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C O N T E N T

ALBANIAN HOUSEHOLD CONSUMPTION RESPONSE TO INCOME AND REAL WEALTH CHANGES, ASSESSED THROUGH MICRO- DATA

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ABSTRACT

The article, based on INSTAT's Living Standard Measurement Survey, analyses the consumption behaviour of 17.000 Albanian households, during the period 2002-2012, focused on income and real wealth changes. The empirical assessment on the micro-data show that the marginal propensity to consume out of income falls in the interval 0.14-0.17, while the one out of real wealth in the interval 0.01-0.06, dependent on the financial situation and real wealth of households.

I. INTRODUCTION

The way consumers adjust their behaviour in response to various shocks of income and wealth is one of the main macroeconomic questions, taking into account the importance and implications of the answer to this question for monetary and fiscal policies, and consequently for the entire economy. Hence, the analysis of consumers' behaviour is particularly interesting for studying this phenomenon in greater details.

Despite the significant contribution of this topic to decision-making, the evidence regarding the assessment of the marginal propensity to consume out of income and real wealth in Albania is limited, particularly that based on the assessment using micro data on households.

The economic theory suggests that consumers' utility depends on their characteristics and their level of consumption. Both the life cycle hypothesis of Modigliani and Brumberg (1954) and of Ando and Modigliani (1963), and the permanent income hypothesis of Friedman (1957) are based on the assumption that consumers prefer to smooth their consumption throughout their entire life.

Based on the life cycle theory, a change in income and current wealth would affect current consumption only if income affected the expected wealth of the households. Similarly, in his permanent income hypothesis, Friedman shows that changes in income will change the consumer behaviour. Thus, according to the life cycle hypothesis, current consumption is a function of the accumulated resources throughout an individual's life, while the permanent income hypothesis agrees that permanent consumption is a function of

[1] *The views expressed herein are solely of the authors and do not necessarily reflect those of the Bank of Albania.*

accumulated wealth and, as a consequence, of permanent income, and that only a change in one or the other would bring a change in consumption.

However, the main difference is that Friedman assumes an infinite horizon, and does not make any distinctions between human and physical capital. It is clear that these differentiations of capital are taken into account in the life cycle theory, where the life horizon of individuals is finite, and they save during all their lives, in order to ensure the same consumption during retirement as well.

There are two main approaches in empirical literature that analyse the relation between consumption, income and wealth of individuals or households: the first approach is based mainly on finding the long-term relationship between consumption, income and real wealth, based on macro-data. The main disadvantage of this approach is that the assessment of the long-term relationship between the indicators mentioned above requires a very long time series. Furthermore, the aggregate data do not take into account the heterogeneity of different household groups. The second approach is based mainly on empirical assessments on micro-data at household level. These data, compared with aggregate data, take into account the heterogeneity of households regarding consumption, income and wealth. Furthermore, as with macro-data, it is assumed that there is equilibrium between consumption, income and wealth.

The best way to see whether the changes in consumption are affected by changes in income and real wealth is through the investigation of the data offered by surveys, which provides simultaneous information on income, consumption and wealth of households. The Living Standard Measurement Survey developed and conducted by INSTAT in 2002, 2005, 2008 and 2012 is a very good source of information on the required data. Despite the lack of information on some important indicators, the Living Standard Measurement Survey (until 2012) remains one of the main sources of data in Albania, regarding households and their behaviour.

Taking into account the advantage of using data at household level, we have assessed the behaviour of households' consumption against changes in income and real wealth, for Albanian households. By using individual data of 17.000 households by the Living Standard Measurement Survey during the four cohorts, the article assesses that the marginal propensity to consume out of income falls in the interval 0.14-0.17, while the one out of real wealth falls in the interval 0.01-0.06. The results vary depending on the age of the head of the household and the characteristics of the household. Thus, estimates show that when the head of a household is aged between 35-65 years, households are less sensitive to changes in income, compared to households whose head is younger (less than 35 years old). Meanwhile, when the head of a household is older than 65 years old, we observe a higher response to changes in real wealth compared to other groups. Assessments are in line with what the life cycle hypothesis suggests on the consumer behaviour.

The paper is organized as follows. The first Section summarizes the methodology and the data used for the assessment of the marginal propensity to consume, the second Section describes the results obtained, and the last Section presents the conclusions.

II. METHODOLOGY AND DATA

DATA

The data used to assess the elasticity of consumption out of income and real wealth are taken by the Living Standard Measurement Survey (LSMS) of 2002, 2005, 2008 and 2012. Overall, the questionnaire has maintained its basic format during the four years, with few changes that do not influence the main indicators that we have taken into account for the assessment of the marginal propensity to consume out of changes in income and real wealth of the households.

The questionnaire is organised in two parts: the first part consists of seven sections that serve to collect information on the composition of the household, education, communication, employment, non-agricultural business, migration and poverty. Meanwhile, the second part consists of ten sections, which aim to collect information regarding food and non-food expenditures, housing, services and long-term use items, social assistance, other income, social capital, as well as the identification of the agricultural household.

According to INSTAT: "A household includes either one person living alone, or a group of people, not necessarily related, living at the same address with common housekeeping". Data on households are collected mainly in two phases. During the first phase data related to the monthly expenditures of the household are gathered, while during the second phase, which is the referring phase, data related with the characteristics of the households, the housing standards, labour income and other sources, etc. are gathered.

Irrespective of the number of questions and data collected, for the purposes and the focus of this study, we have taken into consideration only a handful number of indicators, defined as follows.

Households' annual total expenditure - is an indicator that is based on the self-reported monthly total expenditures of the household on, converted in an annual indicator. The item expenditure includes the total of money that a household spends for the main categories: food, non-food, energy, water, fuel and durable goods.

As an indicator of the wealth of the household, we have taken into consideration only its real wealth, because we lack data regarding other sources of financial actives. The value of real wealth of each household is calculated taking into consideration the size of the house expressed in m2, as well as the reference price published by the National Housing Agency for each area and city of Albania. On the other hand, this indicator does not take into consideration the quality of the construction, as well as the improvements that may have been made to the existing house, where the value of the house is more of a function of the house price change, than its real value.

The household total income indicator is more difficult to measure. To calculate the household annual income, first we looked at the self-reported income of each of its members. There is a significant lack of data regarding this question, since nearly half of the respondents do not declare their level of income. Hence, as a good income proxy, we used the minimum level of income that each household reports as sufficient to afford their expenditures. Furthermore, the minimum reported income was close to the income that households reported to have, and there was a positive correlation between them.

Also, our analysis excluded those households that do not report or report zero income, as well as households where the age of the head of household is younger than 25 or older than 75 years old. Our final sample consists of 16,246 households, which constitute a sufficient sample to assess the marginal propensity to consume out of income and the real value of the house (real wealth).

Since the data of the LSMS Survey is not a panel and there are no households that have taken part in all four surveys, then this way of reporting data creates difficulties for the assessment of the marginal propensity to consume, since households differ from survey to survey. Hence, to correct this issue, we based our analysis on the Deaton approach (1985), which is based on the creation of pseudo-panels. The main idea of this approach is to assess how consumption changes between groups, where household groups are defined based on similar characteristic/s, which are exogenous toward other indicators. As suggested in literature, we have used the year of birth of the head of the household as an indicator to create the groups and assess the consumption function. Also, based on age, we have created three main groups that are: "young age" households - where the age of the head of household is under 35 years old; "middle age" households - where the age of the head of household is between 35 and 65 years old; as well as "old age" households - where the age of the head of household is above 65 years old.

METHODOLOGY

As it was mentioned above, the purpose of this study is to assess empirically how consumption is affected by changes in income and real wealth. Based on the methodology of Teppa, (2014), Cristelis, Gerorgarakos and Jappelli (2014) we have estimated how changes in real consumption are a function of changes in real income, real wealth and households' characteristics, as follows:

$$\Delta lnC_{ht} \quad = \quad \alpha_{ht} + \beta \Delta lnHHI_{ht} + \gamma \Delta lnRFA_{ht} + \eta \Delta lnZ_{ht} + \epsilon_{ht}$$

Where, C_{ht} represents annual total expenditures of household h in time t, calculated as total expenditures of the household. RFA_{ht} represents the real wealth of the household, measured through the real value of the household's house. HHI_{ht} represents the annual income of the household, where as we also mentioned above, we have taken into account the minimal reported income necessary to cover monthly household expenses. Meanwhile vector Z_{ht} includes the entirety of the household characteristics that have to do with the age of the head of household and the size of the household. Carroll, Slacalek and Tokuoka (2013), with a similar approach, argue that the heterogeneity of the household is a key element to assess the marginal propensity to consume, where the models that include more elements of the households heterogeneous behaviours assess higher MPC (marginal propensity to consume), compared with model that have only a representative household.

All indicators are assessed for each household in time t, where the error term is marked with ε_{ht}. The equation is assessed with fixed effects using the panel, where we are particularly interested in the parameters β and γ, which represent the elasticity of consumption with respect to income and real wealth of the household. Also, we have calculated the marginal propensity to consume for the main three household groups divided by the age of the head of the household.

III. RESULTS

Based on the methodology presented above, we have assessed four variations of the equation of the difference of consumption. Hence, in Table 1, first column, we have presented the estimated elasticity of the change of consumption for the entire panel, treating it as a pool. While other coefficients represent the change in consumption by the change of the size of the household and the age of its head. Also, from the second column to the fourth are presented the elasticity with respect to the change of income and the value of the house for the three main age groups, young, middle and old age households.

Table1 Assessments of elasticity and marginal propensity to consume of households in Albania

	Dependent variable: annual consumption expenditures change percentage			
	Total of households	Young age households	Middle age households	Old age households
Households annual income change percentage	0.1816*** (0,000)	0.1915*** (0,000)	0.1659*** (0,000)	0,139*** (0,000)
House value change percentage	0,110*** (0,000)	0.0288*** (0,007)	0,111*** (0,000)	0.1244*** (0,000)
Head of household age	0.0003 (0.9865)	-0,155*** (0,000)	0.0236 (0.3165)	0.1348*** (0,000)
Household size change	0,236*** (0,000)	0,263*** (0,000)	0.2158*** (0,000)	0,251*** (0,000)
MPC out of income	0,169	0,178	0,154	0,129
MPC out of house value	0,013	0,003	0,013	0,015
R^2 corrected	0,147	0,127	0,117	0,156

Note: EGLS panel grouped by households.

The results show a positive and statistically-significant relationship for changes of income and house value. On average, the assessments show that 1% of income growth and house value increases consumption by 0.18 and 0.11, respectively. From the elasticities above, we calculated the marginal propensity to consume, which is obtained by dividing this elasticity with the value of assets and income against aggregate expenditures. The assessments show a marginal propensity to consume out of income at 0.17 and a marginal propensity to consume to house value at 0.013. These results show that for each additional ALL of income, Albanian households increase consumption by ALL 0.17 and for each additional ALL on the house value they increase their consumption by only ALL 0.013.

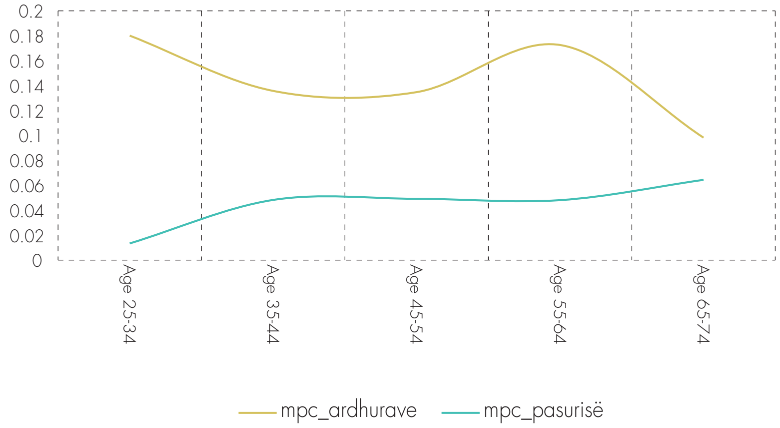
The assessments obtained for the marginal propensity to consume out of annual income indicate that Albanian households do not change or adjust significantly their consumption level to income changes. One of the reasons might have to do with the fact that households, uncertain of their income, prefer mostly to save the additional money instead of spending it. This might also be due to the level of income of Albanian households being relatively low; in addition, a high level of heterogeneity among households is noted.

On the other hand, the low level of the marginal propensity to consume out of real wealth show that not all households may change, increase or decrease consumption, in response to house value changes. In our judgement, this is because most owners expect to live in the house for a long period of time, which makes them indifferent to changes in house prices or rent prices.

To observe the heterogeneity among households, we have analysed how the elasticity of consumption with respect to income and house value changes for the group of young, middle-age and old households, presented by columns 2-4 of the Table. The results obtained are in line with our expectations based on the life cycle hypothesis, whereby it is expected that older households are more sensitive to wealth shocks, since they have a shorter period of time to consume the new accumulated wealth.

Also, to test the robustness of our results, we have reassessed our model for more detailed group-ages of the head of household, specifically for 25-34 years old, 35-44 years old, 45-54 years old, 55-64 years old and 65-75 years old. A summary of the assessment of the marginal propensity to consume out of income and real wealth is presented in Chart 1. The results obtained reconfirm what the life cycle hypothesis claims, that regarding income as well as real wealth, younger heads of households have a higher marginal propensity to consume out of income and a lower marginal propensity to consume out of real wealth, compared with the other groups.

Chart 1. MPC out of income and real wealth for various heads of the households age-groups



Source: Authors' calculations.

IV. FINAL CONCLUSIONS

Through the use of micro data on Albanian households, obtained from INSTAT's Living Standard Measurement Survey, we have assessed how changes in income and real wealth of households affect their marginal propensity to consume.

Assessments based on micro-data show that the marginal propensity to consume out of income falls in the 0.14-0.17 range, while the one out of real wealth falls in the 0,003-0,013 range, depending on the age of the head of the household. The estimates show that for households where the age of the head is in the 35-64 years old range, the marginal propensity to consume is lower than for younger households, which shows that these households are less sensitive to income changes. On the other hand, when the head of the households is older, they react more to house price changes than the other two groups. The assessments obtained are in line with the life cycle hypothesis and model.

However, despite the stability of the results, we must highlight some of the disadvantages related mainly to the data used. First, an important issue remains the quality of the data, their reporting, particularly of data related with households' debt and income source. The households' debt is almost unreported; in fact, there is a very high level of non-reporting. This also applies to income in the form of compensation of employees, from both the private and public sector. Second, non-reporting of other financial sources focuses the assessment of our estimations only on real wealth, without taking into account other financial sources. In order to have a more accurate and comprehensive assessment of the marginal propensity to consume out of income and wealth (real and financial), it is required to have a more comprehensive and accurate source of current information on consumption, income from labour and other alternative sources, as well as data on debt and real and financial wealth of households at the micro level.

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ASSESSING REAL WAGE FLEXIBILITY IN ALBANIA

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1. INTRODUCTION

Wage dynamics play an important role for macroeconomic analyses, and their appropriate measurement is essential in the policy-making process. In the real world, the existence of wage rigidities is expected to translate into persistent responses of wages to disturbances occurring in the economy. The literature has shown that the real wage flexibility may act as a crucial adjustment channel to asymmetric shocks, especially if cross-border labour mobility and fiscal flexibility is limited. Through determining the response of the economy to shocks, the nature of real rigidities has several implications for the conduct of monetary policy. In this context, a deeper knowledge of the extent of rigidity is uncommonly useful for evaluating the performance of monetary policy and building and calibrating more effectively macroeconomic models that are used for policy analysis and forecast.

This study aims at assessing the degree of real wage flexibility in Albania within a vector error-correction model, based on aggregate level data for real wages, where the real wage flexibility is defined through the responsiveness of real wages to shocks in unemployment and productivity. To the best of our knowledge, empirical evidence on wage flexibility in Albania for the whole economy is scarce. The published studies on this topic use a descriptive and graphical analysis of the development of wages over time, but do not use any econometric model to make a more rigorous analysis. In this context, this article contributes to the literature by offering empirical evidence on real wage flexibility in Albania using the latest available data.

2. LITERATURE REVIEW

2.1 THE CONCEPT OF REAL WAGE FLEXIBILITY

The labour market flexibility is a very broad notion. In principle, the disturbances in the labour market can be accommodated via two main channels: quantities (adjustment in number of workers and in working time), or prices (wages), or a combination of both. Due to limited mobility of workers within the labour markets, it is more likely to consider wage flexibility rather than migration as an efficient tool for coping with adverse shocks (Babetskii, 2006).

The flexibility of wages as the price of labour is a key determinant of labour market flexibility (see Hyclak and Johnes (1992), Boeri et al. (1998), and

Blanchflower (2001)) as adjustment in prices towards the equilibrium on the labour market, tends to be quicker and less costly than adjustment in quantities. Wage flexibility characterizes different aspects if defined as either a micro or macroeconomic concept. In the microeconomic framework, wage flexibility is typically assessed using the distribution of wages: a lack of wage decreases, for example, is interpreted as an indication of downward rigidity. In firm-level surveys, the concept of rigidity is related to the proportion of firms which freeze wages (nominal rigidity) or automatically link wages to inflation (real rigidity).

While microeconomic and survey-based estimates of wage flexibility bring valuable evidence on the distributional properties of wages and allow controlling for industry and firm effects, there are important costs involved in data collection and processing, and micro data on the Albanian labour market over time are not accessible. Hence, this article takes a macroeconomic perspective, the objective of which is to assess the degree of real wage flexibility in Albania. The use of aggregate data allows us to infer about real wage flexibility on the economy-wide level, which is of interest for policy makers.

From the macroeconomic point of view, aggregate wage flexibility can be expressed in nominal or real terms. Nominal wage flexibility is the responsiveness of nominal wages to changes in the price level or inflation, meanwhile real wage flexibility can be defined as the responsiveness of real wages to various shocks (e.g. shocks in productivity, unemployment, past wages, etc.) (Babetskii, 2006). Since the difference between real and nominal wage growth is given by inflation, real and nominal wage adjustment approach each other in a low inflation environment.

2.2 EMPIRICAL EVIDENCES ON WAGE FLEXIBILITY

Although substantial research has been performed on wage flexibility in the recent years, there is no full consensus yet in the literature regarding how to measure it. Theoretical models provide insight into the key macroeconomic influences on wage setting, but they are not informative about the particular relationship to be estimated. Most of the literature has therefore been primarily focused on studying empirical relationships.

Babecký and Dybczak (2012) present macroeconomic evidence on the extent of real wage flexibility in 24 EU member countries using Eurostat labour cost data for the period 2000Q1–2010Q2. Following the structural VAR approach, real wage flexibility is measured as the responsiveness of real wages to various shocks (shocks to productivity, unemployment and past wages). A similar approach is also used by Czech National Bank in its yearly assessments of the degree of economic alignment of the Czech Republic with the euro area (see section 2.2.1 in CNB, 2009, 2010), where real wage flexibility at the macroeconomic level is measured with the adjustment of real wages to the unemployment rate (the Phillips curve). The data shows that the impact of the global financial crisis on real wage adjustment has not been uniform across the sample countries, with some evidence for an increase in real wage rigidity.

Heinz and Rusinova (2011) estimate the degree of real wage flexibility in 19 EU countries in a wage Phillips curve panel framework, using a distributed lag structure model. The latest provides e long-run relationship between wages, prices and trend productivity including the error correction term that serves as the speed of adjustment to the equilibrium. The authors measure wage flexibility in a broad way, using two indicators: the response of wages to cyclical unemployment as well as to productivity growth. The degree of real wage flexibility tends to be larger in the Central and Eastern European (CEE) countries than in the euro area; weaker in downturns than during upswings. Moreover, there exists an inflation threshold, below which real wage flexibility seems to decrease. Finally, the authors find that part of the heterogeneity in real wage flexibility and unemployment might be related to differences in the wage bargaining institutions and more specifically to the extent of labour market regulation in different country groups within the EU.

Marques, Martins and Portugal (2010) use both micro and macro data to provide empirical research on price and wage dynamics for the Portuguese economy during the period 1992-2007 based on a VAR model and a shock decomposition analysis. As regards firms' pricing behaviour, the most noticeable finding is that prices in Portugal are somewhat less flexible than in the United States, but more flexible than in the Euro Area. Regarding firms' wage setting practices, the evidence favours the hypothesis of aggregate wage flexibility, but changes in wages occur with less frequency than changes in prices. Recent evidence from both aggregate and disaggregate wage data, however, suggests that the responsiveness of real wages to unemployment changes may have declined over the last years.

Focusing on a cross-country analysis of labour markets in the enlarged European Union, Babetskii (2006) tries to assess empirically the role of aggregate wages as a correction mechanism for dealing with economic disturbances. The study uses three alternative econometric techniques, among which a cointegration analysis and an error correction model. A comparable quarterly data-set is constructed covering 1995–2004 for four central European states (CE-4), four new EU members participating in the Exchange Rate Mechanism-II (ERM-II participants), and three peripheral members of the euro area (EMU3). The macroeconomic data does not seem to support the argument that real wages are flexible in the considered EU member countries. The pattern of rigidities at the micro level does not differ much from the estimated macroeconomic indicators of wage flexibility. A similar finding of no significant wage flexibility is reported in Radziwiłł and Walewski (2003). The authors analyse a broad set of indicators at the macro and micro levels and conclude that wages are not flexible in six new member states (accession countries at that time), except for some evidence of flexibility in Lithuania.

In conclusion, most theories of wage determination (see Bean 1994 for a review) and empirical research confirm that, on average at the macro level, the aggregate nominal wage should be proportional to the (expected) price level and productivity, a decreasing function of the unemployment rate (an indicator of outside opportunities), and also influenced by a set of other factors including demographics, taxation, union power, and labour market institutions and policies.

3. EMPIRICAL ANALYSIS

3.1 WAGE FLEXIBILITY MEASURES

The economic disturbances enforce a change of the wage rate and flexibility is a measure of the pace with which actual wages respond to changed market conditions. In this article, we conceptualize real wage flexibility in a broad way, using two indicators. In the first place, similarly to Babetskii (2006), wage flexibility is defined through the responsiveness of real wages to unemployment rate. On the contrary, wage rigidity implies either an absence of such effect, or a considerably retarded one. The unemployment rate captures mostly supply-side determinants, as wage requests by unions are expected to become more moderate in the presence of higher unemployment. As a second indicator of wage flexibility, we assess the responsiveness of wages to changes in productivity, which has attracte less empirical research than the link with unemployment. Labour productivity is aimed at capturing labour demand: the higher the productivity of labour at given price level, the higher the nominal wages firms are willing to pay. For each of these indicators, we compare not only the coefficient sizes, but also the speed and lag structure of the response. As Kittel (2001) argues, it is important to consider not only the differences in the flexibility outcomes, but also in the way these outcomes have been achieved. The timeliness of the wage response to economic developments is also relevant, since if it is strongly delayed, then the adjustment might not be optimal any longer in the presence of new shocks.

In analogy with existing work (see Nickell, 1988; Manning, 1993; Bell, Nickell, Quintini, 2002; Nunziata, 2005), the estimated dynamic wage equation can be obtained as a reduced form specification incorporating both demand and supply-side labour market determinants. This assumes that there is an equilibrium relationship between the real wage level, the unemployment rate and labour productivity to which real wages will converge even if there are transitory shocks that divert wages from this equilibrium. Real wage flexibility is low when it takes time for a dis-equilibrium in the labour market to be eliminated. Similarly, a high degree of real rigidity entails a situation where real wages or mark-up of price over marginal costs respond little to demand pressures. Such a framework does not exclude the possibility of reverse causation and multiple long-run relations among the considered variables.

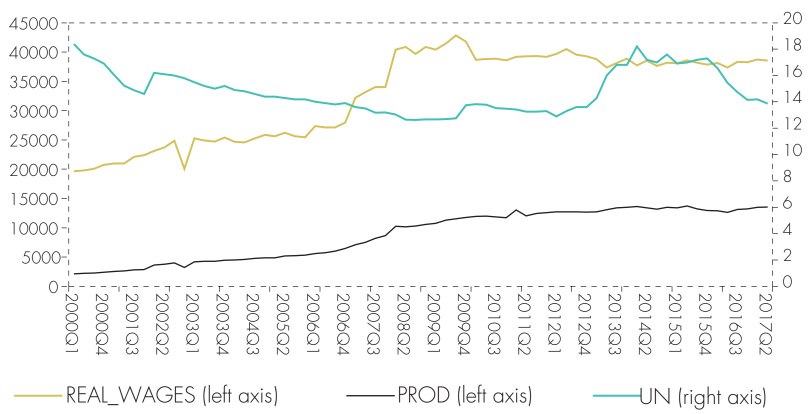
3.2 DATA

This section describes the data used for the empirical analysis of the article. The dataset includes quarterly time series on real wages, unemployment rate and labour productivity for the period 2000Q1-2017Q2. Except the unemployment rate, all the other series are seasonally adjusted using TRAMO/ SEATS method.

The series of real wages is calculated as a ratio of the nominal wages of the private sector to the consumer price index multiplied by 100. The series of nominal wages is published by INSTAT in annual terms since 2000. The quarterly data are interpolated for the period 2000-2002 using the wages of the public sector, while starting from 2003 they are interpolated in line with the wage index from the Survey of Economic Enterprises conducted by the National Institute of Statistics (INSTAT). The unemployment rate is taken from the Quarterly Survey of Labour Force database conducted by INSTAT as well. Labour productivity is calculated by dividing the Gross Domestic Product at constant prices with the number of employees.

Figure 1 illustrates graphically the relation of the real wages with unemployment rate and labour productivity over the considered period. Real wages and productivity tend to follow similar positive trends (similar slopes), meanwhile the negative relation between wages and unemployment rate seems to be present until 2012Q1, and then after 2015Q4, with few short interruptions (short increase in the unemployment rate).

Figure 1 Development of the considered indicators over the period 2000Q1-2017Q2



Source: INSTAT and author's calculations, 2017.

3.3 METHODOLOGY

A Vector Error Correction model (VECM) is used to analyse the relation between real wages, productivity level and unemployment rate as a usual applied methodology in examining more than one endogenous variable, which also incorporates both short-run and long-run dynamics. A limited number of variables was chosen for the regression, because the aim of the analysis is not to explain real wages as much as possible, but rather to see whether wage developments observed in a certain time period were in line with what would be predicted on the basis of fundamentals or whether they were driven by some temporary or structural factors (policy or market driven for instance).

Incorporating demand and supply-side labour market determinants, the long-run wage equilibrium relationship is specified as:

$$\ln(real\_w_t) = \alpha + \beta_1 \ln(prod_t) + \beta_2 un_t + e_t \qquad (1)$$

where *t* indexes the time period, ***real_w*** denotes real wages, un is the unemployment rate, ***prod*** is the GDP per total employment and ***e*** is the disturbance term. While, the dynamic wage equation has the following form:

$$\Delta \ln(real\_w_t) = \mu + \sum_{i=1}^{p} \eta_i \Delta \ln(prod_{t-i}) + \sum_{i=1}^{p} \vartheta_i \Delta\ (un_{t-i}) + \gamma \widehat{e_{t-1}} + \varepsilon_t \qquad (2)$$

where $\widehat{e_{t-1}}$ is the residual from equation (1) and therefore γ measures the speed of adjustment to a random shock.

Except for the unemployment rate, we include the natural logarithm for real wages and labour productivity, in order to obtain the semi-elasticity of real wages with respect to unemployment and their elasticity with respect to labour productivity. In addition to the above basic specification, other authors include terms of trade among the explanatory variables (higher terms of trade expected to be reflected in higher wages, other things being equal). Alternative specifications are also estimated using nominal wages as dependent variable and adding the price level to the list of the explanatory variables.

As a first step of the estimation procedure, the stationarity properties of the data are assessed by applying the standard techniques: the augmented Dickey–Fuller (ADF) unit root tests. Overall, the series of real wages, labour productivity and unemployment can be characterized as integrated of order one I(1). Engel and Granger (1987) note that a linear combination of two or more I(1) series may be stationary (or I(0)), in which case we say that the series are cointegrated. Before testing the existence of such a relationship, it is necessary to determine first the optimal lag length in the vector autoregressive representation of the model. Referring to the different lag selection criteria, the number of lags is mostly found to be four, and only in one case one. As the VEC specification applies to cointegrated series, we should run cointegration tests to determine if there is any cointegrating relation between the variables and further the number of these cointegrating relations. These tests are performed with three lags in our case, as the VEC model is estimated with one lag less than the optimal lag length of the respective VAR specification. Table 1 summarizes the results of the two Johansen tests (Trace and Max-Eigen value). They indicate the existence of one cointegrating relation among the real wages, unemployment and productivity in most of the cases.

Table 1. Johansen tests results summary.

Sample: 2000Q1 2017Q2					
Included observations: 66					
Series: L_REAL_W L_PROD UN					
Lags interval: 1 to 3					
Data Trend:	None	None	Linear	Linear	Quadratic
Test Type	No Intercept	Intercept	Intercept	Intercept	Intercept
	No Trend	No Trend	No Trend	Trend	Trend
Trace	1	1	1	0	0
Max-Eig	1	1	1	0	0
*Critical values based on MacKinnon-Haug-Michelis (1999)					

3.4 MODEL RESULTS AND THEIR INTERPRETATION

After confirming the existence of one cointegrating relation among the variables, we estimate this relation using a VEC model with three lags. The estimated coefficients presented in the Table 2. The specification can be regarded as a reduced form supply-demand system for the labour market. It also assumes that there is an equilibrium relationship between real wages, labour productivity and unemployment rate, to which real wages will converge even if there are transitory shocks that divert wages from this equilibrium.

Table 2 Estimated coefficients of wage equation through a VECM approach.

Dependent Variable:	L_REAL_W(-1)
L_PROD(-1)	0.602720*** (0.05370) [11.2229]
UN(-1)	-0.021958 (0.01780) [-1.23327]
C	4.438153 (0.65385) [-6.78773]
Error Correction Term	-0.120607*** (0.04816) [-2.50448]
R-squared=	0.485214

*Note: *significance at 10%, **significance at 5%, ***significance at 1%.*

In addition, we perform several diagnostic tests to see if the necessary conditions of VEC approach are satisfied. According to these tests, there is no evidence of serial correlation and heteroskedasticity of the VECM residuals, and the stability test was passed successfully. Only the assumption of the normal distribution of model residuals is not satisfied, but this does not have implications for the aim of our analysis.

As it can be seen from Table 2, both explanatory variables have the expected signs, but only productivity results to be statistically significant in the long-run relationship. The insignificant effect of unemployment on wages can be explained by the vertical long-run Phillips curve. Notice that the long-run coefficients cannot be interpreted as elasticities in the strict sense, since each of the coefficients incorporates the effect of shocks to all variables (see Lutkepohl, 1994). However, as the objective of this article is to assess real wage flexibility, in the estimated cointegrating relation, we are more interested in the error correction term (ECT), which represents the speed of adjustment of real wages to the long-run equilibrium relationship rather than the estimated variable coefficients. As it can be seen, the ECT is statistically significant at 1% level of significance and it has a negative sign, which confirms the existence of the long-run relation between real wages, productivity and unemployment rate. Its value indicates that real wages need almost two years (8.3 quarters) to converge to their equilibrium values, in case of any shocks that diverge them from their steady-state.

In order to have a clearer view about the flexibility of real wages in Albania, we deepen the analysis further for the short-run period and investigate how important were the different shocks in accounting for the observed fluctuations in real wages, by looking at the forecast-error variance decomposition of real wages. The results of this analysis are shown in Table 3. As it can be seen, in the very short run (2-3 quarters), most of the variation in real wages forecast errors is explained by their past values (more than 70%) rather than by labour productivity or unemployment rate, with a higher role of productivity relative to unemployment shocks. However, their contribution to forecast error variance of real wages is low, which means that real wages do not respond significantly towards productivity and unemployment in the short run. One should be aware that these conclusions are based on a particular statistical methodology and may be affected by certain statistical problems due to the quality of the underlying data series.

Table 3 Sources of real wage flexibility.

Var. Decomp. L_REAL_W:			
Period	L_REAL_W	L_PROD	UN
1	99.29054	0.000000	0.709458
2	94.74190	4.363262	0.894835
3	81.15108	18.12083	0.728087
4	71.57257	22.28286	6.144565
5	75.38749	17.79730	6.815213
6	72.68193	20.42679	6.891274
7	68.61941	23.60401	7.776580
8	66.23193	24.64845	9.119627
9	67.39298	23.15692	9.450103
10	66.79215	23.71309	9.494765
11	65.58419	24.93315	9.482660
12	64.97251	25.35070	9.676783

Although the model appears to be correct econometrically and satisfies all the necessary conditions dictated by the estimation method, at the end of this session, it is important to stress that the results of this study are subject to a number of caveats. An important limitation of the analysis is the fact that besides real wage adjustment, the adjustment in the labour markets can take many other forms (e.g. migration, changes in labour force, participation and part time arrangements or other factors). In the presence of these other mechanisms, real wages may react less than they would otherwise. A further complication is the existence of relatively extensive grey economy in Albania, which may influence the behaviour of wages, distort the official wage figures and might itself represent an alternative labour market adjustment channel. For instance, in the case of an aggregate demand shock, employers are more likely to fire first the informal employees and to cut wages of the workers unprotected by the labour protection regulations of the formal economy (Heinz and Rusinova, 2011).

4. CONCLUDING REMARKS

Based on aggregate level data for the period 2000Q1-2017Q2, this article aims at assessing the degree of real wage flexibility in Albania within a vector error-correction model, where real wage flexibility is conceptualized using two indicators: the responsiveness of real wages to unemployment rate and their responsiveness to changes in productivity.

The Johansen cointegration tests confirm the existence of one cointegrating relation among the three variables. The error correction term suggests that real wages need more than two years to converge to their long-run equilibrium values, a relatively long period of time which indicates real wage rigidity in Albania. In addition, we perform a variance decomposition analysis for the forecast errors of real wages, in order to identify which of the explanatory variables shocks contributes more to real wage fluctuations. In the very short run (2-3 quarters), most of the variation in real wages forecast errors is explained by their past values (more than 70%) rather than by labour productivity or unemployment rate. This analysis confirms the fact that real wages are slightly flexible towards these two variables. However, these conclusions are based on a particular statistical methodology and may be affected by certain statistical problems due to the quality of the underlying data series.

At the end of this article, it is important to highlight that an overall assessment of wage developments needs to look at a broader set of variables and shouldn't be limited to the indicators used here. Other aspects that could be included in the analysis are labour force shocks and the change in the sectorial structure of the economy. Furthermore, a disaggregated wage analysis based on firm-level surveys data would be helpful to have more detailed information on the frequency of wage changes in Albania. Government policies should also be taken into account when analysing the labour cost developments, as they affect wage dynamics both directly and indirectly through minimum wages, wage indexation laws, social security contributions and direct labour taxation etc.

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A DUAL INDEX OF INCOME INEQUALITY FOR ALBANIAN HOUSEHOLDS

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ABSTRACT

This article introduces an alternative method besides the Gini coefficient to measure income inequality for Albanian households, presenting two distribution indexes: the Rawlsian index and the Gini dispersion index. Results based on micro data for Albanian households suggest that inequality has decreased as a result of the redistributive effect among households with higher income, while the magnitude of inequality remained almost constant.

1. INTRODUCTION

Economic inequality is in the core of economic research. Since inequality of resources is what drives change, development and economic growth, it is very important to pay special attention to it in the economic literature. Welch (1999) states that: "The entire economy stems from inequality. Without inequality of resources and skills, there would be no trade and specialization. As a result, there would be no economy.

Since household income drives consumption and savings, income inequality is one of the main components of economic inequality. We should point out that income inequality is a broader concept than poverty, in the sense that it does not focus only on low-income groups, but on the population as a whole. Income inequality has attracted much interest in the academic literature (see Levy and Murnane (1992); Burtless (1995); Gottschalk and Smeeding (1997)]. Researchers and statisticians have tried to measure income inequality by using direct methods or by constructing and applying indexes to calculate the level of equality or inequality in an economy.

The most direct method of measuring inequality ranks the population from those with lower income to those with higher income, and calculates the percentage of income (or expenses) attributable to each quantile or decile of the population. The poorest quantile accounts for about 6-10 % of aggregate expenditures, while the richest quantile accounts for about 35-50% (World Bank (2008)). Similarly, the Hoover index simply calculates the percentage of all income that would have to be redistributed to achieve a state of perfect equality.

The most used and simple method to be applied among inequality indexes is the Gini coefficient (1936). It is considered a good statistical measure of inequality and welfare. [Gini (1936)] and it is derived from the Lorenz's curve. The Gini coefficient ranges from zero (full equity) to one (full inequality), but usually ranges between 0.3-0.5 per capita. In Albania, Çami (2017) has calculated a Gini coefficient of expenditures of 0.31 for 2012.

Even though the classic Gini coefficient has many desirable features –: independent mean, independence of population size, as well as symmetry and Pigou-Dalton transfer sensitivity can not be easily decomposed to understand inequality sources. Moreover, Gini does not determine in which part of the distribution inequality occurs. As a result, two different income distributions may have the same Gini coefficient.

In the literature, there are also other income inequality measures, such as the Theil and the Atkinson index, which make it possible to examine the effects of inequalities in different areas of the income spectrum, thus enabling more meaningful quantitative assessments of qualitatively different inequalities. Atkinson (1975) developed an index that includes a sensitivity parameter (ε), which ranges from zero (which means the researcher is indifferent to the nature of the income distribution), to infinity (where the researcher is concerned only for the positioning of income at the bottom of the distribution). Atinkson states that this index was a way to include Rawls' concept for social justice when measuring the income inequality. Also, the Theil index (Theil, 1967) is an entropy measure. A Theil index equal to zero implies full equality. A Theil index equal to one means that the population distribution entropy is almost the same to that of a population with a distribution 82:18, which is more unequal than a classical distribution "Pareto 80:20".

On the other hand, poverty and inequality measures, same as the three indexes above, which are based on official data, are often used to drive and prioritize policy actions. However, it is a well-known fact that such "standard" measures based on consumption and household income or on income aggregates have several shortcomings, specifically related to normalization, low dimensionalism and ad hoc weights.

For these reasons, this study considers a contemporary measurement of a dual income inequality index. The paper uses households' micro data from the Living Standard Measurement Survey (LSMS) for 2002, 2005, 2008 and 2012, and calculates the inequality in the household income distribution, using the Rawlsian index and the Gini dispersion index, based on the methodology of Park et al. (2017). These two indexes can be considered as a decomposition of the classic coefficient Gini: which assesses the magnitude (Rawlsian index) and the dispersion (Gini dispersion index) of the unevenly distribution income (UD). Such estimates provide a necessary additional information beyond the classic coefficient Gini on understanding the welfare and income models of household income.

The results suggest for an added value in using the dual index when measuring inequality. The decomposition of the Gini coefficient suggests that in Albania, movements in the total population inequality were a result of the redistribution of the UD income, whilst the magnitude of inequality has remained almost the same between 2002, 2005, 2008 and 2012.

Section II presents the methodology of this article, including an explanation of the concept of unevenly distributed income (UD income) and the application of the dual index of income inequality. Section III presents the data used and the selection of the final sample. Section IV presents the results and Section VI presents the conclusions.

2. METHODOLOGY

2.1 THE RAWLSIAN INDEX

Let's assume a population of n individuals. The income distribution of this population is given through the vector $y = (y_1, y_2, \dots, y_n)$,where y_1 is the income of the i-th household. Without loss of generality, we can assume that the individual incomes are ordered, i.e $y_1 \leq y_2 \leq \cdots \leq y_n$. Under these assumptions, the total income of the population and the mean income of the population can be denoted through the two respective formulas: $S(y) = \sum_{i=1}^{n} y_i$ and $\mu(y) = S(y)/n$.

Income inequality is the difference in income between households. In this sense, all households have (1) a share of their income in common (equal to y_1), and 2) the remaining income (difference from y_1). Since in a population of ***n*** households, all of them have y_1 ***n***-times income in common, it is possible that we analyse inequality considering only the remaining parts of income.

Mathematically, if the population consists of 5 individuals with income distribution: $y = (y_1, y_2, y_3, y_4, y_5)$. Taking into account the above information, y_1 is included in income y_2 until y_5 at least once, therefore $5y_1$ income are distributed evenly and $S(y) - 5y_1$ are distributed unevenly. Perk et al. (2017) only considers $S(y) - 5y_1$ income to measure inequality and refers to it as UD income (Unequally distributed income), as we are going to call it.

More concretely, in this article we will measure: (1) what is the total of UD income; and (2) how unequal is the UD income. These elements will be called (1) magnitude of inequality, or Rawlsian index, after Rawls (2011), and (2) dispersion of inequality, or Gini dispersion index.

Park et al. (2017) the UD income distribution of $\mathbf{y}$ by $d = (d_1, d_2, \dots, d_n)$, where $d_i = y_i - y_1$. As shown above, the total UD income is given by the formula: $S(d) = S(y) - ny_1$. $S(d)$ represents the magnitude of inequality and our first index (Rawlsian). Conversion of $S(d)$ in an index requires a reference magnitude where we use $S(y)$ for this purpose as a reasonable reference.

Therefore, the Rawlsian index, is expressed by the following formula:

$$\frac{S(d)}{S(y)} = \frac{\mu(d)}{\mu(y)} = 1 - \frac{y_1}{\mu(y)};$$

where, $y_1\mu(y)$ is the relative benefit of the poorest household and $1 - y_1/\mu(y)$ is the relative disadvantage of the poorest household. According to the formula, an increase in the index, indicates an increase in the magnitude of income inequality. It should be noted that this index is generally expected to have values lower than one, but it is possible to mathematically also have values greater than one.

This magnitude fulfils the conditions of the Rawl's difference principle, and for this reason we shall call it Rawlsian index and we will denote it with $R(d|y)$. Some of the properties are: (1) scale invariance; (2) replication invariance; (3) anonymity axiom; and (4) the principle of transfer; (5) translation invariance. For more information refer to Perk et al. (2017), who in their article mathematically proved that the Rawlsian and Gini dispersion indexes meet all of these properties and are consistent with economic theories on inequality.

Through the two indexes, we can also calculate the classic Gini coefficient, as we will prove below. However, this is not possible when using a logarithmic function since the first difference of d_1 is always zero.

2.2 DECOMPOSITION OF THE GINI DISPERSION INDEX

As $y_i - y_j = d_i - d_i$, the **Gini G(y)** classic coefficient is written as :

The sub-distribution of **d**, which consists of the i largest UD income, is expressed according to the formula:

$$G(y) = \frac{\sum_{i=1}^{n}\sum_{j\geq i}^{n}|y_i - y_j|}{n(n-1)\mu(y)} = \frac{\mu(d)}{\mu(y)} = \frac{\sum_{i-1}^{n}\sum_{j\geq i}^{n}|d_i - d_j|}{n(n-1)\mu(d)} = R(d|y)G(d).$$

While its distribution is denoted as:
$\mu(d_i) = (d_{n-i+1} - d_{n-i+1}, d_{d-i+2} - d_{d-i+1}, \dots, d_n - d_{d-i+1})$

Then $d = d_n$ and $d = \mu(d)$. We can derive the following formula:

$$G(d_i) = \frac{\sum_{j=n-i+1}^{n}\sum_{k\geq j}^{n}|d_j - d_k|}{i(i-1)\mu(d_i)}$$

$$= \frac{\sum_{k=n-i+1}^{n}|d_{n-i-1} - d_k| + \sum_{j=n-i+2}^{n}\sum_{k\geq j}^{n}|d_j - d_k|}{i(i-1)\mu(d_i)}$$

$$= \frac{i\mu\big(u(d_i)\big) + (i-1)(i-2)\mu(d_{i-1})G(d_{i-1})}{i(i-1)\mu(d_i)}$$

$$= \frac{1}{i-1}\frac{\mu(u(d_i))}{\mu(d_i)} + \frac{i-2}{i}\frac{\mu(d_{i-1})}{\mu(d_i)}G(d_{i-1}). \quad (2)$$

Given that $\mu(d_2) = (0, d_n - d_{n-1})$ and $G(u(d_2)) = 1$ we have:

$$G(d_2) = R(u(d_2)|d_2)G\big(u(d_2)\big) = R(u(d_2)|d_2)$$

Using equations (2) and (3), we can express the Gini dispersion index as:

$$G(d) = G(d_n) = \sum_{i=2}^{n} w_i \frac{\mu(d_i)}{\mu(d_n)} R(u(d_i)|d_i), (4);$$

where $u(d_n) = d_n$, $R(u(d_n)|d_n) = 1$ and

$$w_i = \begin{cases} \left(\prod_{j=1}^{n-i}\frac{n-j-1}{n-j+1}\right)\frac{1}{i-1}, \text{if } i = 2,3,\ldots,(n-1), \\ \frac{1}{n-1}, \text{if } i = n, \end{cases}$$

Then $\mu(d_i)$ can be interpreted as the average benefit of households in d_i. As a result, the average relative benefit of households in d_i is expressed by the formula $\mu(d_i)/\mu(d_n)$. The Rawslan index of d_i is denoted by $R(u(d_i)|d_i)$ Equation (4) states that the Gini dispersion index is a weighted average of Rawlsian indexes for d_i.

Since $w_{i+1}w_i = (i+1)(i+2)$ is less than one and $\mu(d_i)R(u(d_i)|d_i) = \mu(u(d_i))$ an effective way to reduce the Gini distribution index is to diminish $\mu(u(d_i))$ for small i, that is to diminish the UD income gap between the richest households.

Let's denote with y_i the sub-distribution of y, which consists of the i largest incomes, that is, $y_i = (y_{n-i+1}, y_{n-i+2}, \ldots, y_n)$. Then $y = y_n$ and $u(y_i) - u(d_i)$. Using equations (1) and (4), we can write the Gini coefficient as:

$$G(y) = \sum_{i=2}^{n} w_i \frac{\mu(u(y_i))}{\mu(y_n)} = \sum_{i=2}^{n} w_i \frac{\mu(y_i)}{\mu(y_n)} R(u(y_i)|y_i)$$

Similar to the the Gini dispersion index, the Gini coefficient is a weighted average of the Rawlsian Index for y_i. The Gini coefficient can be reduced using the same method similar used for the Gini dispersion index.

3. DATA

The data used for the analysis are secured from the Living Standards Measurement Survey LSMS for 2012. LSMS is a survey conducted by INSTAT every four year with a sample of 6672 households that provide information on a different issues, including data on household expenditures for different categories of products and services and income.

The questionnaire contains extensive information on some different aspects of living in Albania, on an individual and household level. INSTAT definition of the household is "A household includes either one person living alone, or a group of people, not necessarily related, living at the same address with common housekeeping. The LSMS's questionnaire is appropriate to be used for our empirical analysis, as it contains comprehensive information about income and characteristics of households".

The sampling methodology is slightly different between the LSMSs of 2002, 2005, 2008 and of 2012: (i) the sample size in the first three cohorts is equal to 3,500 households while in the last cohort it is equal to 6672 households; (ii) in the first three cohorts the households were sampled based on the region of residence: central, mountain, coastal, Tirana and based on 432 primary sampling units (PSUs) and in 2012 were sampled by district (36 districts) based on 834 PSUs; (iii) in the first three cohorts, the selection of households is based on the 2001 Population and Housing Census and in the last cohort it is based on the year one for 2011.

Despite these differences, the four samples are considered by INSTAT representative of the population at the national, regional and urban/rural level. The raw sample consists of 17121 households.

Despite the extensive information provided by the LSMS, for the purposes of our calculations, we only consider the indicator of household income. The questionnaire collects information about income in three ways: (i) by self-reported total income from labour and other sources in ALL for a one-month period; (ii) self-declared monthly income from labour and other sources collected for each working household member; and (iii) the minimum income necessary to cover the monthly household expenses.

When comparing the information in (i) and (ii), we cannot find any significant difference between the two indicators. On the other hand, there is lack of information on the two indicators for about 6% of the surveyed households, which refuse to report their monthly income from labour. On the other hand, the indicator (iii) is an indicator which, besides the income component, also includes the concept of subjective poverty perceived by households. As such, this indicator is reported non-zero for 100% of households.

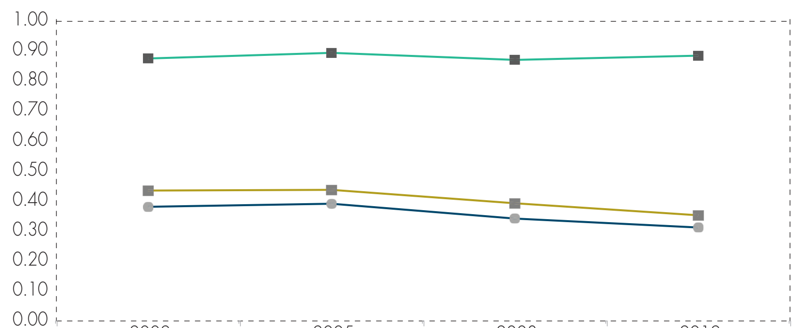
Considering our circumstances, we choose to use the indicator (i) as an approximate for true household income. Although the indicator (iii) is better reported in the survey, we want to study household income by overlooking the subjective elements, as in the case of indicator (iii) of income. We have chosen to multiply the total monthly income in ALL of households with 12, to obtain an annual income variable, as the results are presented in annual terms. It is clear that UD income should be non-negative and should be zero. Hao and Naiman (2010) and Cowell (2011) suggest dropping observations where income is non-positive. OECD (2013) suggests replacing the negative incomes with zero. Bell`u and Liberati (2006b) suggest replacing zero income with small arbitrary positive numbers and calculating income inequality indexes using a logarithmic function. One advantage of the approach on income of Park et al. (2017), using the concept of UD income, is that a non-positive income correction is not needed for obtaining unbiased and significant results. In our case, no household reports zero income and therefore we choose only to drop observations where households have refused to report their income (6%). Moreover, in the sample there are households which are considered to be "outliers". For this reason, we have dropped observations where the age of the head of the household is more than "85" and less than "25", as well as observations where the number of household members is more than "8."

After dropping these observations, the final sample resulted to be a total of 16055 families that are sufficient to provide significant results. Also, all values are weighted by the weights of the survey to estimate representative results at the population level.

4. RESULTS

The graphic below shows the calculated Rawlsian (R (d/y)) and Gini's dispersion (G (d)) indexes, as well as the classic Gini (G (y)) coefficient.

Chart 1. Rawlsian, Gini Dispersion Index and Gini Classic Coefficient.



	2002	2005	2008	2012
G(y)	0.381	0.391	0.342	0.312
R(d/y)	0.876	0.894	0.871	0.885
G(d)	0.435	0.437	0.393	0.353

Source: Author's calculations.

The classical Gini coefficient is a coefficient which, in the present case, according to formula (1), should be equal to the product of the Rawlsian and the Gini dispersion index, in the sense that the two indexes can be described also as a decomposition of the Gini classic coefficient. In this case, the calculated Gini coefficient is exactly equal to this value, since none of the households used in the estimations reported a non-positive income.

As for the indexes trends, the results suggest for a Rawlsian index almost constant over time, fluctuating in the range 0.87-0.90 near the value of 1. On the other hand, as shown in the chart, the Gini coefficient and the Gini dispersion index move together over the years, this being more evident in 2012. Only in 2012 we see a slight increase in the magnitude of inequality even though its dispersion has decreased. These results suggest that the Gini classic coefficient in 2008 and 2012 has not decreased as a result of the decrease of the magnitude of inequality, but as a result of the redistribution of income among households in the UD distribution.

The analysis was repeated also for "per capita" income since the economic literature on inequality supports the idea that per capita income is a better indicator to be used. According to the literature, income/expenditure indicators should be weighted regarding differences in living costs, which vary eg. by the size of the household, and small changes in such assumptions have major effects on estimated poverty rates (see e.g. Lanjouw et al., 2004 and Deaton, 2010). The results obtained are in line with the results obtained from the non-capita analysis, except that the Gini coefficient has been calculated to have a higher value during all four years. These results are shown in Table 1 in the appendix.

Also, we calculated the two indexes and the Gini classic coefficient for different percentages of income (5, 25, 50, 75, 95) for each year. The results are presented in Table 1 below. This analysis is complementary to the above information on the inequality of the population, as it reduces the effect of the outliers statistics on the top and bottom of the distribution, as well as it allows for a comparison between the two groups below and above the mean of the population.

Table 1 Inequality indexes by percentiles 5, 25, 50, 75, 95.

2002			
	G(y)	R(d|y)	G(d)
<5p	0.06	0.549	0.109
<25p	0.08	0.738	0.108
>75p	0.218	0.411	0.53
>95p	0.206	0.325	0.634
<50p	0.159	0.738	0.215
>50p	0.277	0.509	0.544
2005			
	G(y)	R(d|y)	G(d)
<5p	0.176	0.779	0.226
<25p	0.208	0.899	0.231
>75p	0.277	0.482	0.575
>95p	0.235	0.381	0.616
<50p	0.245	0.935	0.262
>50p	0.335	0.586	0.571
2008			
	G(y)	R(d|y)	G(d)
<5p	0.055	0.529	0.104
<25p	0.122	0.677	0.180
>75p	0.21	0.359	0.585
>95p	0.208	0.348	0.597
<50p	0.164	0.756	0.217
>50p	0.244	0.543	0.449
2012			
	G(y)	R(d|y)	G(d)
<5p	0.05	0.471	0.106
<25p	0.103	0.617	0.167
>75p	0.172	0.667	0.257
>95p	0.13	0.322	0.41
<50p	0.15	0.706	0.212
>50p	0.21	0.437	0.48

Source: Authors' calculations.

According to the results shown in the table, referring to the Gini classic coefficient, inequality is more pronounced at the top 25% of the income distribution (especially for households at the 75% -95% level), while at the bottom 25% of the distribution, coefficient estimates suggest for a high level of equality. Generally, households that are in the > 50% of income are more unequal than those in the <50% (see the first column), as the literature would suggest.

As concerning the decomposition by using the two indexes, the magnitude of inequality is higher in the bottom of the income distribution versus the top, and the opposite occurs for the distribution of inequality by the Gini dispersion index. The disaggregated values by the two indexes support the Gini coefficient differences observed between the lower and upper percentiles of the distribution.

Despite fluctuations, these tendencies are present in all four years, although in 2012 we see a smoothing in the differences between percentiles. On the other hand, although these results are complementary and better explain the aggregate results for the total population, they also support the developments of the Gini classical coefficient, Rawlsian index and Gini distribution index observed in aggregated terms.

5. CONCLUSIONS

This article uses micro data from the 2002, 2005, 2008 and 2012 Living Standards Measurement Survey to calculate income inequality using a dual index: the Rawlsian index, which measures the magnitude of inequality, and Gini's dispersion index, which measures the dispersion of inequality.

The results suggest for an added value when using a dual index in measuring inequality. An analysis based on a dual index provides information that cannot be provided by a unique index such as the classic Gini coefficient of income. The decomposition of the Gini coefficient suggests that in Albania, movements in the inequality of the population were a result of the redistribution of the UD income, whilst the magnitude of inequality has remained almost the same between 2002, 2005, 2008 and 2012. These results are also based on a disaggregated percentile estimation (5, 25, 50, 75, 95), where furthermore, it is noted that inequality is more pronounced in the upper part of the income's distribution.

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6. APPENDIX

Table 1 Income Inequality Indexes Per Capita

	G(y)	R(d|y)	G(d)
2002	0.43	0.833	0.167
2005	0.44	0.897	0.103
2008	0.35	0.855	0.145
2012	0.36	0.867	0.133

Source: Authors' calculations.