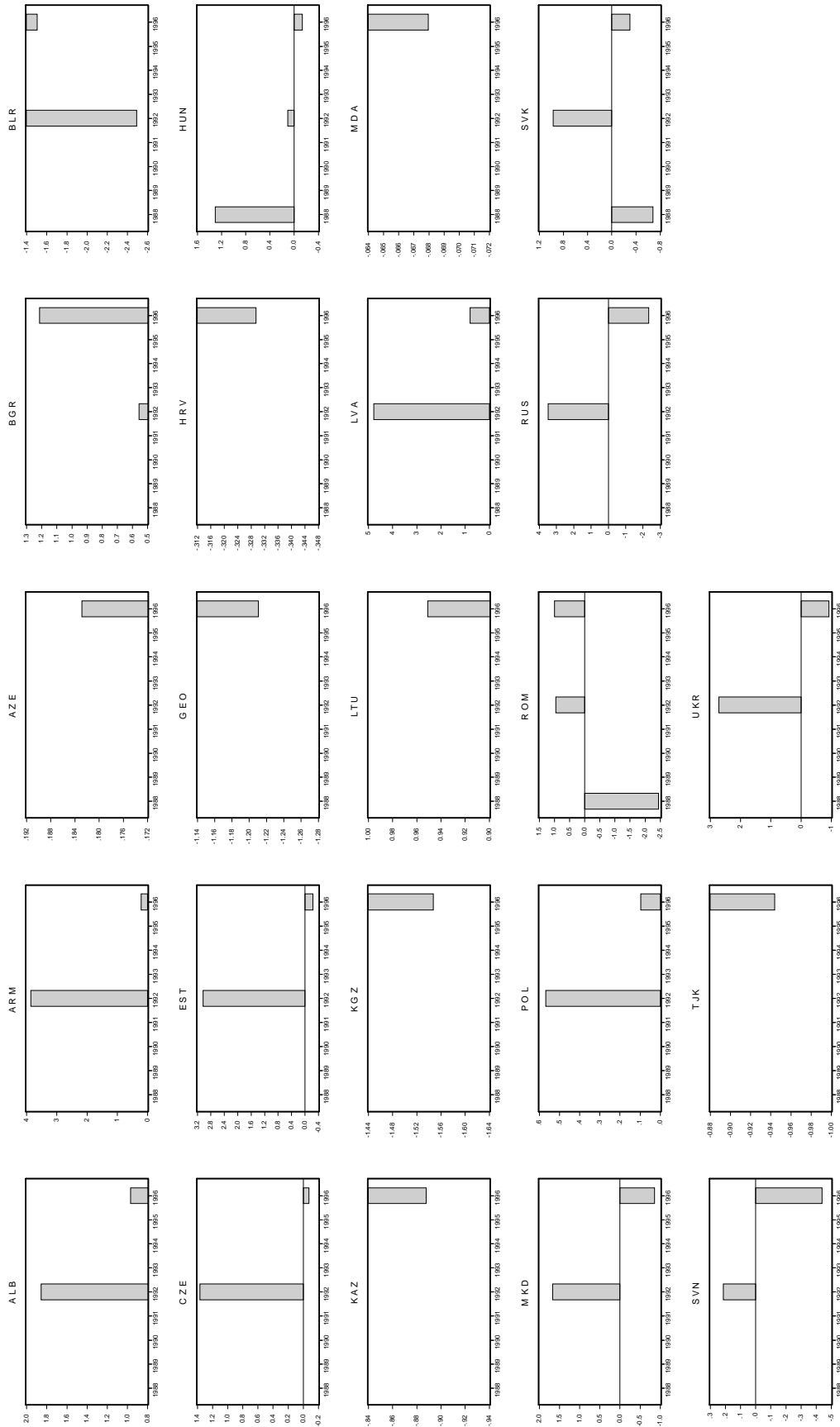
Table A-2: Extrema of the distributions of residuals from regression (6), 1991–2000

	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000
	Upper 10 per cent (upper 5 per cent) of the distribution									
Iran	Iran	Iran	Yemen	Yemen	Nigeria	Bermuda	Nigeria	Nigeria		
Yemen	Yemen	Yemen	Angola	Angola	Yemen	Yemen	Syria	Syria	Syria	
Angola	Syria	Angola	Nigeria	Nigeria	Syria	Tanzania	Tanzania	Tanzania	Tanzania	Syria
Israel	Israel	Dem Rep. Congo	Syria	Syria	Rep. of Congo	Rep. of Congo	Jamaica	Jamaica	Jamaica	Tanzania
Seychelles	Rep. of Congo	Nigeria	Dem Rep. Congo	Dem Rep. Congo	Dem Rep. Congo	Israel	Dem Rep. Congo	Lebanon	Lebanon	Jamaica
Argentina	Angola	Israel	Israel	Rep. of Congo	Israel	Zambia	Israel	Puerto Rico	Costa Rica	Costa Rica
Syria	Argentina	Rep. of Congo	Seychelles	Israel	Zambia	Puerto Rico	Zambia	Israel	Zambia	Nigeria
Rep. of Congo	Taiwan	Taiwan	Taiwan	Seychelles	Singapore	Equatorial Guinea	Puerto Rico	Costa Rica	Venezuela	Israel
Taiwan	Cameroon	Singapore	Singapore	Taiwan	Taiwan	Lebanon	Lebanon	Rep. of Congo	Israel	Venezuela
Dem Rep. Congo	Seychelles	Seychelles	Zambia	Tanzania	Tanzania	Togo	Costa Rica	Rep. of Congo	Grenada	Grenada
Singapore	Singapore	Argentina	Argentina	Japan	Japan	Brazil	Grenada	Grenada	St. Lucia	St. Lucia
Haiti	Senegal	Gabon	Grenada	Rep. of Korea	Rep. of Korea	Angola	Brazil	Venezuela	Rep. of Congo	Zambia
Senegal	Gabon	Hong Kong	Hong Kong	Hong Kong	Argentina	Grenada	Hong Kong	Hong Kong	Hong Kong	Iran
Cyprus	Cyprus	Senegal	Cyprus	Cyprus	Equatorial Guinea	Madagascar	Singapore	Brazil	Iran	Lebanon
							Rep. of Congo	St. Lucia	Nigeria	Yemen

Table A-2 contd.: **Extrema of the distributions of residuals from regression (6), 1991–2000**

		Lowest 10 per cent (5 per cent ) of the distribution									
		1991	1992	1993	1994	1995	1996	1997	1998	1999	2000
China					Nepal	Mauritius	Azerbaijan	Turkey	Russia		
Sierra Leone	Sierra Leone			India	Russia	Vietnam	India	India	Rep. of Korea	Nepal	India
Slovak Republic	Slovak Republic			Albania	China	Lithuania	Tajikistan	India	India	Mauritius	Azerbaijan
Guyana	Czech Republic			Kenya	Estonia	Guinea	Mauritius	Ukraine	Armenia	Tajikistan	Indonesia
Guinea	Romania			Macedonia	Vietnam	Mexico	Uzbekistan	Nepal	Kazakhstan	Armenia	Mauritius
Albania	Nepal			Mozambique	Turkey	Nepal	Nepal	Guinea	Moldova	Indonesia	Armenia
Egypt	Mozambique			Nepal	Lithuania	Bulgaria	Armenia	Vietnam	Nepal	Zimbabwe	Moldova
Czech Republic	Vietnam			Bulgaria	Bulgaria	Uzbekistan	Moldova	Moldova	Guinea	Guinea	Guinea
Slovak Republic	Albania			Vietnam	Uzbekistan	Azerbaijan	Kazakhstan	Bulgaria	Zimbabwe	Moldova	Tajikistan
Nepal	Bulgaria			Estonia	Ukraine	Armenia	Ukraine	Kazakhstan	Mauritius	Kazakhstan	Russia
Russia	Estonia			Latvia	Kazakhstan	Belarus	Belarus	Armenia	Belarus	Russia	Kazakhstan
Vietnam	Russia			Russia	Armenia	Kazakhstan	Georgia	Belarus	Ukraine	Belarus	Ukraine
Bulgaria	Ukraine			Lithuania	Azerbaijan	Moldova	Bulgaria	Tajikistan	Kyrgyzstan	Ukraine	Georgia
Ukraine	Latvia			Ukraine	Kyrgyzstan	Ukraine	Kyrgyzstan	Kyrgyzstan	Georgia	Kyrgyzstan	Belarus
Latvia	Armenia			Armenia	Belarus	Kyrgyzstan	Turkmenistan	Georgia	Indonesia	Georgia	Kyrgyzstan
<i>Number of observations</i>											
		138	141	142	147	149	167	146	146	140	135

Figure A-4: Standardised residuals from estimating equation 7. Table 2, column (2)





## Appendix B. PWT data construction and economic relationships

Making aggregate quantities, such as GDPs, internationally comparable always means comparable in a common currency. Consider there is more than one good in each economy. Expressing price ratios of goods in terms of a foreign currency then necessarily involves the imposition of the foreign price structure when constructing any weighted price index such as PPP over GDP. This problem, of course, has been recognised early in the ICP so that in the course of comparing prices of international goods and weighting them to price indices involves weighting the price structures of all the countries in the project such that the reference currency is not the U.S. dollar but rather a virtual currency resulting from the weighting procedure, the international dollar. In this virtual currency, “relative prices of goods are set at the ‘weighted’ average of relative prices for the same goods in all countries and the level of prices is normalized so that the GDP of the United States is the same in international dollars as in American dollars” (Summers and Heston, 1991, p. 334).

However, there is of course one real world country that comes closest to the price structure of this virtual currency. In fact, for the 1985 ICP benchmark data, this country can be shown to have been Hungary (Nuxoll, 1994), which means that the construction of price indices, such as the PPP over GDP, involves imposing Hungarian relative prices everywhere in the world. Then, stemming directly from the usual index number problem (Paassche versus Laspeyres), the Gerschenkron effect states that measured growth rates of comparable quantities depend on the underlying relative price structure.<sup>9</sup> I.e., if Hungary’s price structure is the relevant price structure for international comparisons, then PPP-adjusted per capita incomes and growth rates of countries richer (poorer) than Hungary are overstated (understated) by using Hungarian relative prices. Of course, there is a danger that any economic relationship involving these internationally comparable prices and quantities may be biased by the construction of the data, be it in the area of convergence debates or in the BS context.

Nuxoll (1994) shows that Gerschenkron effects are indeed present in the ICP data, which underlie the PWT, and that growth rates derived from PWT do differ from those derived from national SNAs. However, this difference is not significantly dependent on PPP-adjusted per capita income levels, most probably due to the high level of aggregation of PWT data (rather, the difference depends on the size of relative price changes in the period under consideration, and this is certainly an issue for transition economies). Still, in the context of the convergence debate, Nuxoll’s recommendation is to use per capita income level data from PWT and growth data from national accounts. So far, an effort to analyse the effects of potential bias from data construction on the  $p$ - $y$  relationship illustrated in Figure 1 is still missing.

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<sup>9</sup> Nuxoll (1996) gives an account of Gerschenkron’s original problem of properly measuring Soviet growth.

## **Balassa-Samuelson effects in the presence of product differentiation and trade barriers. Implications for transition and convergence**

### **Abstract**

: The paper first illustrates the simple arbitrage view of the Balassa-Samuelson effect and its implications on the relationship between comparative prices and real per capita income. Two aspects of this relationship are then illustrated and connected: in the context of trade in differentiated goods, a higher elasticity of substitution is shown to strengthen the BS-effect. Connecting indirect trade barriers to the elasticity of substitution, the strength of the BS-effect increases with trade liberalisation. Considering this result in an extended arbitrage view on prices and productivities reveals that there are no transition-specific forces left to weaken the BS-effect, when defining transition as liberalisation, reallocation, and restructuring. An immediate corollary is that inflation differentials between Central and East European economies and the euro-area should weaken over time. To the extent they do not, these inflation differentials increasingly signal disequilibria, rather than equilibrium phenomena.

*JEL*-Classification: F40, F43

Keywords: Balassa-Samuelson, transition, product differentiation

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## 1. Introduction

The empirical starting point for the paper is the existence of a significant cross-country relationship between comparative prices and real per capita income, as documented in Bergin et al. (2004), where “internationally comparable prices” refers to any measure of the deviation of a country’s multilateral exchange rate from purchasing power parity such as a real exchange rate index. At the very centre of theoretical explanations of this observation is the “Balassa-Samuelson hypothesis” (BS) featuring divergent productivity developments across sectors of tradable versus non-tradable goods. Accepting the validity of BS and assuming that productivity differences occur (almost) exclusively in the tradable goods sector immediately implies the cross-country relationship between comparative prices and real per capita income. In consequence, economic convergence based on productivity gains centred in the tradable goods sector implies inflation differentials and real exchange rate appreciation for catching up economies as equilibrium phenomena.

This consensus view bears significant consequences, as inflationary differentials with richer economies, as long as they are assessed to be grounded in convergence, will not set off economic policy responses. Economic development and stability are thus vitally dependent upon the correctness of this assessment.

The rest of the paper is organised as follows. Section 2 summarises the simple BS-framework based on arbitrage equalising prices – and thus unit labour costs – of tradables across countries, i.e., on incomplete trade specialisation. However, this specialisation argument remains implicit, as there is in fact no trade in this type of model. To check the robustness of the BS result, section 3 presents a simple model of trade with complete specialisation in differentiated tradables. This reveals a modified relationship between prices and productivities, the strength of which varies systematically with the strength of barriers to trade. Section 4 incorporates this insight into a more comprehensive arbitrage view on prices and productivities, extended by key real factors of relevance to transition economies, i.e., liberalisation, reallocation and restructuring.

## 2 Prices and productivities: The simple arbitrage view

In its simplest version, the Balassa-Samuelson hypothesis is based on a framework of two countries, two (homogenous) goods (one tradable,  $X$ , one non-tradable,  $Z$ ), and one factor of production (labour,  $L$ ). In each country  $j$ , production for both goods is linear in labour,

$$Z_j = A_j^N L(z)_j \text{ and } X_j = A_j^T L(x)_j.$$

Perfect competition among producers ensures that real wages equal labour productivity,

$$w_j / P_j^N = A_j^N \text{ and } w_j / P_j^T = A_j^T.$$

with  $w_j$  as the nominal wage rate in country  $j$ .

Assuming equal preferences across countries with constant expenditure shares  $\theta$  and  $1-\theta$ , respectively, for  $X$  and  $Z$ , demand is

$$X_j = \theta Y_j / P_j^T \text{ and } Z_j = (1-\theta) Y_j / P_j^N.$$

A consumer price index compatible with the underlying preferences is given by

$$P_j = (P_j^T)^\theta (P_j^N)^{1-\theta},$$

such that an internationally comparable consumer price index relation consistent with preferences is given by

$$\frac{P_1}{P_2} = \frac{(P_1^T)^\theta (P_1^N)^{1-\theta}}{(P_2^T)^\theta (P_2^N)^{1-\theta}}. \quad (1)$$

Arbitrage in tradables leads to price equalisation,  $P_1^T = P_2^T$ . Using the real wage relationships,

$$\begin{aligned} \frac{P_1}{P_2} &= \frac{(w_1 A_2^N)^{1-\theta}}{(w_2 A_1^N)^{1-\theta}} = \frac{(A_1^T P_1^T A_2^N)^{1-\theta}}{(A_2^T P_2^T A_1^N)^{1-\theta}} \\ &= \frac{(A_1^T A_2^N)^{1-\theta}}{(A_2^T A_1^N)^{1-\theta}}. \end{aligned} \quad (2)$$

The intuition for the link between prices and productivities in equation (2) is rather simple: wages correspond to the marginal value product of labour; with the latter rising in the tradables sector in country 1, e.g. due to technical progress, wages there will also increase, as tradables prices are tied by arbitrage. National labour mobility implies in-

creasing wages also in the non-tradables sector resulting in price increases there, absent any technical change, which correspond to the original increase in productivity in the country's tradables sector. In consequence, both the relative price of non-tradables in country 1 as well as country 1's overall price index relative to country 2 increase with increasing labour productivity in country 1's tradable sector, e.g., due to technical progress.<sup>10</sup>

Therefore, country 1's internationally comparable price index is higher than in country 2 if country 1's productivity in the tradables sector, relative to the non-tradables sector, is higher than in country 2. Assuming that productivity differences occur (almost) exclusively in the tradable goods sector immediately implies the cross-country relationship between comparative prices and real per capita income, cited in the introduction. In consequence, economic convergence based on productivity gains centred in the tradable goods sector implies inflation differentials and real exchange rate appreciation for catching up economies. However, this consensus view, illustrated in equation (2), can be criticised both on theoretical and empirical grounds:

- Even within the simple two-countries-two-goods-one-factor framework, different productivity increases across countries that are, however, balanced in tradables and non-tradables sectors do not imply cross-country inflation differentials; in case productivity differences occur (almost) exclusively in the non-tradable goods sector, converging economies should have even lower inflation than richer ones.
- Deviations from the simplifying assumptions of the simple two-countries-two-goods-one-factor set-up show the knife-edge role of these assumptions for the result of a systematic relationship between prices and productivities (cf., e.g., Podkaminer, 2003).
- Intuitively, much of the systematic relationship between prices and productivities seems to hinge on the assumption of homogenous tradables, with arbitrage equating international prices. Therefore, the following section uses a simple trade model to inquire the behaviour of prices and productivities when tradable goods are differentiated.

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<sup>10</sup> The only alternative, opened up by equation (2), for  $P_1/P_2$  to increase is by a rise in the share of non-traded goods in consumption over time; judging by trade shares of GDP, this seems heavily at odds with empirical developments, however. In section 4 will make a new attempt at isolating demand influences on  $P_1/P_2$ .

### 3 A monopolistic competition model of trade with product differentiation

The trade model underlying this section is Ricardian in nature, enriched by a Chamberlinian approach to product differentiation (cf. Venables, 1987). Again, there are two countries, each with only one production factor (labour); both countries produce a non-tradable good and a finite number of industrial products, which are all tradable.

#### 3.1 Production

The production function for the non-tradable good is again linear in labour, the only input, as in section 2,  $Z_j = A_j^N L(z)_j$ . Perfect competition among producers ensures that real wages equal productivities in both countries' non-tradable sectors,  $P_j^N = w_j / A_j^N$ , with again  $w_j$  as the nominal wage rate in country  $j$ .

In the industrial sector, monopolistic competitors manufacture industrial products, each of which is produced with a linear technology subject to internal economies of scale; industrial technologies are identical within a country but may differ between countries according to the specification used in Fitzgerald (2003),

$$x_j = A_j^T (a l_j - b). \quad (3)$$

Total costs of the representative industrial producer in country  $j$  are given by wage costs,

$$K(x_j) = \frac{w_j}{a} \left[ \frac{x_j}{A_j^T} + b \right], \quad (4)$$

#### 3.2 Demand

According to the usual formulation of consumer preferences over varieties of differentiated products (e.g., Frensch, 2002), let

$$M_j = \left[ n_j x_{jj}^\beta + n_i x_{ij}^\beta \right]^{1/\beta}, \quad (5)$$

be an index of differentiated tradable  $n_j$  domestic and  $n_i$  foreign-produced industrial consumer goods, with the relevant price index,

$$P_j^T = \left[ n_j q_{jj}^{\frac{\beta}{\beta-1}} + n_i q_{ij}^{\frac{\beta}{\beta-1}} \right]^{\frac{\beta-1}{\beta}}, \quad (6)$$

where  $q_{ji}$  describes the price of an industrial good produced in country  $j$  sold in country  $i$ , and where the nominal exchange rate is normalised to one;  $\beta = 1 - 1/\sigma$  and  $0 < \beta < 1$ , with  $\sigma$  as the constant elasticity of substitution between any pair of industrial products, i.e.

$$\frac{\partial(X_{ij} / X_{jj})}{\partial(q_{ij} / q_{jj})} \frac{q_{ij} / q_{jj}}{X_{ij} / X_{jj}} = -\sigma .$$

Assuming equal preferences across countries with constant expenditure shares  $\theta$  and  $1 - \theta$ , respectively, for  $M$  and  $Z$ , aggregate demand for the index of industrial consumer goods in country  $j$  is

$$M_j = \theta Y_j / P_j^T , \quad (7)$$

with  $Y_j = w_j L_j$  as total income in country  $j$ . Aggregate consumer demand in country  $j$  for an industrial product from either country can then be derived as (see, e.g., Frensch, 2002 or Venables, 1987),

$$X_{ij} = q_{ij}^{1/(\beta-1)} (P_j^T)^{\beta/(1-\beta)} \theta Y_j . \quad (8)$$

### 3.3 Barriers to trade

The assumption of a constant elasticity of substitution,  $\sigma = 1/(1 - \beta)$ , between any pair of industrial products gives rise to the possibility of linking product differentiation to the existence of indirect barriers to trade. By our definition, the elasticity of substitution between any pair of home-produced industrial products is assumed to be the same as between a domestically produced and a foreign industrial product. However, there are reasons to assume that in reality this is not so. The enormous weight of non-tariff barriers to trade in industrial goods, in the form of different national technical standards, norms etc., as compared to tariffs, can be taken as evidence that empirically the elasticity of substitution between any pair of home-produced industrial products is in fact higher than between a home-produced and a foreign industrial product, where the latter increases towards the former in the course of a reduction in trade barriers. This implies an increasing “average” elasticity of substitution between any pair of industrial products with decreasing barriers to trade.

For reasons of analytical tractability, in what follows we have to uphold the assumption of a constant elasticity of substitution. However, in line with the reasoning above, we will identify decreasing *indirect* barriers to trade with an increasing elasticity of substitution,  $\sigma$ , between any pair of industrial products, and thus with a rising  $\beta$ .

### 3.4 International equilibrium

Short-run equilibrium is defined by goods and factor markets clearing in both countries as well as balanced trade; a long-run equilibrium meets the additional condition that the monopolistically competitive producers make zero profits; our interest will be only in long-run equilibria.

Each monopolistic competitor specialises on exactly one variant of industrial goods and the short-run equilibrium condition, marginal revenue = marginal cost, holds for each producer on each market. Assuming a large number of industrial producers, each one's own price elasticity on each market is equal to  $1/(1 - \beta)$ . His marginal revenue on each market is accordingly  $MR(x_{ji}) = \beta q_{ji}$ ; absent *direct* barriers to trade, and with the nominal exchange rate normalised to one, prices of domestically produced industrial products are equal at home and abroad, i.e.,

$$\begin{aligned} q_{11} = q_{12} = q_1 &= \frac{1}{\beta} \frac{w_1}{aA_1^T} \\ q_{22} = q_{21} = q_2 &= \frac{1}{\beta} \frac{w_2}{aA_2^T} \end{aligned} \quad (9)$$

Profits of the representative component producer in country  $j$  are thus

$\pi_j = q_j(x_{jj} + x_{ji}) - w_j l_j$ . The cost function (4) and (9) imply

$$\pi_j = (1 - \beta)q_j(X_{jj} + X_{ji}) - w_j b_j / a.$$

Total value of industrial production in country  $j$  is  $R_j = n_j q_j (X_{jj} + X_{ji})$ . With (9),

$$R_j = \frac{n_j}{(1 - \beta)} \left[ \pi_j + \frac{w_j b_j}{a} \right]. \quad (10)$$

Total (value of) demand for industrial products in country  $j$  is  $\theta Y_j = n_j q_j X_{jj} + n_i q_i X_{ij}$ ; trade balance requires  $n_j q_j X_{ji} = n_i q_i X_{ij}$ . Therefore,  $\theta Y_j = \theta w_j L_j = R_j$ .

In the long-run, profits are eliminated by free market access and  $\theta w_j L_j = R_j$  implies via (9),

$$n_j = (1 - \beta)\theta L_j \frac{a}{b}, \quad (11)$$

illustrating that with internal returns to scale the division of labour is limited by the extent of the market.



### 3.5 Prices and productivities with trade in differentiated goods

For deriving an internationally comparable consumer price index relation in the framework of (1), consider that the demand equation (8) implies

$$\frac{X_{21}}{X_{12}} = \frac{q_2^{1/(\beta-1)} (P_2^T)^{\beta/(\beta-1)} w_1 L_1}{q_1^{1/(\beta-1)} (P_1^T)^{\beta/(\beta-1)} w_2 L_2}$$

and consequently

$$\frac{n_2 q_2 X_{21}}{n_1 q_1 X_{12}} = \frac{q_2^{\beta/(\beta-1)} (P_2^T)^{\beta/(\beta-1)} w_1}{q_1^{\beta/(\beta-1)} (P_1^T)^{\beta/(\beta-1)} w_2}$$

Due to balanced trade,

$$\frac{q_2^{\beta/(\beta-1)} (P_2^T)^{\beta/(\beta-1)}}{q_1^{\beta/(\beta-1)} (P_1^T)^{\beta/(\beta-1)}} = \frac{w_2}{w_1}. \quad (12)$$

Absent direct trade barriers, and according to equation (6),  $P_1^T = P_2^T$ , such that (12) simplifies to

$$\frac{q_2^{\beta/(\beta-1)}}{q_1^{\beta/(\beta-1)}} = \frac{w_2}{w_1}.$$

Substituting for  $q_2/q_1$  according to (9), this can be rewritten as

$$\frac{w_2}{w_1} = \left( \frac{A_2^T}{A_1^T} \right)^\beta. \quad (13)$$

Substituting tradables' prices and productivities for the wage ratio,

$$\frac{P_2^N}{P_1^N} = \left( \frac{A_1^N}{A_2^N} \right) \left( \frac{A_2^T}{A_1^T} \right)^\beta.$$

With  $P_1^T = P_2^T$ , this can be extended to give an internationally comparable consumer price index relation in the framework of (1), with product differentiation in the absence of trade barriers,<sup>11</sup>

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<sup>11</sup> For an analogous result in a slightly different multi-country setting, see Fitzgerald (2003).

$$\frac{P_2}{P_1} = \frac{(P_2^T)^\theta (P_2^N)^{1-\theta}}{(P_1^T)^\theta (P_1^N)^{1-\theta}} = \left[ \left( \frac{A_1^N}{A_2^N} \right) \left( \frac{A_2^T}{A_1^T} \right)^\beta \right]^{1-\theta}. \quad (14)$$

Comparing equation (14) to equation (2), i.e. to the classic BS-hypothesis without product differentiation, reveals two important differences.

First, complete specialisation and the presence of market power due to product differentiation imply a terms-of-trade effect that is absent in the classic BS set-up with homogenous goods. In consequence, this terms-of-trade effect might weaken the BS-effect even to the extent of a negative relationship between aggregate productivity ratios and relative prices. To illustrate this, again consider the case of equal productivity in the tradables and the non-tradables sectors of each country such that, according to equation (14)  $P_2/P_1 = (A_2/A_1)^{(\beta-1)(1-\theta)}$ : as  $0 < \beta < 1$  and  $0 < \theta < 1$ , the richer country has the lower price level.

Now assume that in equation (14) productivities in both countries' non-tradable sectors are constant and equal. Then, the elasticity of relative prices with respect to the productivities ratio in the tradable sectors equals  $\beta(1-\theta)$ . I.e., the higher  $\beta$ , i.e., the higher the elasticity of substitution between any pair of industrial products, the smaller the terms-of-trade effect and the closer equation (14) comes to the original classic BS set-up with homogenous goods. Thus, the higher  $\beta$ , the stronger the BS-effect in the context of differentiated goods. Identifying decreasing indirect barriers to trade with an increasing  $\beta$ , we can conclude that the lower the barriers to trade, the stronger the BS-effect in the context of differentiated goods. Relating barriers to trade to trade liberalisation, we may finally conclude that the strength of the BS-effect in the context of differentiated goods increases with trade liberalisation.

Obviously, this result fits in quite well with the empirical observation that a strong, systematic relationship between prices and productivities, as postulated in BS, has been only a relatively recent one, while it was practically unobservable until the mid-forties of the twentieth century (Bergin et al., 2004) when a substantial decline in trade barriers to international trade started to set in.

## 4 Prices and productivities: An extended arbitrage view for transition economies

The theoretical *caveats* at the end of section 2 as well as the results of section 3 should caution one to accept the existence of a strong, BS-based systematic relationship between prices and productivities as some kind of law of nature. For further analysis, we return to the arbitrage view of the classic BS set-up, however, extending the framework to incorporate the result of the previous section and the effects of transition. In the spirit of Blanchard (1997), we define transition as (trade) liberalisation, (resource) reallocation and (corporate) restructuring, and investigate whether there are transition-specific reasons to strengthen or weaken the BS-effect.

Following the definition in equation (1) of an internationally comparable consumer price index relation,

$$\Pi = p_1 - p_2 = \ln P_1 - \ln P_2, \quad (15)$$

with small fonts indicating logarithmic values. Rather than differentiating only between tradables and non-tradables, now we assume economies to have two sectors, industry (I) and services (S), with products entering national price levels with potentially different weights such that,

$$p_j = \phi_1 p_j^I + (1 - \phi_1) p_j^S. \quad (16)$$

In order to derive conclusions on transition-specific versus general BS-effects on the development of internationally comparable consumer price indices, we make a few simplifying, but nevertheless well-grounded assumptions to modify the simple two-countries-two goods-one factor of production set-up of section 2.

(A1) While all services are non-tradable, only part of industrial goods are tradable due to the existence of barriers to trade, i.e.,  $p_j^I = \tau_j p_j^{I,T} + (1 - \tau_j) p_j^{I,NT}$ .

(A2) Prices are proportional to unit labour costs,  $p_j^k = \lambda^k + \omega_j - a_j^k$ , where  $k = S, I, T$ , and  $I, NT$ ;  $\omega$  is the wage rate and  $a$  is labour productivity (both in logarithmic values),  $a_j^T = a_j^{NT} = a_j$ .

(A3) Exposure to international trade increases the intensity of competition, i.e.,  $\lambda^S = \lambda^{I,NT} = \lambda^{NT} > \lambda^T = \lambda^{I,T}$ .

(A4) Country 1 product quality of tradables,  $\kappa$ , is defined relative to (higher) country 2 quality,  $0 < \kappa < 1$ . Relative country 1 quality increases with aggregate relative productivity,  $\ln \kappa = \gamma(a_1 - a_2)$ .

(A5) Purchasing power parity, as usually, does not hold for non-tradables; for tradables, PPP is restricted by quality differentials according to  $p_1^{I,T} = \ln \kappa + p_2^{I,T}$ .

Definitions (15) and (16), together with (A5), imply long-run internationally comparable consumer price indices according to,

$$\Pi = \gamma(a_1 - a_2) + (p_1 - p_1^{I,T}) - (p_2 - p_1^{I,T}), \quad (17)$$

where equation (16) implies that

$$p_j - p_j^I = (1 - \phi_j)(p_j^S - p_j^I).$$

Inserting from (A1) results in

$$\begin{aligned} p_j - p_j^{I,T} &= (1 - \tau_j)(p_j^{I,NT} - p_j^{I,T}) + (1 - \phi_j)(p_j^S - p_j^I). \\ &= (1 - \tau_j)(p_j^{I,NT} - p_j^{I,T}) + (1 - \phi_j)p_j^S - p_j^I + \phi_j p_j^I \end{aligned}$$

and again using (A1),

$$\begin{aligned} p_j - p_j^{I,T} &= (1 - \phi_j)p_j^S - p_j^{I,T} + \phi_j p_j^I. \\ &= (1 - \phi_j)p_j^S - p_j^{I,T} + \phi_j p_j^{I,T} + \phi_j(1 - \tau_j)(p_j^{I,NT} - p_j^{I,T}) \\ &= (1 - \phi_j)(p_j^S - p_j^{I,T}) + \phi_j(1 - \tau_j)(p_j^{I,NT} - p_j^{I,T}) \end{aligned}$$

Substituting for prices according to (A2) and collecting terms yields

$$p_j - p_j^{I,T} = (1 - \tau_j \phi_j)(\lambda^{NT} - \lambda^T) + (1 - \phi_j)(a_j^I - a_j^S),$$

implying

$$\Pi = \gamma(a_1 - a_2) + (1 - \phi_1)(a_1^I - a_1^S) - (1 - \phi_2)(a_2^I - a_2^S) + (\tau_2 \phi_2 - \tau_1 \phi_1)(\lambda^{NT} - \lambda^T) \quad (18)$$

After total differentiation and again collecting terms, we may finally decompose the rate of change of the domestic (i.e., country 1 relative to country 2) internationally comparable consumer price index into four separate effects (where a  $\Delta$  of a logarithmic value indicates a growth rate),

$$\Delta \Pi = \quad (19)$$

$$(1 - \phi_1)(\Delta a_1^I - \Delta a_1^S) - (1 - \phi_2)(\Delta a_2^I - \Delta a_2^S) \quad \text{(a) Balassa-Samuelson}$$

$$\begin{aligned}
 & + \gamma(\Delta a_1 - \Delta a_2) && \text{(b) quality improvement due to restructuring} \\
 & + \Delta\phi_2[(a_2^I - a_2^S) + \tau_2(\lambda^{NT} - \lambda^T)] - \Delta\phi_1[(a_1^I - a_1^S) + \tau_1(\lambda^{NT} - \lambda^T)] && \text{(c) sectoral reallocation} \\
 & + (\phi_2\Delta\tau_2 - \phi_1\Delta\tau_1)(\lambda^{NT} - \lambda^T). && \text{(d) trade liberalisation}
 \end{aligned}$$

Referring to (a) as “Balassa-Samuelson,” may appear a bit sloppy, as it is differential increases in productivities between industry and services that drives international price differentials here, rather than between tradables and non-tradables. However, we want to refer to the general notion of sectoral productivity differentials as BS effect, which, as usual, also in this framework implies an increasing domestic (i.e., country 1 relative to country 2) internationally comparable consumer price index (e.g., a real appreciation in case prices are made comparable by market exchange rates). Compared to section 2, the slight twist in sectoral decomposition, by adding economic activity categories to the tradable-non-tradable dichotomy, is quite fruitful: it allows us to show that reallocation from domestic industry towards services ( $\Delta\phi_1 < 0$ ) *cet. par.* also implies an increasing domestic relative price index, as long as productivity in domestic industry is higher than in domestic services.

Remember from the argument in section 2, that in the simple set-up there, the only alternative to BS for  $P_1/P_2$  to increase was by a rise in the share of non-traded goods in GDP, which seems heavily at odds with empirical developments. The argument here, however, allows to separate tradability from income shares spent on services and industrial goods. In fact, referring to the experience of transition economies, it seems well documented that both effects, sectoral productivity differentials and an increasing share of services in income, were relevant forces especially during the early stages of transition,<sup>12</sup> while for later stages of transition as well as for the experience of non-transition economies in general this seems less certain.

In the much richer context here, equation (19) also illustrates that in addition to differential productivity growth and reallocation, corporate restructuring during transition and convergence and the ensuing quality improvements in domestically produced tradable industrial goods also imply an increasing domestic relative price index.<sup>13</sup>

The argument so far leaves us only trade liberalisation as a potential transition-specific force to weaken real appreciation. As equation (19) shows, a unilateral reduction in domestic versus foreign barriers to trade in industrial products ( $\Delta\tau_1 > 0$  and  $\Delta\tau_2 = 0$ ) implies a decreasing domestic relative price index. Symmetric reduction in barriers to trade ( $\Delta\tau_1 = \Delta\tau_2 > 0$ ) implies a decreasing domestic relative price index as long as the share of the services sector in total production is larger abroad than domestically, which is certainly relevant in early stage of transition.

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<sup>12</sup> For a documentation of the former effect, see Halpern and Wyplosz (2001), for the latter, cf. Frensch (2000).

<sup>13</sup> For a theoretical motivation, see Frensch (2004).

However, the results of section 3 illustrated that trade liberalisation, in the sense of reducing trade barriers, has a positive effect on the strength of the BS-effect in the context of differentiated goods. While it is not possible to explicitly build this result into the arbitrage decomposition analysis of this section, we can still take this result to considerably weaken the argument stated above on trade liberalisation as a potential transition-specific force against real appreciation.

Thus taking the discussion of section 3 into consideration, the extended arbitrage view of this section illustrates that in fact there are transition-specific reasons to strengthen the BS-effect such that there is a specific inflation differential between transition economies and non-transition economies. Also, due to the nature of transition specificities analysed in this section, we may conclude that this transition-specific inflation differential is higher in earlier than in later stages of transition.

## 5 Some conclusions

This paper illustrates and connects two aspects of the relationship between internationally comparable prices and productivities, based on the BS effect: in the context of trade in differentiated goods, a higher elasticity of substitution is shown to strengthen the BS-effect. Identifying decreasing indirect barriers to trade with an increasing elasticity of substitution, we conclude that the strength of the BS-effect in the context of differentiated goods increases with trade liberalisation. Considering this result in an extended arbitrage view on prices and productivities, we show that there are in fact no transition-specific forces to weaken the BS-effect, when defining transition as (trade) liberalisation, (resource) reallocation and (corporate) restructuring.

While this implies that transition-specific inflation differentials are higher in earlier than in later stages of transition, an immediate policy-relevant corollary of this is that inflation differentials between Central and East European economies and the euro-area should weaken over time with transition effects fading out. To the extent that inflation differentials persist, they increasingly signal disequilibria, rather than equilibrium phenomena connected with economic transition and convergence, calling for stabilisation efforts.

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