Financial Development and Growth: Panel Cointegration Evidence from South-Eastern and Central Europe

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Abstract

Ever since Schumpeter, macroeconomists have argued that financial development has a large and direct effect on the long run wealth of a nation. In this paper, we empirically investigate this relationship for a panel of 16 South-Eastern and Central European countries over the period 1995-2014 by employing a state-of-the-art panel cointegration technique. We find that financial development has a positive effect on the income per capita. The effect is statistically robust to other estimation methods and is economically large since it is almost twice the size of the gross capital formation. Nevertheless, the panel cointegration tests indicate a possibility of an endogenous relationship between the phenomena.

Keywords: panel-cointegration, financial development, economic growth
JEL Classification Codes: C23, C51, O11, O47, O52, E50

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1. Introduction

For Schumpeter [35], growth was related to innovation of products and continual improvements in the existing ones. Well-functioning banks accelerate this process of innovation by identifying and funding the entrepreneurs with the best chances of successfully implementing innovative products or production processes. The magnitude of acceleration depends on the financial development of the economy. The financial development is expressed in the composition of a country's policies, factors and institutions and reflects the structures that lead to efficient intermediation and effective financial markets. Thus, it should directly promote stable, long term economic growth.

There is a large part of economic theory discussing this relationship [18, 21]. Although conclusions must be stated carefully, the preponderance of theoretical reasoning and empirical evidence suggest that Schumpeter was right. In other words, most papers conclude that there is a positive, first-order relationship between financial development and economic growth. The growing body of literature would convince even the biggest skeptic that development of financial markets and institutions is a critical and inextricable part of the growth process, leaving behind the belief that the financial system is an inconsequential shadow, responding passively to economic growth and real sector needs. However, to our knowledge, the scholars from South-Eastern and Central Europe have left this subject almost untouched. We believe that this due to the lack of country specific data; i.e. the span of the time series is short.

Motivated by all these advancements in the topic, in this paper we aim to discover the long run relationship between the financial development, measured as the M2 to nominal GDP ratio, and the income per capita (ergo, long run growth) in South-Eastern and Central Europe. We do this by employing a panel cointegration technique, since we believe that there is a possibility of an endogenous relationship between the variables. Particularly, the endogeneity could arise because of an interdependence between financial development and income per capita and would definitely bias our estimates. In that aspect, the advantage of using panel cointegration estimators over others is that they are robust under cointegration to a variety of estimation problems that often plague empirical work, including endogeneity [24].

The remains of the paper proceeds as follows: Section 2 provides a short review of the growing literature on this subject; Section 3 presents the empirical model and the data; Section 4 is consisted of the estimated model and the results from the model; and finally Section 5 concludes and gives suggestions for future research in this area.

2. Literature Review

The effect of financial development as a driver of economic growth, was indirectly introduced by Schumpeter [35]. However, the first documented positive correlation between growth and indicators of financial development was recorded by Goldsmith [8]. Since the documentation, a debate on the issue of whether financial development plays a critical role in determining long run economic growth rates was risen. A unique resolution for this debate still cannot be found. Yet, resolving this debate would definitely provide guidance in distinguishing among theoretical models, but even more importantly, the information for the importance of finance for growth would affect the intensity with which researchers and policy makers attempt to identify and construct appropriate financial sector reforms around the world.

Following [8], King and Levine [18] conclude that better financial systems improve the probability of successful innovation and consequently accelerate economic growth. On the other side, the authors also find that distortions in the financial sector reduce the rate of innovation and therefore reduce the growth rate. Furthermore, Levine and Zingales [32] discover that industrial sectors grow faster in countries with relatively better developed financial markets than in countries in less developed financial markets.

Out of the panel studies, we can point out Beck and Levine [2] in which the authors argue that stock markets and banks positively influence economic growth; and then again Levine [23] where the author concludes that both financial intermediaries and markets matter for growth. In addition, based on a panel data for a set of 4 Latin American countries, Bittencourt [3] again confirms the Schumpeterian view. What is new in his paper is the highlighted importance of macroeconomic stability as an essential precondition for financial development.

Cross-country analyses document extensive periods when financial development or the lack of it crucially affects the speed and pattern of economic development. For example, Gregorio and Guidotti [7], find a positive effect of financial development over long-run economic growth, measured as real GDP per capita. The positive effect is particularly strong in middle and low-income countries, since large extent of the financial development in high-income countries occurs outside the banking system.

Conversely to all previous findings, there are empirical researches who find little or no evidence of a positive correlation between financial development and growth. For instance, Shan and Morris [36] examine 19 OECD countries and China, and barely find an evidence that financial development precedes economic growth, either directly or indirectly; casting a doubt on claims that financial development is a necessary and maybe sufficient precursor of economic growth. Another paper done by Boulila and Trabelsi [4] on the issue of causality in the Middle East and North Africa presents little support to the view that financial development is a leading factor in determination of long-run growth for the countries of this region at least. We believe these findings are due to country specifics, the organization of the financial system itself, and other factors relevant to the sampled countries.

Closely related to our region, Yucel [39] finds bidirectional causality between financial development, trade openness and growth in a country-specific study about Turkey for a period of 18 years, concluding that economic policies aimed at financial development and trade openness have a statistically significant impact on economic growth. The latest empirical examination of this subject, to our knowledge, targets 8 countries of Central and Eastern Europe. In it, Dudian and Popa [6] prove empirically positive relationship between financial development and economic growth by using panel data for the period of 1996 - 2011.

In this paper, we go a few steps further by widening the time spread from 15 to 19 years and enlarging the sample of countries from 8 to 16. More importantly, differently from most previous studies, we employ the panel cointegration technique, for which we argue that is the correct estimation procedure when investigating the long run effect of financial development over growth.

3. Empirical model and data

In this section, we adopt an empirical specification that captures the long run relationship between a set of three variables: income per capita, financial development and gross capital formation; and describe the data.

The dependent variable, income, is measured as real GDP per capita corrected for Power Purchasing Parity. Its changes usually represent the economic growth of a country and they can be explained with various factors. However, our main goal is to explain it with an indicator of financial development. For that purpose, we follow the existing literature [1, 18, 19, 26, 37, 41] and use the ratio of the broad measure of the monetary stock M2 to the level of nominal GDP as our measure of financial development. Using this simple monetized variable has two advantages: (i) data for it is very easily obtainable; and (ii) it best reflects the savings function [17]

Nevertheless, financial development alone is not enough to explain economic growth, as it fails to explain various effects. Therefore, we include the Gross Capital Formation as a percent of nominal GDP. This indicator represents a simplification of the investments in the country which have been extensively utilized [11] as a crude approximation for a number of factors that can affect both financial development and economic growth by evolving smoothly over time. On the long run this ratio should promote technology indirectly and increase the wealth of a nation [20].

3.1 Empirical specification and econometric issues

Given the variables, we try to find their long run relationship with the help of panel cointegration technique. Many of the endogenous growth proponents, such as Romer [34], suppose that an economy grows exponentially. We accept their opinion and assume that our basic empirical model is given by:

$$log(gdp_{ct}) = \alpha_c + \beta_1 log(fd_{ct}) + \beta_2 log(gcf_{ct}) + u_{ct} \quad (1)$$

where i = 1, 2, ..., C and t = 1, 2, ..., T are country and time notations, fd_{ct} stands for the log of M2 as a percent of nominal GDP and $log(gcf_{ct})$ is the logarithm of gross capital Formation as a percent of GDP. The level of economic development is represented by real GDP per capita, gdp_{ct}, measured in logs. The β coefficients in equation (1) capture the long run effects between the variables, while α_c are country specific fixed effects that help controlling any omitted factors that are stable over time.

Equation (1) assumes a long run trivariate relationship between financial development, investments and the level of GDP per capita. For this assumption to hold, it is necessary that the individual time series for each of three variables (M2, Gross Capital Formation and per capita income) are nonstationary, integrated of the same order and that $log(fd_{ct})$, $log(gcf_{ct})$ and $log(gdp_{ct})$ form a cointegrated system [10].

By definition, two or more non-stationary variables are cointegrated if there exists a linear combination of these variables that is stationary. Therefore, cointegration in the traditional sense, indicates that the long-run relationship between the variables is linear[1]. Moreover, it implies that a regression consisting of cointegrated variables has a stationary error term, hence, no relevant integrated variables are omitted. Any omitted non-stationary variable that is part of the cointegrating relationship would enter the error term u_{it}, thereby producing non-stationary residuals and failure to detect cointegration.

On the other hand, if there is cointegration between a set of variables, then the same stationary relationship exists also in an extended variable space (see, e.g., Johansen [14]); if the variables are nonstationary and not cointegrated, the error term is nonstationary as well, and equation (1) would in this case represent a spurious regression in the sense of Granger and Newbold [9]. Our basic model (1) has three variables, and therefore, the existence of one cointegrating relationship implies that there are two permanent shocks, or common trends, and a transitory shock (Stock and Watson [38]). A number of factors and mechanisms could be the driving forces behind permanent and temporary shocks. Potential permanent shocks could be advancements in financial services or technology trends, while changes in the foreign exchange policy could be treated as transitory shocks.

[1] In our case the relationship is log-linear.

3.2 Data and descriptive statistics

For the purpose of examining the effect of financial development over growth in South-Eastern and Central Europe we collect annual data from the World Development Indicators Database (http://databank.worldbank.org/) for 16 countries from that region: Albania, Belarus, Bulgaria, Bosnia and Herzegovina, Czech Republic, Estonia, Croatia, Hungary, Latvia, Macedonia, Moldova, Poland, Romania, Slovakia, Slovenia and Ukraine.

We focus on the period from 1995, when most of the sampled countries started reporting the data, until 2014, when was the last time they reported it. Thus, we end up with an unbalanced panel of 314 observations. The panel is unbalanced because in some years some country data was missing.

Table 1 gives the country and total sample summary statistics. They reveal that the countries of South-Eastern and Central Europe are characterized with low-to-medium financial development, as well they are part of the low-to-medium income group of countries. Between the cross sections, Albania[2] has the highest average M2 to GDP ratio, followed by Czech Republic and Slovakia, while Romania and Belarus have the lowest average ratio. Average per capita GDP is highest in Slovenia, followed by Czech Republic, Hungary, Estonia and Slovakia. Moldova is the poorest country in the sample. The summary statistics suggest that, overall, the M2 to GDP ratio, the gross capital formation to GDP and the income per capita have grown constantly through the years, so we expect a positive relationship between them.

Table 1: Summary Statistics

Country	gdp	fd	gcf	Country	gdp	fd	gcf
ALB	6979.27	0.70	0.25	LVA	15365.12	0.35	0.27
BLR	11036.88	0.23	0.30	MDA	3256.80	0.38	0.26
BGR	12241.08	0.57	0.22	MKD	9706.24	0.36	0.23
BIH	7678.70	0.43	0.24	POL	17079.10	0.45	0.21
CZE	24527.18	0.64	0.30	ROU	14163.90	0.33	0.24
EST	19843.14	0.43	0.30	SVK	19704.82	0.59	0.28
HRV	17995.36	0.56	0.24	SVN	25148.19	0.49	0.26
HUN	20012.43	0.52	0.24	UKR	6653.86	0.38	0.21

	gdp	fd	gcf
Sample mean	14497.92	0.46	0.25
Sample standard deviation	7239.00	0.18	0.06
Sample maximum	30822.97	0.85	0.42
Sample minimum	2276.10	0.11	0.00

[2] The fact that Albania has the highest M2 to nominal GDP ratio seems a bit counter intuitive. Nevertheless, we do not treat Albania as an outlier and keep it in the sample.

4. Empirical analysis

4.1 Stationarity tests

Prior to conducting cointegration tests, all variables should have same time series properties. Particularly, they should have a unit root in levels and be integrated of the same order - $I(d)$.

Unit root examination is done with two tests: Im, Pesaran and Shin (IPS) [12] and Maddala and Wu (MW) [25]. The tests use a modification of the augmented Dickey-Fuller (ADF) regression:

$$\Delta y_{it} = w_i y_{it-1} + \sum_{L=1}^{k_i} \delta_{iL} \Delta y_{it-L} + \varphi_i z_{it} + \varepsilon_{it} \tag{2}$$

where k_i is the lag length, z_{it} is a vector of deterministic terms, explaining the fixed effects or the individual trends, and φ_i is the corresponding vector of coefficients[3]. The w_i coefficients are substitutions for $\rho - 1$. Under the null hypothesis the time series are non-stationary, while the alternative assumes the opposite. The hypotheses may be written as:

H_0 : $w_i = 0$; for all i

H_a : $w_i < 0$; for at least one i

Both tests represent the second generation of panel stationarity tests as they relax the assumption that the first order autoregressive parameter must be same across countries [22]. IPS test the hypotheses with the standardized t-bar statistic described in (3).

$$\bar{t}_{IPS} = \frac{\sqrt{N}[\frac{1}{N}\sum_{i=1}^{N} t_i - \frac{1}{N}\sum_{i=1}^{N} E(t_i|\rho_i = 0)]}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} var\,(t_i|\rho_i = 0)}} \Longrightarrow N(0,1) \tag{3}$$

Their test takes the average of the individual t_i Dickey Fuller statistics across sections and standardizes it with the expected mean and variance. However, Maddala and Wu [25] find that their test is superior to IPS. Because of that we also calculate the MW ADF Fisher type test which is the sum of the logs of the p-values of each individual cross section unit root test. The test statistic is shown in (4).

$$P = -2\sum_{i=1}^{N} \ln(p_i) \Longrightarrow X_{2N}^2 \tag{4}$$

[3] Throughout the explanation of the methodology i is used to denote a particular cross-section and N the total number of cross-sections. In our specific case, $c = i$ and $C = N$.

Table 2 reports the panel unit root tests. Every test concludes that the variables are non-stationary in levels and integrated of order one. On the one hand, the conclusions for $\log(gcf_{ct})$ and $\log(fd_{ct})$ should be treated with caution as they are not pure unit root processes (their values are bounded between 0 and 1). Yet, on the other hand, Jones [15] states that a variable may act as a unit root process within its boundaries. In fact, similar investment ratios as the log of the gross capital formation were already used in Pedroni [31] and Herzer and Vormer [11].

Table 2: Panel Unit Root Test statistics

	IPS		MW	
Variable	level	difference	level	Difference
$\log(gdp_{ct})$	-0.33	-4.05***	34.47	68.23***
$\log(fd_{ct})$	0.01	-4.65***	24.80	73.28***
$\log(gcf_{ct})$	0.59	-4.78***	21.49	74.60***

Notes: Two lags were selected to adjust for autocorrelation. Individual intercepts were included in every test.
*** Indicates significance at 1% level.

4.2 Cointegration tests

Also prior to estimating the long run model, a cointegration relationship between the variables needs to be confirmed. To inspect this property we use two types of tests, the panel accommodated Johansen-Fisher test and Pedroni's Engle-Granger based tests.
Maddala and Wu [25], with the help of Fisher (1932), adjusted the Johansen [13] test to panel data.

$$\Delta y_{it} = \Pi_i y_{it-1} + \sum_{j=1}^{k} \Gamma_{ij} \Delta y_{it-j} + \varphi_i z_{it} + \varepsilon_{it} \tag{5}$$

In (5) y_{it} is a $p \times 1$ vector of endogenous variables (in our case $y_{it} = [\log(gdp_{ct}), fd_{ct}, \log(gcf_{ct})]'$); p is the number of variables and Π_i represents the long run $p \times p$ matrix. If $1 < rank(\Pi_i) < p$, the matrix can be written as $\alpha_i \beta_i'$, where β_i' is a $r \times p$ matrix which rows are the cointegrating vectors, while α_i is a $p \times r$ matrix that gives the amount of each cointegrating vector entering the error correction model.

The Johansen-Fisher test statistic is computed in a similar way as in (4), just now it is summed over the p-values of the cross sectional trace or maximum eigenvalue cointegration tests. The difference between those two tests is the formulation of the hypotheses. The trace test is a one sided test with an alternative of more than r cointegrating vectors, whereas the maximum eigenvalue performs separate tests on each eigenvalue with an alternative hypothesis of exactly r + 1 cointegration vectors[4]. The advantage of these tests is that they do not specify the cointegrating vectors. Instead they search for how many stationary combinations can be made with the set of variables. Therefore, if we conclude that there are one or two cointegrating vectors, there is still the problem of deciding which ones are they.

To solve this problem we additionally calculate Pedroni's [27] within-dimension and between-dimension ADF and PP test statistics. The tests are four of the seven statistics proposed by the author. We estimate only these because Pedroni [30] concludes that in samples with small time dimension, such as ours, they have the best properties.

The tests' estimation method is an extension of the Engle and Granger's methodology where first for each cross-section, the dependent variable is regressed on the explanatory variables, $y_{it} = \alpha_{it} + \gamma_i t + \beta X_{it} + \epsilon_{it}$. After that, the stationarity of $\hat{\epsilon}_{it}$ is examined with either a technique similar to the Dickey-Fuller tests or to the correction terms in the single equation Phillips-Perron tests. The difference between the dimensions is that the within-dimension has a homogenous alternative, $p_i = p < 1$ for all i whilst the between-dimension has a heterogeneous alternative hypothesis, $p_i < 1$. In other words, the calculated Pedroni test statistics represent a group of four different, yet similar asymptotically normally distributed tests which diverge to negative infinity.

The results of the panel cointegration tests are given in Table 3. Clearly, the failure of both, the trace and the maximum eigenvalue test, to reject the null hypothesis of less than two cointegrating vectors at any level, means that there is a cointegration relationship between the variables. Nevertheless, the need of a less than 5% significance level to infer that there is only one cointegration vector in both tests may not be enough. Therefore, it must be taken into account that the relationship specified in equation (1) could possibly not be the only long run relationship between the variables.

Yet, three of the four Pedroni tests conclude that, at any level, the relation we defined is a long run relationship between the logs of income per capita, financial development and Gross Capital Formation. Only the ADF between-dimension test needs a 5% significance level to infer the same.

4.3 Long run relationship

Since the pre-tests for unit-roots and cointegration suggest that the variables are non-stationary and cointegrated as assumed in equation (1), we proceed to estimation of the long run relationship using the dynamic ordinary least squares (DOLS) within-dimension (pooled) estimator suggested by Kao and Chiang [16]. We opt for this estimator since it yields unbiased and asymptotically efficient estimates of the long run relationship, even if there are endogenous regressors, thus allowing us to control for the potential endogeneity of financial development

[4] Recall, if the Johansen-Fisher tests concludes that $r = p$, then the inspected time series are stationary in levels. Hence, no cointegration.

Table 3: Panel Cointegration Tests

Johansen-Fisher Panel Cointegration Test			
	r = 0	r ≤ 1	r ≤ 2
Trace statistic	143.80***	52.37**	24.62
Max-Eigen statistic	128.80***	52.20**	24.62
Pedroni Panel Cointegration Tests			
	Within-Dimension	Between-Dimension	
ADF statistics	-3.71***	-2.10**	
PP statistics	-3.61***	-3.72***	

Notes: The Johansen-Fisher test is has X^2 distribution with 2N degrees of freedom. Pedroni's test has to be adjusted with terms derived in [27] and then it is asymptotically normally distributed.
***(**) Indicates rejection of the null hypothesis at 1%(5%).

and per capita income [10]. Additionally, it has been established that in panel data samples with small time dimension the DOLS estimator performs better than other available estimators[5], like, for instance, the non-parametric fully modified ordinary least squares (FMOLS) estimator developed by Pedroni [28]. The DOLS model, given in (6), is a modification of equation (1).

$$log(gdp_{ct}) = \alpha_c + \beta_1 log(fd_t) + \beta_2 log(gcf_{it}) + \sum_{j=-p}^{q} \varphi_{1cj} \Delta \log(fd_{ct+j}) + \sum_{j=-p}^{q} \varphi_{2cj} \Delta \log(gcf_{c,t+j}) + \epsilon_{ct} \quad (6)$$

In the equation φ_{1cj} and φ_{2cj} represent coefficients of lead (q) and lag (p) differences which help generate unbiased estimates of β_1 and β_2 by eliminating asymptotic endogeneity and serial correlation.

The within-dimension DOLS estimates for the coefficients on the Gross Capital Formation rate and the M2 to GDP ratio are reported in column 1 of Table 4. As expected, both $log(fd_{ct})$ and $log(gcf_{ct})$ are positive and highly significant. The long run elasticity of income per capita to our measure of financial development is 0.55, is implying that, ceteris paribus, an increase of 1% in the M2 to nominal GDP ratio, on average increases the real GDP per capita by 0.55%. Similarly, if the rate of the Gross Capital Formation to GDP increases by 1%, an economy's income will grow by 0.38%.

Although, estimated this way, the coefficients measure income per capita's long run elasticity with respect to financial development and gross capital formation, a better comparison of their magnitude could be made by standardizing their values. Therefore, we make a standardization by multiplying them with the standard deviation ratio of the independent and dependent variables[6]. In the long run, one standard deviation increase in

[5] More about the performance of DOLS and other panel cointegration estimators can be read in [40].

[6] $\beta_m \times \sigma_{Xct}/\sigma_{\log(gdp_{ct})}$ for *m=1, 2* and $X_{\text{ct}} = \log(fd_{ct})\,, \log(gfc_{ct})$.

$log(fd_{ct})$ promotes income per capita by 41% of its standard deviation, while one standard deviation increase in $log(gcf_{ct})$ increases the same dependent variable by 21% of its standard deviation. The effect of the financial development is almost twice the size of the gross capital formation effect. Thus, we can conclude that on the long run the financial development has a large effect on economy's wealth.

Furthermore, Table 4 also provides the estimates of the between-dimension DOLS and the within and between-dimension FMOLS. They are presented for investigating two possible problems in our specification: (i) our estimates may not be robust to alternative panel cointegration estimation techniques; and (ii) there is a possibility of a bias since we fail to recognize that the β coefficients may be heterogeneous.

To handle the first problem we use the distinction in the estimation procedure of FMOLS and DOLS; they use different solutions to deal with the bias and endogeneity problems. The first method uses non-parametric corrections, whilst the second method adds leads and lags of the differenced regressors in the regression as parametric corrections[7], as specified in equation (6).

As a solution to the second possible problem we include the within and between-dimension estimators. In contrast to the within-dimension estimator, the between-dimension estimator allows for cross-sectional slope variation and is calculated as the average of the individual cross section $\hat{\beta}_{mc}$ coefficients and the t-statistic is the average of the individual t-statistics. Obviously, the robustness check shows that the elasticity of per capita income on both the gross capital formation rate and the M2 to GDP ratio does not vary over different estimation methods. This allows us to conclude that there are no specification problems.

Table 4: DOLS and FMOLS estimates of the long run relationship

Dependent variable: $log(gdp_{ct})$				
	DOLS		FMOLS	
Variable	Within-dimension	Between-dimension	Within-dimension	Between-dimension
fd_{ct}	0.55	0.55	0.56	0.58
	(0.05)***	(0.07)***	(0.02)***	(0.06)***
$log(gcf_{ct})$	0.38	0.21	0.17	0.19
	(0.10)***	(0.07)***	(0.02)***	(0.06)***

Notes: Standard errors in parentheses.
Lead and lag lengths for DOLS are suggested by Akaike Info Criterion.
***(**) Indicates significance at 1%(5%) level.

[7] Although, the usage of the leads and lags in the DOLS procedure is asymptotically equivalent to FMOLS's non-parametric corrections.

5. Conclusion

In this paper we empirically investigated the Schumpeterian view about financial development's effect over growth. We did this by using panel cointegration techniques which are designed to deal with problems that could possibly plague studies such as ours. By employing annual data for 16 South-Eastern and Central European countries over the period 1995-2014, we found that the long run effect of financial development on growth is positive and robust to alternative panel cointegration estimation techniques.

The effect of financial development on income per capita in South-Eastern and Central Europe is not only statistically significant, but also economically large. Particularly, it is almost twice the size of the effect of gross capital formation rate on income per capita. Therefore, we can conclude that monetary policies aimed at developing the financial system in the region, not only stimulate efficiency in the intermediation and effectiveness in the financial markets, but also directly increase the wealth of the South-Eastern and Central European nations.

However, it has to be emphasized that we did not prove any causality. In fact, the Johansen-Fisher Test for panel cointegration inferred that there may be an additional cointegration vector in our set of variables, i.e the relationship between financial development and growth could be possibly endogenous. In the future, this question should definitely be addressed in more detail. Additionally, there should be a country-specific cointegration analysis which would allow us to correctly grasp the magnitude of the effect of financial development over income per capita separately for every country from South-Eastern and Central Europe. Nevertheless, this will require time series data that spans for much longer period of time than those that are presently available.

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