DEVALUATION AND OUTPUT IN FIVE TRANSITION ECONOMIES: A PANEL COINTEGRATION APPROACH OF POLAND, HUNGARY, CZECH REPUBLIC, SLOVAKIA AND ROMANIA, 1993-2000

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Abstract

This paper assesses the impact of devaluations on aggregate output for a group of five transition economies during the period 1993-2000. An application of panel unit root tests and panel cointegration establishes the presence of a long run relationship between real output, real exchange rates, real money and real wages, while the estimation of the long run relationship reveals that devaluations are contractionary in the long run. This finding is in contrast with a large part of the literature, which discern no long run effect on output.

Keywords: Exchange Rates, Devaluation, Output, Panel Cointegration, Transition Economies, EU Accession

JEL Classification: F31

1. Introduction

Devaluations are an important element of economic adjustment and stabilization programs and are frequently used to improve a country's balance of payments position, boost domestic employment, and accumulate more foreign exchange reserves. However, while there is consensus that devaluation is a useful instrument for balance of payments adjustment, wide controversy surrounds the issue of how devaluations impact aggregate output.

The output reaction to devaluations and depreciations becomes all the more important for transition economies aspiring to join the European Union. These EU accession/candidate countries, among other objectives, are laboring to boost output so as to accelerate the process of economic convergence. In the second half of the 1990s, while transition economies showed more than a healthy growth performance, real output convergence resurfaced as an important issue in the policy and theoretical literature. Gács (2003) points out that during the 1988-1999 period the relative position of most Central and East European Countries (CEECs) vis-à-vis the EU worsened, thus there was no convergence. He shows that the per-capita GDP of 10 CEECs[1] as a percentage of the EU 15 average declined from 53% in 1988 to 38.8% in 1999. Halpern and Wyplosz (1997) note that, in most transition economies, liberalization was followed by sharp real exchange rate

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Acknowledgement: The author thanks Peter Pedroni for providing his panel cointegration code, and gratefully acknowledges Taggert J. Brooks and I-Ming Chiu for their helpful suggestions.

Research for this article was supported by a generous grant from the Horace H. Rackham School of Graduate Studies at the University of Michigan

[1] This group includes Bulgaria, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia, Slovenia.

depreciation and a subsequent appreciation. Did this subsequent appreciation hurt output through lost competitiveness?

Large exchange rate movements in transition economies have prompted several empirical assessments but have not put an end to the controversy surrounding their effect on real output. To mention some of the more recent studies on the topic, Mitchell and Pentecost (2001) find devaluations contractionary in a panel study of Bulgaria, Czech Republic, Poland and Slovenia in the short-run as well as the long run. The long run contractionary effect is somewhat mitigated by a rise in output one year after the devaluation. In contrast, Karadeloglou et al. (2001), using a wage-price-GDP model, find devaluations to be slightly expansionary in Slovenia, only initially expansionary, but with no long run effect, in Bulgaria, and contractionary in Poland. Bahmani-Oskooee (1998) investigates 23 LDCs in a time series cointegration framework and validates the hypothesis of neutral devaluations with respect to output in the long run. Chou and Chao (2001) employ panel unit root tests in a bivariate framework and conclude that devaluation hurt Asian economies' output in the immediate aftermath of the 1997 crisis, however it left no prints in the long-run.

Most authors that have resorted to panel type regression techniques have done so without regard to the stationarity of the series involved running the risk of obtaining spurious regression estimates. On the other hand, that part of the literature that employs cointegration analysis on time series data may have been undermined by the low power of the tests applied to very short time series. The issue of the short time span is even more critical in studies of transition economies since many studies are only interested in the post 1992 period.

This paper analyzes the impact of devaluations/depreciations on output for five transition economies: the Czech Republic, Hungary, Poland, Romania, and the Slovak Republic[2]. In contrast to the literature this paper intends to improve upon the traditional approaches of the existing econometric literature by introducing cross-section variation and employing panel unit root tests and panel cointegration under multiple regressors to test for stationarity and presence of cointegration in a panel setting[3]. Section 2 outlines the general theoretical framework and model in its reduced form equation. The econometric methodology is explained in section 3. Section 4 reports model estimation followed by a summary in section 5.

2. The Theoretical Framework

The "orthodox" school advocates the argument that devaluation is expansionary because of its expenditure switching effects and the increased production of tradables that it stimulates. But exports of transitional economies may not be as responsive to devaluations since their products are not of the same quality as those of industrial economies. In

[2] Except for Romania, which is expected to join the EU in 2007, all are EU members since May 2004.

[3] In contrast to Chou and Chao (2001) that employ panel cointegration in a bivariate framework, this paper uses Pedroni's panel cointegration for heterogeneous panels in a *multivariate* framework. This latter not only allows for heterogeneity among panel members, but also for the inclusion of more than one right-hand-side variable.

addition, devaluations can cause output to contract because of other factors. First, devaluation can cause a contraction of aggregate demand because, among other things, it redistributes income towards economic entities with high marginal propensity to save (Krugman and Taylor, 1978), it makes capital investment more expensive (Branson, 1986), and increases debt and debt service payments in local currency (Cooper, 1971). Second, devaluations may also reduce aggregate supply as the price of imported production inputs increases (Bruno, 1979), wages increase when based on price levels (Hanson, 1983), and working capital grows costlier as real balances decline (Bruno, 1979).

This study uses a testable reduced form equation for output based on a macroeconomic model with IS-LM and aggregate supply equations derived by Mills and Pentecost (2001):

$$y_t = a_0 + a_1 m_t + a_2 q_t + a_3 w_t + e_t \qquad (1)$$

where m is the real money supply, q the real exchange rate and w the real product wage. The rationale for incorporating the real wage rate is that in transition economies wage income occupies a very large share of total income. Therefore increased real wages lead to higher real incomes, which in turn yield a greater demand for domestic output. As indicated above, the signs of coefficients a2 and a3 are ambiguous and necessitate an econometric approach to estimate the net effect of real exchange rates and real wages on output. Indeed, as the real exchange rate appreciates output is affected in two ways.

First, it may suffer from a decrease in net exports, as the Marshall-Lerner condition would suggest. Second, it causes a lower price for consumer goods because of cheaper imports, which in turn boosts the real wage. On the one hand, a higher real wage triggers a reduction in output supplied, but stimulates aggregate demand through increased consumption. Naturally, the net effect of these counteracting channels has to be measured empirically.

3. Methodology and Estimation Procedure

A panel framework is chosen to estimate the effects of devaluation on output mainly because it can control for heterogeneity in individual behavior. It offers more variation, less collinearity among regressors, and more efficient estimators. Moreover panel models intrinsically present less measurement error problems as well as a mitigated omitted variable bias.

While using a panel framework has obvious benefits, one has to consider whether the panel members present unacceptable heterogeneity. Indeed, Mills and Pentecost (2001, p.430), note that "it is unwise to generalize about the effects of devaluation on output for a set of economies as diverse as the transition economies of Eastern Europe." While this statement undermines the case for using panel data techniques, Gács (2003) observes that, despite their obvious differences, CEECs seem to be a less heterogeneous block than the EU economies. To this effect, he notes that there was remarkable similarity across the economies of Central and Eastern Europe in terms of the dominating heavy industry, prioritizing investment in the utilization of income, and the distinct patterns of trade impressed by membership in the Council of Mutual Economic Assistance (CMEA). During the 1990's the CEECs have struggled to break away from these structural straight-jacket-like similarities of the 1980's. In spite of everything, they have performed similar structural reforms like the emancipation of services, the move away from agriculture and toward more 'progressive' industries[4], as well as the increased reliance on foreign savings to finance domestic investments. Therefore, the use of panel data is justified to a great extent by the structural similarities of these economies.

The explanatory variables chosen to explain variation in real output are real effective exchange rates (REER), real money (M) and real wages (W). Inevitably, the estimation of this reduced form equation entails the regression of nonstationary variables such as output, and could potentially produce spurious results. According to Granger and Newbold (1974), the usual t and F tests have a tendency to reject the hypothesis of no relationship between these variables even when there is none. As a matter of fact, regressing two independent random walks will almost invariably result in a significant relationship. The literature of contractionary devaluations has for the most part used least squares estimation techniques on levels from pooled cross-section and time series data. Although in a panel setting, these studies can still suffer form the so-called spurious regression problem, which necessitates the use of panel cointegration analysis.

As indicated by Engle-Granger, a variable is considered integrated of order d if it becomes stationary after being differenced d times. A set of variables, integrated of order d, can be considered cointegrated if the residuals from the regression of one variable on the others are integrated of order less than d. Since conventional cointegration tests are designed to examine the existence of long run relationships in time series data, this study resorts to a more recently developed test of cointegration in panel data. Panel unit root tests and panel cointegration tests have been developed on the same principles that underlie the conventional ADF test. The integrating property of each panel variable is first examined by means of employing one of several panel unit root tests. Their most prized feature perhaps is the degree of homogeneity that they allow. For example, a test by Levin and Lin (1992) allows for heterogeneity of the intercepts across members of the panel, a more recent test by Im, Pesaran, and Shin (1997) (IPS test hereafter) allows for heterogeneity in intercepts as well as in the slope coefficients. The Im, Pesaran, and Shin test is based on the equation below:

$$\Delta y_{it} = m_i + b_i y_{i,t-1} + \sum_{k=1}^{p_i} q_{i,k} \Delta y_{i,t-k} + g_i t + e_{it} \tag{2}$$

where i = 1, 2, …,N and t = 1,2,…,T.

The null hypothesis is $\beta_i = 0$, for all i's, while the alternative hypothesis is $\beta_i < 0$. The IPS statistic is an average of the individual ADF statistics computed as follows:

$$\bar{t} = \frac{1}{N} \sum_{i=1}^{N} \frac{\hat{b}_i}{\hat{s}_{\hat{b}_i}} \tag{3}$$

In a further step, the above t-bar statistic is standardized so that it converges to a standard normal distribution, as N grows large.

[4] Industries that are prevalent in more advanced industrialized countries.

Additional adjustments become necessary when the test is applied to the residuals of a reduced form model like equation (1) with multiple regressors. Unlike the rest of the panel cointegration tests developed to date, Pedroni (1995, 1997) has constructed a framework that allows testing for cointegration of homogeneous and heterogeneous panels with multiple regressors. Following Pedroni (1999), consider the following model:

$$y_{it} = \alpha_i + \beta_i t + \gamma_{1i} x_{1i,t} + \gamma_{2i} x_{2i,t} + \ldots + \gamma_{Mi} x_{Mi,t} + e_{i,t} \tag{4}$$

for i = 1, 2, …,N cross-sections; t = 1,2,…,T observations; and m = 1,2,…,M regressors. In the above equation, α_i represents the fixed effect or the individual-specific effect that is allowed to vary across individuals. The slope coefficients γ_{mi} and the time effect β_i are modeled heterogeneously as well.

The two statistics developed by Pedroni, which this study uses, differ in that the first is considered to be a within-dimension statistic or panel t-statistic, while the second is a between-dimension statistic or group t-statistic. Their labels are based on the way the autoregressive coefficients are manipulated to arrive at the final statistic. The panel-t statistic is constructed from estimators that pool the autoregressive coefficient across different individuals for unit root tests on the estimated residuals. The Group-t is built on estimators that merely average the individually estimated coefficients for each i. Hence, while Panel-t statistic virtually averages the numerator and denominator terms of the individual t-statistics separately, the Group-t statistic averages the entire ratio of individual t-statistics. Even though the null hypothesis is the same for both tests, the alternative hypothesis is different. In the case of Panel-t statistic, the alternative hypothesis assumes that the stationary autoregressive parameter is homogenous, unlike the alternative hypothesis of the Group-t statistic, which allows for a heterogeneous stationary autoregressive parameter. Pedroni (1995, 1997) uses the moments of the underlying Brownian motion functions that describe the individual statistics under the null hypothesis to normalize the distributions of these test statistics. The null of no cointegration is then tested based on these standard normal statistics. Under the alternative hypothesis, these two statistics diverge to negative infinity. Hence a large left tail value implies a rejection of the null hypothesis. More details about the critical values or the approximate standardization can be found in Pedroni (1999).

The presence of cointegration, as detected by the panel cointegration test, would signal that there is a long-run equilibrium between these variables, which in turn can be estimated using a least square dummy variable model or a random-effects GLS regression. Such a model would yield estimates of the long run effects of devaluations on output.

4. The Results

Naturally, even in a panel framework, the investigation of a long run relationship begins with stationarity tests for all the variables involved. The quarterly dataset includes five transition economies[5]: Czech Republic, Hungary, Poland, Romania, and Slovak Republic,

[5] This selection was primarily dictated by quarterly data availability. Since quarterly GDP data are difficult to find, data on quarterly industrial production was used instead.

spanning the period from the first quarter of 1993 to the third quarter of 2000. Panel unit root tests with heterogeneous lag truncation that allow for heterogeneous trends were applied. Table 1 displays the results of the panel unit root tests as given by the Group-t statistic as well as the individual ADF statistics for each country over time.

Table 1. Panel Unit Root Test Results

	Czech R.		Hungary		Poland		Romania		Slovak R.		
	ADF	Lags	ADF	Lags	ADF	Lags	ADF	Lags	ADF	Lags	t-Stat
Y	-2.50	4	0.30	1	-1.80	0	-3.48	4	-0.49	4	1.74
M	-2.02	4	-3.33	4	-1.74	4	-3.77	4	-2.57	4	-1.51
REER	-3.13	1	-2.48	0	-2.14	4	-2.35	1	-4.96	3	-2.49
W	-2.33	4	-0.51	4	-1.15	3	-3.06	4	0.33	3	2.50

Note: The Group-t statistic presented is an adjusted test result that can be compared to the N(0,1) distribution. Because the test is left tail-sided, the 1% critical value is –1.96, the 5% critical value is –1.64, and the 10% critical value is –1.28.

All the variables but the real effective exchange rate have a Group-t statistic that is greater than the critical value of –1.96 from the standard t-table, indicating that the null of non-stationarity cannot be rejected. Two more panel unit root tests were conducted on the real effective exchange rate to ascertain its stationarity properties. The Group-t statistics of a second test that does not assume heterogeneous trends and a third test that includes time dummies in addition to heterogeneous trends were respectively 0.23 and –1.95. While the results of these additional tests make a borderline case for the nonstationary properties of real effective exchange rates, it is safe to assume they are nonstationarity based also on findings of previous research. Properties of real exchange rates in transition economies have been examined by other studies like Barlow (2004) that indicates the purchasing power parity does not hold between accession economies and developed market economies.

As a next step, panel cointegration tests for the presence of long-run relationships among our variables are conducted based on the following log-linear model:

$$Y_{it} = \alpha_i + \beta_i t + \gamma_{1i} REER_{it} + \gamma_{2i} M_{it} + \gamma_{3i} W_{it} + e_{it} \qquad (5)$$

For both, the panel t-statistic and the group t-statistic developed by Pedroni (1995, 1997), two results are presented: one that pertains to the standard case that allows for variation only in country-specific fixed effects (ai), and another more general specification where the slope coefficients, gmi, and the time effect, bi, are modeled heterogeneously just like the intercept terms. Table 2 reports the results of the panel cointegration tests, where both statistics presented are standard normal and will reject the null of no cointegration if they are large negative numbers (smaller than –1.96). Regardless of the presumption of the standard case or the heterogeneous deterministic trends, the above test statistics strongly reject the null of no cointegration. Hence the null of non-stationary residuals in equation (5) is rejected, which implies that real output, real effective exchange rates, real money and real wages are cointegrated. Therefore, these tests reveal that the stochastic trends of these variables cancel each other out in the long run yielding a stable equilibrium relationship.

Table 2. Panel Cointegration Tests

Standard Case:	
panel t-statistic	-3.87094*
group t-statistic	-12.38318*
Heterogeneous Deterministic Trends:	
panel t-statistic	-11.27322*
group t-statistic	-40.47072*

Notes: [a] The Panel-t and Group-t statistics presented are adjusted test results according to Pedroni's procedure that can be compared to the N(0,1) distribution. [b] The asterisk implies that the null of no cointegration can be rejected at the 1% level. Because the test is left tail-sided, the 1% critical value is –1.96, the 5% critical value is –1.64, and the 10% critical value is –1.28.

Having established the presence of cointegration, the estimation of the long run relationship becomes feasible. Applying the principles of the Engle and Granger (1987) methodology on pooled data, when the variables are cointegrated, any OLS-based estimates of the cointegrating vector are consistent. The estimates of the cointegrating equation (5) from the LSDV and random-effects GLS models are presented in Table 3.

It is evident from Table 3 that in both models the real exchange rate carries a significant and positive coefficient. Its positive sign indicates that devaluations or depreciations have a contractionary impact on real GDP (an increase in the real effective exchange rate index is synonymous with appreciation). Moreover, the real exchange rate coefficient[6] is relatively sizeable suggesting that 1% devaluation would lead to a 0.68% reduction in real output in the long run. This effect is comparable to that estimated by Mitchell and Pentecost (2001) who use a panel data set on four transition countries.

Table 3. Estimates of the Cointegration Equation

	LSDV Model				Random-Effects GLS Model			
Variable	Coeff.	Std.Err.	T-Stat.	P-val.	Coeff.	Std.Err.	T-Stat.	P-val.
CONST	0.658	0.471	1.397	0.164	1.957	0.404	4.846	0.000
REER	0.680	0.167	4.076	0.000	0.473	0.171	2.761	0.006
M	0.443	0.094	4.689	0.000	0.042	0.041	1.024	0.306
W	-0.816	0.133	-6.112	0.000	-0.503	0.123	-4.090	0.000

5. Conclusion

The objective of this work is to add to the existing empirical literature on the effect of devaluation on aggregate output. It examines the issue of contractionary devaluations for a group of five eastern European countries during the period 1993-2000. Since the existing theoretical literature recognizes that devaluations have the potential to become

[6] The Hausman specification test favors the use of the fixed effect model (LSDV) versus random-effects.

contractionary, this paper makes an attempt to estimate a reduced form equation for output with real effective exchange rates, real money, and real wage rates as explanatory variables.

The application of recent techniques in panel unit root tests and panel cointegration, that avoid spurious regression results and offer increased power, establishes the presence of a long run relationship between these four variables. The estimation of this long run linear relationship lends support to the contractionary devaluation hypothesis and stands in contrast with a large part of the literature which holds that devaluations do not affect output in the long run.

The contractionary effect of devaluations in these transition economies may be part of the rationale behind a long-standing reluctance China has shown in devaluing its own currency, instead of pursuing export tax rebates to stimulate its external sector. The substantial output effects of devaluation in these transition economies may similarly induce some policies of exchange rate rigidity in European Union potential candidate countries like Albania, Bosnia and Herzegovina, Former Yugoslav Republic of Macedonia, and Serbia and Montenegro in their efforts to join the EU.

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Apendix A: Data

Data were predominantly extracted from the International Financial Statistics (IFS) of the International Monetary Fund, in CD-ROM format. IFS Is the International Monetary Fund's principal statistical publication.

The dataset contains quarterly information for Czech Republic, Hungary, Poland, Romania and Slovak Republic and spans the period from 1993 Q1 to 2000 Q3, allowing 31 observations on each country. Since quarterly data on real GDP for these countries is largely unavailable, the industrial production series from line 66 of the IFS is used.

Real Effective Exchange Rates (REER) were obtained from the section of Exchange Rates and Exchange Rate Arrangements. Money (M) represents a broad measure of money comparable to what is commonly referred to as the M2 monetary aggregate. This variable was extracted from subject codes 34 and 35. Money wage rates were obtained from line 65 of the IFS.