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Testing for price convergence: how close are EU New Member's States to euro zone?

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Abstract

The purpose of my paper is to empirically explore the validity of the relative version of purchasing power parity (PPP) for nine EU New Member States (NMS) using unit root and panel unit root frameworks. Three periods are considered: the period before NMS joined the EU (1995:01-2004:04), the period after NMS joined the EU (2004:05-2008:12), and the sum of the two periods (1995-2008). We find evidence that real exchange rates follow generally unit-root processes implying that PPP doesn't hold even as a long-run relationship. In addition, results show that price co-movements increase strongly after NMS joined the EU.

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1. Introduction

This paper examines the price convergence through the PPP concept for nine EU new member's states (NMS). After joining the EU as a result of their progress during a still unfinished transition and convergence process, the NMS have as objective to join euro zone for some reasons: this entry could guarantee them lower interest rate, a certain monetary stability and a more easily access to foreign founds. As highlighted in the Maastricht Treaty, the accession to Euroland requires the achievement of five criteria including inflation convergence and nominal exchange rate stability within its member states. The nominal exchange rates and inflation convergence became consequently a crucial key issue for the future integration of NMS into euro zone.

The most common way to investigate the empirical validity of purchasing power parity (PPP) doctrine consists in searching unit roots in real exchange rates. If the unit root can be rejected in favour of level stationarity, then deviations from parity are temporary and PPP is said to hold in long run. Given its importance in international finance, many studies have analysed long-run PPP relationship during the last decade using significant developments in econometrics[1]. Despite them, the slow speed of convergence documented in international markets remains a puzzle[2]. Possible explanations of these conflicting results are related to the insufficiently longer time span to overcome the possible problem of 'small sample distortions', the low power of unit root tests, the degree of cross correlation and heterogeneity of the series in the panel, appropriate price indices and the potential structural breaks because data could cover both the fixed and floating exchange rate regimes.

Given these diverging results, this paper investigates how European economic integration has affected the price convergence process between NMS and euro zone using new econometric developments. More precisely, the aim of the study is to empirically test the hypothesis on price convergence in the NMS's case employing the concept of beta and sigma convergence. Three periods are considered: the whole period (1995-2008), the sub-period before NMS joined the EU (1995:01-2004:04) and the sub-period after NMS joined the EU (2004:05-2008:12). Results show evidence that real exchange rates follow generally unit-root processes and additionally, that price's co-movements are stronger during the post-accession period of NMS into EU.

The rest of the paper is structured as follows. Section 2 describes the methodological aspects concerning the price convergence process. Section 3 presents the results obtained using conventional time series procedures and panel data unit root tests in the context of a set of somewhat homogeneous developing countries. Concluding remarks are reported in section 4.

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[1] Such as conventional time series, econometric techniques as well as recent panel unit root tests (Levin and Lin, 1992, 2002; Im, Pesaran and Shin, 1997), heterogenous panel unit root tests with cross-section dependence (Bai and Ng, 2001; Pesaran, 2003, Moon and Perron, 2004) and panel data cointegration tests (Baltagi and Kao, 2000; Banerjee, 1999; Pedroni, 2001b). See also Hurlin and Mignon (2005) for a review of panel data unit root tests.

[2] While a number of recent papers claim that PPP does hold in the long run [Drine and Rault (2005)], other studies [Asplund and Friberg (2001)] highlight that the LOOP does not even hold for identical goods sold at the same location, as long as these goods are denominated in different currencies (the evidence comes from price data of products sold in Scandinavian duty-free stores). See also Obstfeld and Rogoff, (2000), Taylor (2001, 2002), Parsley and Wei (1996), Cecchetti et al (2002) etc.

2. Methodological Discussions

The strict form of PPP says that the nominal exchange rate is proportional to the relative price ratio so that the real exchange rate is constant over time. Let's take s_t the nominal exchange rate, and p and p^* *are*, respectively domestic and foreign prices, all measured in natural logarithms. If the real exchange rate defined as: $q_t = s_t - p_t + p_t^*$ follows a stationary process then deviations from parity are temporary and will disappear with time, and PPP is said to hold in the long run. Contrary, if q_t contains a unit root, then it does not have an unconditional mean, and when a deviation of the real exchange rate from its equilibrium occurs, this deviation will not be eliminated over time. In this case, PPP is said to fail in the long run. A large body of literature on empirical exchange-rate studies has reported evidence that real exchange rates follow unit-root processes, implying that PPP does not hold even as a long-run relationship. This section presents the time series process for real exchange rates to investigate this assumption.

2.1 Testing for β-convergence

The analysis about convergence is traditionally based on Solow's (1956) neo-classical growth model, which predicts that a poor economy tends to grow faster than a rich one. In this model, the assumption of diminishing returns is fundamental for convergence hypothesis to hold because economic agents will allocate resource (i.e. labor and capital) across different locations to maximize their wealth. As a result, differences in returns to labor or to capital among different countries will diminish over time. Based on Solow's model, a huge "convergence literature" has been developed, typified by the seminal papers by Barro and Sala-i-Martin (1992) and Mankiw et al. (1992), exploring two measures of convergence: the sigma (σ) and the beta (β)-convergence in its conditional and unconditional version. This section briefly discusses the methodological aspects referring to (β)-convergence.

The literature distinguishes two *β*-convergence versions: the absolute and relative β-convergence. Absolute (unconditional) beta convergence occurs when economies with different levels of a given variable approach an identical point, while conditional (weak) beta convergence can be identified when economies with different levels of a given variable approach each other but, the identical point is not reached. To measure beta convergence or the capacity of developing countries to growth, we estimate the so-called Barro's regression, in which we replace income levels with price differentials according to the following equation:

$$\Delta p_{ij,t} = \alpha_{ij} + \beta p_{ij,\,t-1} + \varepsilon'_{ij,t} \qquad (1)$$

where $p_{ij,t} = \left| \ln P_{it} - \ln P_{jt} \right|$ is the measure of international price dispersion between country i and j and is computed for all pairs of countries. As noted above, the dependent variable of equation (1) is the first difference in the log-price in country *i* relative to euro zone. The key parameter is β and it captures the speed of convergence.

To investigate the empirical validity of PPP concept, we rewrite the relation (1) in order to estimate the following AR(1) (panel) regression:

$$p_{ij,t} = \alpha_{ij} + \beta^* p_{ij,\,t-1} + \varepsilon_{ij,t} \qquad (2)$$

Our interest is in the parameters of the relation (2): α_{ij} - a constant (which captures the countries' steady state), β* which refers to the rate or speed of convergence computed as λ = -

$\ln(\beta^*)$ and $\varepsilon_{ij,t}$ which is a disturbance term. One of the advantages of this technique is that it takes into account not only the cross-sectional dimension but also the time dimension, hence providing greater degrees of freedom.

Under the null of no convergence, β^* is equal to zero. In this case, a shock to $p_{ij,t}$ is permanent and the PPP does not hold in the long run. Convergence implies a negative β^*, with the approximate half-life of a shock to $p_{ij,t}$ given by $t^* = -\ln(2)/\ln(1+\beta^*)$. The half-life indicates the number of periods needed to halve the distance from the mean. When α_{ij}-a country pair specific effect equals zero, the absolute LOOP is tested. Under an alternative assumption, when the individual effect is included ($\alpha ij \neq 0$), the conditional version of LOOP is calculated. In the last case, the individual effect permits to each price difference to converge to a single, country specific, mean.

To estimate the speed of convergence, we firstly check for unit roots in each series and in the panel. For this purpose, we apply unit root tests (Augmented Dickey-Fuller test) and panel unit root tests based on a more conservative works of Levin and Lin (1993), Taylor and Sarno (1998), Maddala and Wu (1999), Im, Pesaran and Shin (2003) and on a relatively recent econometric development of Pesaran (2003). The main reason of this choice is that the univariate unit-root tests, initiated by Dickey and Fuller, have notoriously low power and it becomes difficult to reject the unit root null when it is in fact incorrect. Because of the weakness of the univariate unit-root test, researchers exploited the panel dimension of data available in certain applications.

If these tests and their different versions, with respect to the form of deterministic components (with and without individual effects, time effects and a time trend) show that the null hypothesis of unit root is rejected in all cases at standard levels of confidence, it can be possible to investigate the speed of convergence. Now we describe briefly the main features of the panel unit root tests considered.

Levin and Lin (1993) is a first generation panel unit root model and assumes that each individual unit in the panel shares the same AR(1) coefficient, but allows for individual effects, time effects and possibly a time trend. This test may be viewed as a pooled Dickey-Fuller test, or an Augmented Dickey-Fuller (ADF) test when lags are included, with the null hypothesis of non-stationarity (I(1) behavior). After transformation, the t-star statistic is distributed standard normal under the null hypothesis of non-stationarity.

Taylor and Sarno (1998) proposed a Multivariate ADF test, where the sum of the autoregressive coefficients may vary across countries under the alternative hypothesis. The test applies Zellner's seemingly unrelated equation estimator to N equations, defined for the N units of the panel. Each equation is specified as a k-th order auto-regression. This test involves testing the hypothesis, for each equation, that the sum of coefficients of the autoregressive polynomial is equal to unity. The null-hypothesis implies that this condition is satisfied over the N equations. Under the null-hypothesis, all of the series under consideration are I(1), or non-stationary, stochastic processes. A rejection of the null hypothesis should be considered carefully because this result doesn't indicate that each of the series is stationary.

Maddala and Wu (1999) test, which is referred as the Fisher test, combines the p-values from N independent unit root tests and assumes that all series are non-stationary under the null hypothesis against the alternative that at least one series in the panel is stationary. ZFisher follows a chi-square distributions with N*2 degrees of freedom. It is worth noting that due to the pooling of p-values from independent unit root tests, ZFisher could be applied to unbalanced panels.

Im, Pesaran and Shin (2003) is another first generation model that we apply in our study. It allows for individual effects, time trends, and common time effects for heterogeneous panels. Based on the mean of the individual Dickey-Fuller t-statistics of each unit in the panel, the IPS test assumes that all series are non-stationary under the null hypothesis. The number of lags of the dependent variable may be introduced to allow for serial correlation in the errors and the W[t-bar] statistic is distributed standard normal under the null hypothesis of non-stationarity. The test proposed by Im, Pesaran and Shin permits to solve Levin's and Lin's serial correlation problem in assuming heterogeneity between units in a dynamic panel framework.

Pesaran (2003) is a second-generation model and is applied to check for unit roots in heterogeneous panels with cross-section dependence. This method assumes that inter-individual correlations of the variable considered are not only due to an inter-individual correlation of the residuals, but come from the presence of one or more common components. Parallel to Im, Pesaran and Shin (IPS, 2003) test, it is based on the mean of individual DF (or ADF) t-statistics of each unit in the panel. Null hypothesis assumes that all series are non-stationary. To eliminate the cross dependence, the standard DF (or ADF) regressions are augmented with the cross section averages of lagged levels and first-differences of the individual series (CADF statistics).

2.2 Testing for price co-movements

In this section, we explore other features of the price movements in an attempt to gain additional perspective into how European economic integration has affected the price convergence process in the case of NMS's partners. Here, our interest is to explore the co-movements of prices between nine NMS and euro zone in order to fit some possible cyclical features of the evolution of these economies.

We follow the methodology developed by Alesina, Barro and Tenreyro (2002) and Shin and Lee (2004) to examine the price co-movements between each NMS and euro zone. We compute the relative prices P_{it}/P_{jt} where *i* correspond to each NMS and *j* to euro zone. For each pair of countries (i, j), we use monthly time series to compute the first order auto-regression:

$$\ln P_{it}/P_{jt} = \alpha_{ij} + \beta^* \ln P_{i,t-1}/P_{j,t-1} + \varepsilon_{ij,t}. \qquad (1)$$

The estimated residual $\hat{e}_{t,i,j}$ measures the relative prices that cannot be predictable from the first prior value of relative price. The co-movement of prices VP_{ij} is calculated using the root-mean-squared error:

$$VP_{ij} = \sqrt{(1/T\text{-}2)\sum(\hat{e}_{ij,t})^2} \quad \text{where } T = \text{the number of observations} \qquad (2)$$

The lower VP_{ij}, the greater the co-movement of price between *i* and *j*. In other words, if price series are affected by similar shocks, the conditional variance of the ratio of these series should be near to zero. A positive response of the co-movements to the monetary union would lead to a higher degree of consensus around a common monetary policy without involving imbalances between participants and to lower costs generated by the loss of the monetary autonomy. A negative response of the co-movements would have the opposite effect: it would generate divergent fluctuations prejudging that the exchange rate policies will be important to support the economic activity.

Here, $|\ln P_{it}/P_{jt}|$ represents a measure of international price dispersion between country i and j and is computed for all individual pairs of NMS and euro zone. The concept of price dispersion is derived from the literature of real convergence (Barro and Sala-i-Martin, 1992) and originally concerned cross-sectional dispersion of income to measure sigma convergence. Sigma convergence occurs when the dispersion of the levels of a given variable between different economies tends to decrease over time. In our case, sigma convergence occurs if the dispersion of prices declines over time and is usually measured using the standard deviation.

3. Results

3.1 Data

To perform unit root tests in order to verify whether or not the theory is empirically valid we employ monthly data on nominal exchange rates and consumer price index from nine NMS (the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia, Slovenia and Cyprus) covering the period from 1995 to 2008. The data are taken from Eurostat. All series are transformed into natural logarithms and consumer price variables are in index form with 2005 as the base year. Note that the real exchange rate (quoted to incertain) is calculated as the ratio of the consumer price indices (CPI) and that the euro is taken as a benchmark.

3.2 Empirical Results: β – convergence

In order to determine the presence of a unit root in individual country specific data we employ standard ADF test. For a panel unit root we conduct Levin-Lin (1992), Taylor and Sarno (1998), Maddala and Wu (1999), Im, Pesaran and Shin (2003) and Pesaran (2003). The panel tests include a constant and a heterogeneous time trend in their specifications. Note that by estimating bilateral regressions, we give up the country dimension corresponding to a panel data set - we only exploit the time variation in price differences. In light of this, the standard ADF test show that the unit root null could not be rejected when we consider the full sample, the sub-sample ex-ante (i.e. the sub-period *before* NMS joined the EU) and the one allowing the homogenous exchange rate regime. In these cases, the series are generated by an I(1) process and hence, the PPP doesn't hold in the long run and the relative prices will diverge.

While previous estimations on price convergence failed to reject the null of a unit root in the price series when time series data were used, the estimations on the *ex-post* sub-sample show that the null hypothesis of unit root is rejected in some country pair cases. This finding on relative price convergence implies that it is possible for them to estimate the speed of convergence β. Table 1 reports the convergence coefficients (standard errors in parentheses) obtained by estimating convergence with standard unit root method. It is worth noting that the β coefficients were obtained using euro zone as the base country. The striking feature of this table is that the estimated speeds of convergence are very high for selected countries. The estimated autoregressive coefficients imply a very high speed of price convergence of 58,7%, 52,4% and 48,2% per month for Cyprus, Czech Republic and respectively, Slovenia and consequently, a half life a price shocks of 1,2 month, 1,3 month and 1,4 month.

Table 1: Results for Country-Pairs

	Chypre	Czech Rep.	Slovenia
β coefficients	-0,4133* (0.103)	-0,4759* (0.115)	-0,5176* (0.166)
speed of convergence (λ)	0,587	0,524	0,482
half-life of price shocks (t*)	1,182	1,323	1,437

Note: *- significantly at 1% level.

Because results of the univariate unit root tests lead to a mitigated diagnosis, we will apply panel unit root tests mentioned above that combine the information from time series with the information from cross-sectional units. The main advantage of the addition of cross-sectional variation to time series variation is that both dimensions improve estimation efficiency leading to smaller standard errors and to higher t-ratios.

Table 2 reports the results of panel unit root tests for the full sample and both sub-samples: the ex-ante period (1995:01-2004:04) and the ex-post-period (2004:05-2008:12).

We begin with the presentation of the results of the Levin and Lin procedure, which is the traditional panel unit root test with the hypothesis that all the cross-section units are non-stationary against the hypothesis that all are stationary. The evidence from our simulations suggests that the null hypothesis of unit root is rejected only for the model without trend (for the whole period and the ex-ante period) in which case the PPP hold in the long run. The null hypothesis of unit root is accepted for both models (with and without a time trend) for the ex-post period. We have to note that the rejection of the null hypothesis of unit root means that at least one from the series of our panel is stationary.

To detect the presence of a unit root in our panel, we also studied the Multivariate ADF (Mad-Fuller) statistic, typified by Taylor and Sarno (1998). Table 2 presents the Mad-fuller statistics and the critical values at 5% level derived by Taylor and Sarno from Monte Carlo simulation. We conclude that at least one from the series is stationary in all cases considered. Results based on the Fisher test of Maddala and Wu (1999) only fail to accept the null hypothesis of unit roots for the model without trend in the ex-ante period.

For the full sample and the two sub-samples, Im, Pesaran and Shin (2003) tests allow us to accept the unit-root null in a vast majority of the cases. That is to say, regardless of the sample (sub-samples) period, there is very little evidence of a stationarity in our panel.

Results of the Pesaran (2003) procedure illustrate that the null hypothesis of unit root is rejected only for the model without trend, in which case the PPP hold in the long run.

To summarize our results, regardless of the econometric method, we strongly reject the null hypothesis that real exchange rates between NMS and euro zone in our sample, contain a unit root only in the case of the model without trend as highlighted by Levin and Lin (1992), Taylor and Sarno (1998) and Pesaran (2003) procedures.

In this model, the explanatory variable is the lagged value of the accepted measure of price dispersion, so it can be classified as dynamic and consequently, the standard panel data estimators— Pooled OLS levels, Within Groups (Fixed Effects)—cannot be used. Nevertheless, we report the results of the estimated autoregressive parameter (even if it is biased upwards or downwards) to compare them with those obtained by 2SLS and GLS.

Table 2: Panel Unit Root Results

	Tests	Period	Statistics	Model without trend	Model with trend	Lags
First generation models	Levin and Lin (1992)	Whole	t_ρ^*	-2.86 (0.002)	-0.34 (0.37)	1
		Ex-ante	t_ρ^*	-1.35 (0.09)	-1.12 (0.13)	1
		Ex-post	t_ρ^*	0.46 (0.66)	0.20 (0.58)	1
				MADF_stat	Approx 5% CV	
	Taylor and Sarno (1998).	Whole	MADF	34.97**	15.19	1
		Ex-ante	MADF	56.47**	17.06	1
		Ex-post	MADF	33.23**	20.54	1
	Maddala and Wu (1999)	Whole	P_{MW}	14.19 (0.72)	21.37 (0.26)	1
		Ex-ante	P_{MW}	30.41 (0.03)	23.62 (0.17)	1
		Ex-post	P_{MW}	6.71 (0.99)	13.17 (0.78)	1
	Im, Pesaran and Shin (2003)	Whole	Z_{tbar}	-1.80 (0.19)	-2.30 (0.33)	1
			W_{tbar}	-0.88 (0.19)	-0.45 (0.33)	
		Ex-ante	Z_{tbar}	-1.77 (0.22)	-2.06 (0.67)	1
			W_{tbar}	-0.77 (0.22)	0.45 (0.67)	
		Ex-post	Z_{tbar}	-1.45 (0.59)	-2.00 (0.74)	1
			W_{tbar}	0.23 (0.59)	0.64 (0.74)	
Second generation models	Pesaran (2003)	Whole	t_ρ^*	-2.33 (0.06)	-2.53 (0.29)	1
			Z [t_bar]	-1.58 (0.06)	-0.54 (0.29)	
		Ex-ante	t_ρ^*	-2.59 (0.009)	-2.73 (0.12)	1
			Z [t_bar]	-2.36 (0.009)	-1.18 (0.12)	
		Ex-post	t_ρ^*	-2.26 (0.09)	-2.53 (0.29)	1
			Z [t_bar]	-1.37 (0.09)	-0.56 (0.29)	

Note: - the p-value associated with different statistics is in parenthesis; ** - means significant at 5% level.
- the null hypothesis is a unit root process (no convergence).

Table 3 illustrate the half-lives of a shock to $p_{ij,t}$ and the speed of convergence. For the 2SLS model, the estimated autoregressive parameter of 0.449 implies an average speed of price convergence of 8.01% per month and therefore a half-life of 3.5 months.

Table 3: The estimations of beta convergence

Method	OLS	FE within	2SLS	GLS with AR(1)
β*	0,449 (0.097)	0,201 (0.21)	0,449 (0.097)	0,398 (0.036)
Speed of convergence	*0,801*	*1,606*	*0,801*	0,921
Half-life (t*)	*3,491*	1,144	*3,491*	*8,783*
Observations	1549	1549	1549	1239

Note: - the p-value associated with different statistics is in parenthesis

3.3 Empirical Results: price co-movements

Table 5 provides the results of the price's co-movements for selected countries with the euro zone. It is worth noting that a higher number (i.e., more variance) means less co-movement and then higher costs to join the monetary union. We aim to emphasize how European economic integration has affected the magnitude of co-movements among participating countries. For that, we compute the co-movements for the full period and three sub-samples: the sub-period *before* NMS joined the EU, the sub-period *after* NMS joined the EU and the one allowing for homogenous exchange rate regimes. We have also considered three groups of countries according to their exchange rate systems: an intermediary exchange rate regime (group 1: Cyprus and Hungary), a fixed exchange regime (group 2: Estonia, Latvia and Lithuania) and a floating exchange rate regime (group 3: Czech Rep., Poland, Slovakia and Slovenia). Table 4 presents this classification and the periods selected for estimations.

Table 4: Exchange rate regimes and sub-periods selected by countries

Countries	Period		Exchange rate regimes
	Start	Fine	
Chyprus	2001:01	2008:12	peg to euro with +/-15% margins
Estonia	1996:01	2008:12	currency board (euro)
Hungary	2001:06	2008:12	peg to euro with +/-15% margins
Latvia	1996:01	2008:12	peg on DTS
Lithuania	2002:02	2008:12	currency board (euro)
Poland	2000:01	2008:12	free floating/inflation targeting
Czech Rep,	1999:01	2008:12	managed floating/inflation targeting
Slovakia	1999:01	2008:12	managed floating
Slovenia	1996:01	2004:01	narrowly managed floating

Table 5 shows our measures of co-movements according with these four criteria. Results firstly suggest that the co-movements of prices in *ex-post* conditions are higher than those in *ex-ante* conditions. The co-movements on the whole period are generally lower than those obtained on the sub-periods, results which are in line with our main hypothesis that successful integration should translate to lower price differentials. We find also that the co-movements of prices according with homogenous exchange rate regimes are higher than those on the whole period but generally lower than those on the sub-periods.

Table 5: Co-movements of prices with euro zone

Countries	Co-movements homogenous exchange rate regime	Co-movements ex-ante	Co-movements whole period	Co-movements ex-post
Group 1				
Chyprus	0,0034	0,0058	0,0062	0,0033
Hungary	0,0140	0,0117	0,0159	0,0118
Group 2				
Estonia	0,0047	0,0036	0,0047	0,0020
Latvia	0,0125	0,0110	0,0125	0,0042
Lithuania	0,0067	0,0139	0,0149	0,0017
Group 3				
Poland	0,0177	0,0174	0,0214	0,0114
Czech Rep,	0,0122	0,0122	0,0147	0,0065
Slovakia	0,0106	0,0138	0,0156	0,0051
Slovenia	0,0039	0,0034	0,0043	0,0017

Note: the table illustrates the value of VP, the standard error of the residual for the AR(1) regression with Augmented Dickey-Fuller test for the log of the real exchange rate.

4. Concluding remarks

In this study we employ standard and panel unit root methods for evaluating the purchasing power parity doctrine and prices co-movements for nine NMS. The analysis is mainly realized for three periods: the full period (1995-2008), the ex-ante sub-period (1995-2004) and the ex-post sub-period (2004-2008). The empirical findings of this paper do not support the relative version of PPP in the most part of cases. The analysis of country pair separately furthermore indicates that this failure of the PPP is not driven by the data from only a few countries: Cyprus, Czech Rep. and Slovenia allowing the ex-post period. Results provided by the panel unit root find evidence for the PPP only in the case of the model without trend. In addition, the co-movements of prices are generally high and differ between sub-periods; results suggesting that co-movements are higher in the *ex-post* period than those in *ex-ante* period. The homogenous exchange rate regimes seem to increase the co-movements of prices compared to the full period selected.

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