SEEJE

South-Eastern Europe Journal of Economics

THE OFFICIAL JOURNAL OF THE ASSOCIATION OF ECONOMIC UNIVERSITIES
OF SOUTH AND EASTERN EUROPE AND THE BLACK SEA REGION

Vol 17 | No 1 | SPRING 2019

web site: http://www.asecu.gr
ISSN 1109-8597
ISSN On line: 1792-3115

ASECU

FUNDING INSTITUTION
UNIVERSITY OF MACEDONIA
THESSALONIKI, GREECE

SEEJE
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ISSN 1109-8597
ISSN On line: 1792-3115
Thessaloniki, Spring 2019

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ASECU South-Eastern Europe Journal of Economics 1 (2019) 7-31 SEEJE

LARGE FIRE DISASTER AND THE REGIONAL ECONOMY: THE 2007 CASE OF THE PELOPONNESE

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Abstract
The article explores the evolution of annual personal incomes in the Peloponnese, in southern Greece, at the disaggregated (local community) level from 2001 to 2010, i.e., before and after the 2007 fires, in order to better understand the medium-term economic effects of these fires in the burned and other areas of the region outside the fire path. The paper considers a number of econometric approaches and ends up engaging in a series of cross-sectional regressions of income-filer figures and average incomes to study the situation year after year. Findings indicate that, by and large, no inordinate drop or rise in average income figures or income-filer numbers is detected in the aftermath of the fires, especially in the communities damaged by them.

JEL Classification: C21, Q54
Keywords: Employment, Income, Large Wildfires, Economic Impact, Greece

The article has benefited from suggestions made by an anonymous referee, and by comments made to an earlier version by participants in the 2nd International Conference in Applied Theory, Macro and Empirical Finance held in Thessaloniki (May 2016). Additional conversations with the economic development personnel of the regional administration of Western Greece are greatly appreciated. The usual disclaimer applies.

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1. Introduction

The article examines the medium-term effects of the large fires that occurred in 2007 in the Peloponnese on the region's economic welfare. To that end, it explores the evolution of personal local incomes in the region before and after these fires. To our knowledge, the impact of the particular fires on people's income has not been empirically studied yet.

Large fires, like other large scale disasters, bring about many more consequences over and above those that are classified as "economic", with local income being an easily and frequently measured variable, though not the only economic variable affected. In fact, measuring other economic dimensions may be quite challenging (Butry *et al*., 2001; Hesseln *et al*., 2004; Loomis, 2004; Masvar *et al*., 2012; Richardson *et al*., 2012; Kalabokidis *et al*., 2014). A good number of large fires, especially the very large fires with extreme behavior characteristics that cause significant loss of assets and high fire suppression costs, are classified as megafires, partly on account of their economic impact (Bartkett *et al*., 2007). A number of the fires that occurred in the Peloponnese in 2007 are, indeed, regarded as megafires, while the rest are related as large fires or wildfires (Maditinos and Vassiliadis, 2011; San-Miguel-Ayanz *et.al*., 2013). That said, the issue of a fire's impact on income or output is not straightforward. In this day and age, disasters may have quite dissimilar short- and medium-term consequences around the world (e.g., Noy, 2009).

The rest of the article is organised as follows: Section 2 reviews the literature on the issue. Section 3 provides background information on the region and the 2007 fires. Section 4 describes the data. Section 5 deals with certain methodological issues. Section 6 describes the model considered for the empirical analysis. Section 7 discusses the results, and Section 8 provides the conclusions.

2. Literature review

The literature underscores different large fire effects on average wages and incomes, as well as on the number of employees and income earners (i.e., the two components of overall output or income) depending on regional or local idiosyncrasies, such as population density and size, sectoral concentration and specialisation, spatial proximity to the fire, and other factors.

For instance, Kent *et al*. (2003) studied the effects of Colorado's 2002 Hayman wildfire on economic activity (wages, employment and retail sales) in three sectors and the overall economy during two periods: the two-month fire period and the two-month post-fire period. The geographic area taken into account included the four counties where the fire occurred (primary impact area) and thirteen counties bordering the primary impact area (secondary impact area). The authors used monthly data and linear regression models in order to estimate the economic activity that could be expected without the fire. To the extent that the difference between the estimated and the actual economic activity provides a measure of the fire impact, no strong evidence of positive or negative effects was detected.

Mosley *et al.* (2013) studied the effects of large wildfires on county-level employment and wages across eleven western states of the USA (346 large wildfires in 124 counties) from 2004 to 2008. The authors used quarterly data and linear regression models to estimate the effects of suppression spending on average employment and wage rate growth by controlling for past fires, population size and sectoral specialisation, and found that employment and average wages in the affected counties increased during the fire period. Neighbouring counties were also affected, exhibiting modest increases in both employment and average wages. However, the results varied depending on the characteristics of each county. Although the short-term effects of the wildfires were mostly positive, the medium-term effects (one to two years after the wildfire) tended to exhibit increased volatility, suggesting that wildfires amplified existing seasonal economic patterns.

Nielsen-Pincus *et al.* (2014) studied the impact of large wildfires that occurred in eleven western states of the USA during 2004-2008, on total and sectoral local employment growth. Monthly data from 413 counties were examined via linear regressions aiming to compare (a) the average change in the 24-month period that followed the end of wildfire suppression activities in the counties that experienced wildfires to (b) the average change in the counties that did not experience wildfires, while controlling for population size and the impact of the economic recession. The results suggest that, in counties with over 250,000 people, total employment growth was not significantly affected, and most sectoral effects were relatively small. However, leisure and hospitality services were associated with significant positive effects, while federal employment was associated with significant negative effects. On the other hand, in smaller counties, federal employment growth and employment growth in natural resources and mining increased significantly, while employment growth in leisure and hospitality, as well as manufacturing, shrank. As a rule, employment growth increased during the summer the fire occurred in and during autumn, returned to expected growth rates during winter, and dropped below expected growth rates during the next 12-18 months.

Nielsen-Pincus *et al.* (2013) also analysed quarterly county-level employment and earnings growth data from 122 western US counties affected by 346 wildfires during 2004-2008. By using generalised autoregressive conditional heteroskedasticity (GARCH) models and controlling for state growth and suppression spending, the authors found that employment and average wages increased during the quarter the wildfire occurred in. At the county-level the wildfire effects depended on population size and sectoral specialisation. For instance, counties with smaller population, specialising in recreation activities, exhibited an increase in employment growth, while service-specialised counties faced a decrease in employment growth. Moreover, employment and average wage volatility increased in the year following the fire as seasonal fluctuations were amplified. By and large, this loss-and-gain cycle was repeated for a second year to a lesser (and sometimes non-significant) magnitude.

Davis *et al*. (2014) used quarterly employment and wage data, as well as qualitative interview information, to study the effects of large wildfires that occurred in 2008 in Trinity County, California. The authors used a GARCH model to estimate employment and wage growth rates and volatility, and found that employment and wage growth rates increased more than expected during the fire period, but the effects varied across sectors. For instance, employment growth in the government sector was not affected, while average wage growth declined in the natural resource sector during wildfires. For the most part, the interviews confirmed the outcomes of the empirical analysis.

With these techniques and findings in mind –findings regarding diverse, predominantly short-term sectoral reactions and comparisons between places that were affected by wildfires and places that were not- we turn to the situation in the Peloponnese –a different part of the world– to study the economic impact of large fires through a different *lens*. The available data allow the consideration of several factors on an annual basis (from year to year), i.e., from a medium-term perspective, albeit lacking disaggregated qualitative and suppression/restoration cost elements, and monthly or quarterly frequency that permit the kind of empirical analysis carried out in the aforementioned literature. So, a new approach, one that exhibits high model fitness, has been devised.

Next, in order to better fill in the reader, we proceed with a description of the region and of the 2007 wildfires.

3. Background information

Situated at the edge of southeastern Europe, on the southernmost part of continental Greece, the Peloponnese (i.e., Greek for the "Island of (ancient king) Pelops") is home to 1.0 to 1.1 million people (according to the 2011 and the 2001 *Censuses*, respectively), spans an area of 21.4 thousand square kilometers (8.3 thousand square miles), is joined to the mainland by bridges at two places (one at the isthmus of Corinth, another at the Rion-Antirrion strait), and connected to nearby islands, the mainland, and Italy by ferries operating from several ports (north, south, east and west).

The terrain is dominated by high mountain-ranges, small valleys traversed by rivers, narrow coastal strips[1], and a very jagged coastline that extends for 1.4

1. Specifically: 48.3% of the area is situated at an elevation of over 800 meters (hereinafter referred to as *mountainous*), another 25.3% is below 800 meters featuring altitudinal differences of 300 meters or less (hereinafter referred to as *flat lowland*), while the remaining 26.4% is below 800 meters and features rather large altitudinal differences (hereinafter referred to as *intermediate/hilly terrains*).

thousand kilometers. These natural features greatly fragment the area into a host of tiny places[2] which, for the purpose of our analysis, will be studied at the disaggregated level in terms of 158 districts. Of these districts, the city of Patras (in the prefecture of Achaea) hosts 15.4 to 16.7% of the population (according to the 2001 and the 2011 *Censuses*, respectively); the other administrative (prefectural) seats, namely, Corinth (in Corinthia), Tripolis (in Arcadia), Kalamata (in Messenia), Sparta (in Laconia), Pirgos (in Ilis), Nafplion (in Argolis), and two towns with over 20-25 thousand people each (Aegion in Achaea, Argos in Argolis), host 19.6-21.0% of the population; another 149 smaller districts host over 62% of the population. For the sake of brevity, in the pages that follow, Corinth, Tripolis and Kalamata will be referred to as *districts A*, and the six towns listed after them as *districts B*.

At the time of the fires, Kalamata and Tripolis were linked to Corinth (and, from Corinth, to Greece's capital, Athens) via a newly-built multilane highway, route A7, while the city of Patras and Aegion were linked to Corinth via an older, narrower highway, route A8[3]. Two main roads near Tripolis connected Sparta and Argos-Nafplion to route A8, and two main roads at either end of route A8 linked up Pirgos and the UNESCO world heritage sites of ancient Olympia (in the west) and ancient Epidaurus (in the east). Of these, ancient Olympia was narrowly saved from the wildfires, as will be discussed below.

The wildfires consisted of a series of forest fires which erupted on (a) July 17th-18th 2007 in the north-eastern part of the region (namely, the prefecture of Corinthia), (b) July 23rd in the north (the prefecture of Achaea), and (c) August 24-25th along a belt which, by and large, runs from the north-eastern to the south-western part of the region (across the prefectures of Ilis, Arcadia, Messenia, Laconia). The flames spread and by the time they were put out (early September) had devastated 8% of the Peloponnese: 9% of its forest and bush territories, 8% of its agricultural land and 4% of its other (mostly inhabited) places (WWF, 2007; Gitas *et al*., 2008; Athanasiou and Xanthopoulos, 2010). See Map 1. At the same time, news of 49 people dead and images of the destruction sparked an outpour of sympathy for those who survived. To address the immediate needs of the latter, the Greek Government (2007) issued an instant relief allowance of 3,000 euro to each family affected, a higher allowance to large families, additional compensation for the immediate replacement of lost crops, animals, household and other belongings, for irreparable injuries, etc., while individuals, corporations, foreign governments also helped with reconstruction projects and donations.

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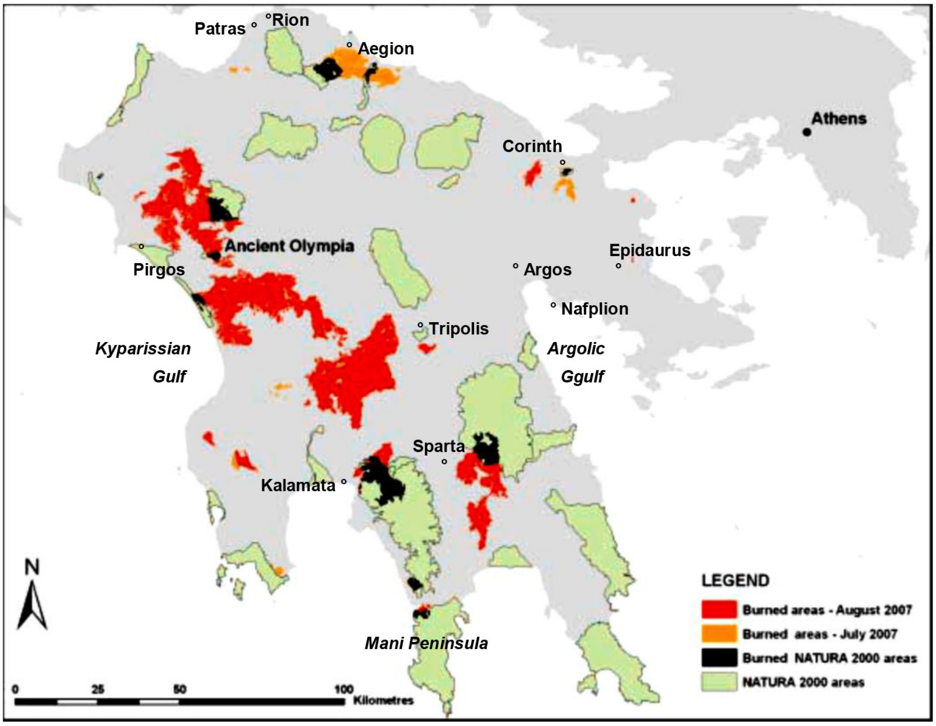
2. The same landscape gave birth to the patchwork of city-states and self-governing tribes of classical antiquity, and a mosaic of jurisdictional cantons in early modern times. Prodromidis (2010) finds very little inter-municipal commuting flows outside the urban centres of Patras, Corinth, Tripolis, Kalamata and their immediate surroundings.
3. The old A8 route was replaced by a newly-build multilane highway ten years later.

Map 1. The extent of the forest fires of July and August 2007 in the Peloponnese



Note: The Natura 2008 areas are breeding and resting sites for rare and threatened species.
Source: Gitas (2008), used with permission.

To put some of these figures in context: (a) In the course of the twenty years that preceded the fire, land used for agriculture had risen from 43% to 46% of the total, at the expense of forest and bush lands, which had shrunk from 11% to 10% and from 44% to 42%, respectively (WWF, 2012). (b) In the years after the fire, both grassland and cropland areas recovered quickly, while areas with widely spaced trees and forests recovered at a much slower pace (Rahman, 2014). (c) The effects on economic activity are not easily discernible at the (fairly aggregated) territorial level at which most statistics are usually collected, since: (i) The labor force and other state-of-the-economy surveys are conducted at a supra-prefectural scale. (ii) Six out of the region's seven prefectures comprise localities that lie in the actual path of the flames, as well as places that do not, while in all but one prefecture GDP figures per capita continued to grow and did not decrease prior to 2009 (i.e., the time that the Greek economy as a whole entered a recession period). See Table 1. (iii) The decline in population figures detected between the *Censuses* of 2001 and 2011, cannot be directly or exclusively attributed to the wildfires. Other prefectures in Greece, not affected by wildfires, experienced steeper reductions compared to the Peloponnese[4]. See Tables 2a and 2b. (iv) According to Greek Revenue Service (GRS) records, in the course of 2001-10, both the number of income-filers and the average level of declared personal income (the product of which comes to the overall amount of nominal personal income) in all seven prefectures increased.

Table 1. The evolution of per capita GDP in current prices in the Peloponnese (change %)

Territorial units	2000 -01	2001 -02	2002 -03	2003 -04	2004 -05	2005 -06	2006 -07	2007 -08	2008 -09	2009 -10	2010 -11*
Achaea pref.	6	8	9	9	2	14	4	1	-5	-1	-9
Argolid pref.	13	11	7	1	6	11	7	3	-1	-7	-10
Arcadia pref.	4	6	11	6	2	14	5	-3	5	-7	-10
Ilis prefecture	14	8	9	1	4	7	8	2	-3	2	-8
Korinthia pr.	8	0	8	1	3	7	6	7	-5	-7	-4
Laconia pref.	8	7	6	6	-1	8	8	1	1	-7	-6
Messenia pr.	8	3	10	4	7	10	7	2	0	-4	-4
Peloponnese	9	7	9	5	3	10	6	2	-3	-2	-8
Greece	7	6	10	8	3	9	7	4	-2	-4	-8

* Provisional data *Source:* Hellenic Statistical Authority (ELSTAT), own calculations.

Table 2a (left). Changes in population figures between 2001 and 2011 in the Peloponnese

Territorial units	Usual residents thousands	%
Achaea pref.	-9	-3
Argolis pref.	-5	-5
Arcadia pref.	-5	-5
Ilis pref.	-24	-13
Corinthia pref.	1	0
Laconia pref.	-4	-4
Messenia pref.	-7	-4
Peloponnese	-53	-5
Greece	-118	-1

Table 2b (right). Changes in population figures between 2001 and 2011 in other Greek prefectures with population drops in excess of 11 thousand people, i.e., 7%

Territorial units	Usual residents thousands	%
Athens (in Attiki)	-164	-6
Pireaus (in Attiki)	-30	-5
Serre (in Central Macedonia)	-18	-9
Fthiotis (in Central Greece)	-11	-7
Arta (in Epiros)	-6	-8
Kilkis (in Central Macedonia)	-6	-7

Source: ELSTAT, own calculations.

4. In the ten-year period between the censuses, the fire-stricken Ilis, in the west (where ancient Olympia is located), experienced a 13% reduction in population (the highest in the country), neighboring Arcadia (in its east), the fire-free Argolis (in the east), and the Peloponnese as a whole experienced a 5% reduction. At the same time, in considerable distance north of the Peloponnese, the prefectures of Arta, Fthiotis (where a minor fire was reported on August 26th), Kilkis and Serre experienced reductions in excess of 7%.

4. Available data

To properly investigate the matter, one needs disaggregated observations. To the extent that the international scholarship, by and large, attributes the effects of large wildfires on (a) people's average wages or incomes, and (b) employment or earner numbers to the size of local population and workforce-related features, merging the GRS and *Census* data provides a solution. The former data contain information on the evolution of the two components of personal income, number of income-filers and average income (i.e., the likely regressands), are annual, solicited at the postal-district level, and are associated with minor policy shifts in personal income taxation rules[5]. The latter data contain information on the terrain (flat/lowland, mountainous, intermediate/hilly), the demographics, the concentration and distribution of the population, as well as the sectoral composition of the workforce (i.e., the likely regressors), are organised in terms of rural and urban municipalities[6], and are updated every ten years.

Matching up the two datasets yields 158 amalgamated spatial units or districts[7]. Of these districts, 69 were damaged, either partially or completely, in the 2007 wildfires. This means that 69 of the 158 observations concern either damaged areas (communities) or damaged and undamaged communities, while the remaining observations concern undamaged communities only. Each community is either urban (containing one or more points with concentrations of 2,000 people or more) or rural (consisting of smaller or no such concentration points), situated either at high or low or intermediate elevation, and features an initial male and female workforce with a specific sectoral composition (for instance, 40 men and 25 women involved in agriculture, 30 men and 10 women involved in manufacturing, and so on, at the time of the 2001 *Census*).

The basic features of the damaged and undamaged districts in terms of (a) their natural settings, initial population densities and working population compositions, as well as (b) their initial working population and income-filing population numbers, and the evolution of their income-filing population numbers and of average incomes over time are provided in Tables 3a and 3b, respectively.

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5. Mainly involving an increase of threshold exemptions in 2004. Ceteris paribus, a rise in income-exemption levels ought to affect a fall in the number of income-filers and declared incomes, and may explain the findings regarding 2004, reported in the second paragraph of Section 7.
6. The city of Patras and the town of Kalamata contain several postal districts. Each of the remaining postal districts consists of rural communities (each with a population of fewer than two thousand inhabitants) and/or urban municipalities.
7. Securing a fair number of observations across space allows for a more thorough empirical examination in terms of explanatory variables within an acceptable range of degrees of freedom (e.g., Tabachnick and Fidell, 1989; Bartlett *et al.*, 2001).

Table 3a. Districts in the Peloponnese, damaged and undamaged by the 2007-fires, in terms of their natural setting features, initial (2001 census) population density, gender and sectoral working population distribution vectors (in shares, %)

	Damaged districts: 69					Undamaged districts: 89				
	Min	Max	Median	Mean	St.Dev.	Min	Max	Median	Mean	St.Dev.
Geography (consists of)										
Flat lowlands	0	100	**77**	**59**	40	0	100	**4**	**28**	38
Hilly country	0	92	**8**	**19**	24	0	100	**11**	**29**	36
Mountainous	0	100	**6**	**23**	34	0	100	**31**	**44**	43
Pop. density										
Urban	0	100	**40**	**36**	36	0	66	**0**	**4**	14
Rural	0	100	**60**	**64**	36	34	100	**100**	**96**	14
Working pop.										
Male	59	76	66	66	3	53	78	66	67	5
Female	24	41	34	34	3	22	46	34	33	5
Agriculture,etc.	2	73	42	42	17	4	82	45	45	17
Fishing, etc.	0	13	0	0	2	0	10	0	1	2
Mining, etc.	0	2	0	0	0	0	3	0	0	0
Manufacturing	0	19	4	6	3	2	19	5	5	3
Electricity, etc.	0	26	1	2	4	0	19	0	1	2
Construction	2	18	8	8	3	2	27	8	9	5
Trade, etc.	3	20	9	9	4	2	17	7	7	3
Hotels, etc.	1	11	4	4	2	1	27	4	5	4
Transport, etc	1	9	4	4	2	1	12	5	5	2
Financ. interm.	0	3	1	1	1	0	3	1	1	1
Real estate, etc.	1	6	3	3	1	0	6	2	2	1
Public admin.	1	12	5	5	2	1	14	5	5	3
Education	0	3	10	4	2	0	3	9	3	2
Health, etc.	0	8	2	2	1	0	7	2	2	1
Other services	0	5	2	2	1	0	9	2	2	1
Household act.	0	4	1	1	1	0	3	1	1	1
Extra-territorial	0	0	0	0	0	0	0	0	0	0
Unknown act.	0	18	5	6	4	1	35	5	6	5
Part damaged in 2007	0.4	100.0	7.8	25.5	31.6					

Note: **Bold** denotes cases in which the two groups on average (median compared to median, mean compared to mean) vary by 3 percentage points or more.

Source: See Table 1.

Table 3b. The 158 districts of the Peloponnese in terms of their initial working population and income-filing population sizes, and the evolution of their income-filing population figures during 2001-10

Number of districts		Initial size (2001) of the working pop. average (1)	income-filing population average (2)	min (3)	max (4)	Five patterns of change in terms of income-filing pop. 2001-4 (5)	2004-7 (6)	2007-10 (7)
1	not damaged	3,630	4,811	4,811	4,811	↓	↓	↓
54	not damaged [i]	1,705	1,781	103	17,130[a]	↑	↓	↓
28	damaged [ii]	4,240	4,598	88	19,241[b]	↑	↓	↓
23	not damaged [iii]	846	855	36	3,865	↑	↓	↑
3	not damaged [iv]	538	750	122	1,409	↑	↓	–
21	damaged [v]	1,023	1,190	84	4,967	↑	↓	↑
5	not damaged	628	620	231	1,558	↑	↑	↓
11	damaged [vi]	709	674	55	2,135	↑	↑	↓
3	not damaged	1,667	1,402	832	2,227	↑	↑	↑
9	damaged	14,816	17,684	496	95,123[c]	↑	↑	↑

Notes: Drops in average incomes were witnessed only in 2001-4, involving, in the cases under item (i) 13 districts, (ii) 7 districts, (iii) 6 districts, (iv) 1 district, (v) 3 districts, (vi) 3 districts.

The largest districts in terms of initial income-filing population figures were as follows: Under item (a) Tripolis and 2nd largest: Sparta (12,845). Under item (b) Argos and 2nd largest: Pirgos (17,661). Under item (c) Patras and 2nd largest: Kalamata (31,392).

Sources: ELSTAT (column 1), GRS (other columns), own calculations.

The numbers suggest that: Between 2004 and 2007 (and/or between 2007 and 2010), most districts with communities that would be (or were) damaged in the 2007 wildfires experienced reductions in their income-filing population figures. However, many more non-damaged districts exhibited the same pattern. One third of damaged districts and one fifth to a quarter of districts that were not damaged experienced reductions in average income levels between 2001 and 2004 (i.e., some time prior to the fires), but not after that. On average, the districts damaged in the fires were more urban and situated at a lower elevation compared to those that were not damaged in the fires; however, their initial working population structure was quite similar in terms of gender composition and sectoral orientation, consisting mostly of men (66-67%), engaging primarily in agriculture (42-45%).

Since this does not say much about the sources of income variation in the two sets of districts over time (before and after the fires), we turn to econometrics through which we estimate/isolate, year after year, the increased or reduced impact of every aspect for which data exist. Basically, we consider and compare the effects of all possible combinations of (a) urban or rural, (b) lowland, mountainous or intermediate (c) male or female, (d) NACE 1.1 sectoral workforce inputs[8], and (e) damaged or undamaged areas. We call these *elements z* for short. The same or similar elements are routinely used as explanatory variables (regressors) in econometric analyses of personal earnings, average household incomes, per capita GDP, median household income growth and the like (e.g., Schofield, 1975; Miles, 1997; Bhatta and Lobo, 2000; Aronson *et al*., 2001; Prodromidis, 2006; Gebremariam *et al*., 2012; the sources cited therein).

5. Methodological issues

Understandably, the presence of regressors that do not vary over time limits the ability to carry out typical panel data analyses on either component of personal income in a fixed effects and (more or less) in a random effects setting[9]. In particular, the fixed effects analysis cannot be carried out (consequently, the Hausman test cannot be performed) and the random effects analysis can only be carried out in a very limited form (see Appendix 1). At the same time, the attempt to explain (analyse) the two regressands and their annual changes over time via pooled longitudinal and cross-sectional data in a difference-in-difference mindset, with the aim of estimating (comparing) the autonomous and time-variant effects in the fire-damaged and other areas before and after the 2007 fires, turns out to produce equivalent or inferior fits compared to cases in which the regime-switching event occurs a year or two earlier: in 2005 or 2006 (see Appendix 2). With these limitations and thoughts in mind, and recognising the need to keep an eye on the changing impact of the factors considered, we turn to a series of cross-sectional analyses (regressions) –one for every year for the period under examination– and compare the parameters estimated from one year to the next. Thus, we are able to work out each factor's impact over time in damaged and undamaged communities, and identify shifts before or after the fires. As we shall see in the analysis, all regressions exhibit high levels of econometric model fitness.

8. The *Nomenclature statistique des Activités économiques dans la Communauté Européenne* 1.1 is the four-level sectoral classification of economic activities that was applicable in the European Union at the time the 2001 census results were published. The categories were: (a) agriculture, hunting and forestry; (b) fishing; (c) mining and quarrying; (d) manufacturing; (e) electricity, gas and water supply; (f) construction; (g) trade (wholesale and retail) and repair of motor vehicles, motorcycles and personal and household goods; (h) hotels and restaurants; (i) transport, storage and communication; (j) financial intermediation; (k) real estate, renting and business activities; (l) public administration, defense, and compulsory social security; (m) education; (n) health and social work; (o) other community, social and personal service activities; (p) activities of households (e.g., domestic staff); (q) extra-territorial organisations and bodies. For the purpose of the analysis, this means 18 sectors plus vague responses.

9. The regressands (i.e., the number of income-filers and average income) are in time-series form, while the regressors (altitudinal features, spatial dummies, the initial population and workforce concentration characteristics) are not.

The initial composition of the working population, in terms of demographic, sectoral, residential (in damaged or undamaged communities) and other characteristics, is provided in Table 4 (starting from the top white part, moving to the right, then the bottom light gray part and moving to the right):

1) 1.21% consisted of (i) men engaging in manufacturing activities and living in (subsequently) damaged urban hilly areas; or engaging in construction activities and living in undamaged urban hilly areas; or engaging in agriculture, hunting, forestry, electricity, gas and water supply activities and living in undamaged rural lowland areas; or engaging in financial intermediation and living in undamaged rural hilly areas; (ii) women engaging in household activities and living in (subsequently) damaged rural hilly areas, or engaging in hotel and restaurant activities and living in undamaged urban lowland areas; and (iii) men and women living in mountainous urban areas: All attached to (or being part of) Patras or the region's other main towns mentioned in Section 3.
2) 40.79% consisted of the remaining population living in the undamaged areas of Patras or the other main towns (*districts A* and *B*).
3) 5.93% consisted of the remaining population living in the damaged areas of Patras or the other main towns.
4) 2.82% consisted of the counterparts of those listed under items (i)-(iii), living in the rest of the Peloponnese; 41.37% consisted of the remaining population living in the undamaged areas of the rest of the Peloponnese; and 7.88% consisted of the remaining population living in the damaged areas of the rest of the Peloponnese. The rest of the Peloponnese comprises: (a) five districts near Patras, Corinth, and between Tripolis and Kalamata (*districts C*, hereinafter)[10], (b) 33 districts forming a belt around Patras and the main towns (*districts D*)[11], (c) eleven districts forming clusters or strings of localities by the south Kyparissian Gulf, the west Argolic Gulf, and on the Mani Peninsula[12], which exhibited differentiated effects compared to other places in the empirical analysis (*districts E*), and (d) the remaining districts.

10. Namely, Vrachneika (in Achaia), Ancient Corinth, Isthmia (in Corinthia), Megalopolis, Dirachion-Leontation (in Arcadia).
11. These include ancient Epidaurus, Ligourion, Nea Kios, Tolon (in Argolis), Athikia, Kiaton, Likoporia, Perigialion, Pitsa, Sofikon, Velon, Xilokastron, Zevgolation (in Corinthia), Akrata, Diakopton, Kamares, Kalavrita, Kato Achaea, Sageika (in Achaea), Varda (in Ilis and Achaea), Amalias, ancient Olympia, Avdravis, Katakolon, Kilini, Lehena, Vartholomion (in Ilis), Dimitsana, Eleohorion, Isaris, Lagkadia, Stemnitsa, Vitina (in Arcadia).
12. The first involves Zaharo (in Ilis), Dorion, Gargaliani, Kiparissia and Psarion (in Messenia). The second consists of Astros-by-the-sea (in Arcadia). The third involves Areopolis, Gerolimen, Githion, Kotronas and Pirgos-by-Diros (in Laconia).

Table 4. The distribution (in %) of the initial (2001 census) workforce in the Peloponnese across the sectors and types of districts which appear to affect the fitness of the income-filing population and average income functions the most

	male workforce (wf) in					female wf in		male and female wf in			
	Agricul. at RF areas	Constr. at UH areas	Manuf. at UHD areas	Finan. at RH areas	Electic. etc. at RF areas	Hotels etc. at UF ar.	Hholds at UH areas	UM not D areas	UM D areas	elsewhere not D areas	elsewhere D areas
1. Patras	0.01	0.00	0.00	0.00	0.00	0.21	0.00	0.00	0.00	18.58	0.00
2. Towns along route A7 [a]	0.08	0.00	0.16	0.00	0.00	0.08	0.00	0.00	0.00	7.44	5.13
3. Other main towns [b]	0.49	0.00	0.00	0.00	0.01	0.15	0.00	0.00	0.00	14.77	0.80
4. Five districts near 1 and 2[c]	0.02	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	1.20	0.58
5. Belt adjacent to 1-4 [d] and three formations [e]	0.89	0.03	0.00	0.01	0.01	0.08	0.00	0.00	0.10	15.88	3.48
6. Other places	1.35	0.12	0.02	0.01	0.03	0.02	0.00	0.15	0.00	24.29	3.81

Notes: R: rural. U: urban. F: flat lowlands. H: hilly/intermediate terrain. M: mountainous terrain. D: damaged in the 2007 fires. A district may consist of R and/or U, F and/or H and/or M, D or other parts.

[a] Districts A (see Section 3). [b] Districts B (see Section 3). [c] Districts C (see fn. 10) .
[d] Districts D (see fn. 11). [e] Districts E (see fn. 12).

Source: See Table 1.

6. Modeling strategy

As already mentioned, the empirical analysis is carried out through a series of cross-sectional analyses due to the nature of the available data. In this, the two components of people's overall personal income, Y, reported across districts, are explained in terms of the factors mentioned at the end of Section 4 (see *elements z*). Besides these factors, the only other piece of information that is both available at the disaggregated level and suited to explain Y, takes the form of categorical (dummy) variables that capture the administrative and transportational nodes discussed in Section 3 or denote the spatial formations mentioned at the end of Section 5 (*element x*, hereinafter).

For the two sets of *elements* to satisfy a basic assumption regarding regressor independence, x and z are made orthogonal to each other. (See Zeller, 1974; Gujarati, 1995.) In essence, instead of regressing Y or its components on arguments x and z, z (the continuous portion) is first regressed on x (the categorical portion), then a z' is predicted, and a *residual* $\zeta = z - z'$, orthogonal (hence, entirely independent) to x, is estimated. Thus, Y or its components may be explained in terms of x and ζ at any given time, *t*, *t+1*, *t+2* etc[13]. In shorthand functional form notation:

$$Y_{t+k} = Y(x_t, \zeta_t) \tag{1}$$

where *t* refers to 2001 and *k* takes values from 0 to 9.

To the extent a district's *Y* is equivalent to the product of its mean, *y*, times the size of the income-declaring population, *n*, it is probably reasonable to study the initial and subsequent levels of *y* and *n* both before and after the fires, in terms of the same explanatory variables:

$$n_{t+k} = n(x_t, \zeta_t), \tag{2}$$
$$y_{t+k} = y(x_t, \zeta_t). \tag{3}$$

Thus, the two components of Y observed over the years will be studied/explained in terms of vectors pertaining to 2001. This means that each of the *n*'s and *y*'s observed in 2001, 2002, and the other years of the period under examination, will be regressed separately on groupings of *x* and *ζ* dating to 2001. In theory, explaining phenomena in terms of earlier data is quite acceptable. Here, the specification (i.e., the grouping of elements *x* and ζ) in each of the two equations will be allowed to vary so as to best fit the data.

For the sake of explaining the respective results of the two expressions in terms of people and euro, rather than in terms of elasticities and semi-elasticities, a simple linear arrangement will be adopted; and, in order to address heteroscedasticity concerns, the econometric analysis is performed within a robust standard errors framework.

In the interest of brevity, only explanatory variables (regressors or characteristics) that improve the econometric goodness of fit the most are taken into account. The other variables or characteristics generally exhibit lesser (similar to one other) effects; so, with negligible loss of information, the following are grouped together: spatial categorical variables with spatial categorical variables, workforce composition variables with workforce composition variables. Consequently, the expression (regression) concerning the size of the income-filing population, *n*, is found to depend very much on ten regressors, and the expression regarding the average level of personal income, *y*, if found to depend on eleven regressors. One regressor (namely, the initial number of men living in undamaged, hilly, urban areas who engaged in construction) is common in both regressions, while the other regressors vary.

13. The technique is employed in a multitude of analyses across disciplines (e.g., Kentor, 2001; Cutright *et al.*, 2006, 2007; Jaeger, 2010; Cohen-Goldberg, 2012; Riley, 2012; Bradford and Stoner, 2014). In this setting, when the empirical analysis is carried out, only workforce regressors are modestly correlated with each other (|r|<15% in Table 5, |r|<25% in Table 6). In Table 6, elements x_t correspond to the top four regressors, and element ζ_t to the other regressors. In Table 5, elements x_t correspond to the top three regressors, and element ζ_t to the other regressors. (Due to the presence of high r's between the ultimate and penultimate regressors and the other five workforce regressors considered in Table 5, the ultimate and penultimate regressors are made orthogonal to the other five workforce regressors as well).

It turns out that the factors which are important for 2001 are also important for 2002, 2003, and following years (i.e., both before and after the fires); besides, regression fitness does not change much from year to year: all R^2s are in the 98-99% range, and adjusted R^2s are lower by no more than a decimal. The estimated coefficients and p-values are provided in Tables 5 and 6; however, to the extent the data involve the entire population rather than a sample of districts, the need to turn to p-values and estimate measures of the recovered coefficients' statistical closeness to the parameters observed in the population probably diminishes.

Table 5. Econometric regressions with robust standard errors of each year's number of income-filers in the 158 districts of the Peloponnese, 2001-10

	2001		**2002**		**2003**		**2004**		**2005**	
Explanatory variables	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values
City of Patras	93,711	0.00	95,331	0.00	96,240	0.00	101,467	0.00	100,416	0.00
Main towns	16,711	0.00	17,151	0.00	17,395	0.00	19,257	0.00	18,705	0.00
Other places	1,412	0.00	1,452	0.00	1,457	0.00	1,655	0.00	1,593	0.00
Initial male pop. in										
• Agricult. at RFN	4	0.00	4	0.00	4	0.00	5	0.00	5	0.00
• Constr. at UHN	15	0.00	16	0.00	16	0.00	17	0.00	17	0.00
• Manufact. at UHD	13	0.00	13	0.00	13	0.00	15	0.00	15	0.00
• Financial at RHN	195	0.00	205	0.00	208	0.00	234	0.00	224	0.00
Initial female pop. in										
• Hotels etc. at UFN	49	0.00	51	0.00	52	0.00	57	0.00	56	0.00
Other pop. in										
• Undamaged areas	1	0.00	1	0.00	1	0.00	1	0.00	1	0.00
• Damaged areas	1	0.00	1	0.00	1	0.00	1	0.00	1	0.00
Model fit (R^2)	*99.7%*		*99.7%*		*99.7%*		*99.7%*		*99.7%*	

	2006		**2007**		**2008**		**2009**		**2010**	
Explanatory variables	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values
City of Patras	101,567	0.00	102,470	0.00	103,395	0.00	104,067	0.00	103,645	0.00
Main towns	18,863	0.00	19,119	0.00	19,254	0.00	19,345	0.00	19,265	0.00
Other places	1,594	0.00	1,608	0.00	1,610	0.00	1,613	0.00	1,601	0.00
Initial male pop. in										
• Agricult. at RFN	5	0.00	5	0.00	5	0.00	5	0.00	4	0.00
• Constr. at UHN	17	0.00	17	0.00	17	0.00	17	0.00	17	0.00
• Manufact. at UHD	15	0.00	15	0.00	15	0.00	16	0.00	16	0.00
• Financial at RHN	226	0.00	232	0.00	235	0.00	239	0.00	234	0.00
Initial female pop. in										
• Hotels etc. at UFN	57	0.00	57	0.00	57	0.00	58	0.00	58	0.00
Other pop. in										
• Undamaged areas	1	0.00	1	0.00	1	0.00	1	0.00	1	0.00
• Damaged areas	1	0.00	1	0.00	1	0.00	1	0.00	1	0.00
Model fit (R^2)	*99.7%*		*99.7%*		*99.7%*		*99.7%*		*99.7%*	

Notes: R: rural. U: urban. F: flat lowlands. H: hilly/intermediate terrain. M: mountainous terrain. D: damaged in the 2007 fires. N: not damaged in the 2007 fires. Districts are heterogeneous: with R and/or U, F and/or H and/or M, D or other parts. All parts were considered. The 4th-8th regressors are orthogonal to the top three regressors, and the two bottom regressors are orthogonal to all previous regressors. Since categorical variables were specified for all places, the regressions were estimated without a constant. Ramsey's regression equation specification error test suggests an absence of relevant variables at the 1% error probability level.

Table 6. Econometric regressions with robust standard errors of each year's average income (in euro) in the 158 districts of the Peloponnese, 2001-10

	2001		**2002**		**2003**		**2004**		**2005**	
Explanatory variables	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values
Patras; districts A, C	12,364	0.000	13,228	0.000	13,771	0.000	13,822	0.000	15,181	0.000
Districts B	10,887	0.000	11,535	0.000	11,795	0.000	11,670	0.000	12,937	0.000
Districts D, E	9,429	0.000	9,946	0.000	10,309	0.000	10,294	0.000	11,456	0.000
Other places	7,512	0.000	7,964	0.000	8,173	0.000	7,878	0.000	8,941	0.000
Male pop. in										
• Electr. etc.at RFN	92	0.000	95	0.000	106	0.000	99	0.000	108	0.000
• Constr. at UHN	9	0.000	9	0.000	9	0.000	10	0.000	11	0.000
Female pop. in										
• Households at UFD	247	0.000	265	0.000	285	0.000	317	0.000	343	0.000
Pop. in UMN	-0	0.099	-0	0.099	-0	0.055	-1	0.000	-0	0.001
in UMD	0	0.003	1	0.000	0	0.040	-0	0.009	-0	0.543
elsewhere N	0	0.958	-0	0.996	-0	0.760	0	0.679	-0	0.899
elsewhere D	0	0.181	0	0.350	0	0.614	0	0.650	0	0.912
Model fit (R^2)	*99.1%*		*99.1%*		*99.1%*		*98.9%*		*99.1%*	

	2006		**2007**		**2008**		**2009**		**2010**	
Explanatory variables	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values
Patras; districts A, C	16,170	0.000	17,308	0.000	18,113	0.000	18,684	0.000	20,732	0.000
Districts B	13,791	0.000	14,707	0.000	15,328	0.000	15,677	0.000	18,116	0.000
Districts D, E	12,157	0.000	13,074	0.000	13,503	0.000	13,815	0.000	16,055	0.000
Other places	9,547	0.000	10,268	0.000	10,812	0.000	11,040	0.000	13,729	0.000
Male pop. in										
• Electr. etc. at RFN	117	0.000	121	0.000	119	0.000	124	0.000	111	0.000
• Constr. at UHN	12	0.000	12	0.000	12	0.000	11	0.000	12	0.000
Female pop. in										
• Households at UFD	355	0.000	390	0.000	423	0.000	448	0.000	400	0.000
Pop. in UMN	-0	0.180	-0	0.004	-0	0.101	-0	0.248	-0	0.014
in UMD	0	0.961	-0	0.364	0	0.136	0	0.126	0	0.797
elsewhere N	-0	0.671	-0	0.528	-0	0.428	-0	0.289	-0	0.353
elsewhere D	0	0.999	-0	0.846	-0	0.657	-0	0.528	-0	0.917
Model fit (R^2)	*99.1%*		*99.1%*		*99.1%*		*99.1%*		*99.5%*	

Notes: R: rural. U: urban. F: flat lowlands. H: hilly/intermediate terrain. M: mountainous terrain. D: damaged in the 2007 fires. N: not damaged in the 2007 fires. Districts are heterogeneous: with R and/or U, F and/or H and/or M, D or other parts. All parts were considered. The 5th-11th regressors are orthogonal to the top four regressors. Since categorical variables were specified for all places, the regressions were estimated without a constant. Ramsey's regression equation specification error test suggest the equations have no omitted variables at the 1% error probability level.

7. Empirical findings

It appears that a small number of factors suffices to explain a good part of the overall variation year after year, while there is no inordinate change in the sizes of the estimated coefficients for the years following the fires (2007-08 and 2008-09), particularly in areas damaged by the fires. (See also Table 7: it consolidates the information of Tables 5 and 6.)

Table 7. The results of Tables 5 and 6 in terms of changes detected from year to year, 2001-10

Income-filer numbers	01-02	02-03	03-04	04-05	05-06	06-07	07-08	08-09	09-10
City of Patras	1,620	909	5,227	-1,051	1,151	903	925	672	-422
Main towns	440	244	1,862	-552	158	256	135	91	-80
Other places	40	5	198	-62	1	14	2	3	-12
Initial male pop. in									
• Agricult. at RFN	0	0	1	0	0	0	0	0	-1
• Constr. at UHN	1	0	1	0	0	0	0	0	0
• Manufact. at UHD	0	0	2	0	0	0	0	1	0
• Financial at RHN	10	3	26	-10	2	6	3	4	-5
Initial female pop. in									
• Hotels etc. at UFN	2	1	5	-1	1	0	0	1	0
Other pop. in									
• Undamaged areas	0	0	0	0	0	0	0	0	0
• Damaged areas	0	0	0	0	0	0	0	0	0
Average incomes	**01-02**	**02-03**	**03-04**	**04-05**	**05-06**	**06-07**	**07-08**	**08-09**	**09-10**
Patras; districts A, C	864	543	51	1,359	989	1,138	805	571	2,048
Districts B	648	260	-125	1,267	854	916	621	349	2,439
Districts D, E	517	363	-15	1,162	701	917	429	312	2,240
Other places	452	209	-295	1,063	606	721	544	228	2,689
Initial male pop. in									
• Electr. etc. at RFN	3	11	-7	9	9	4	-2	5	-13
• Constr. at UHN	0	0	1	1	1	0	0	-1	1
Initial female pop. in									
• Households at UFD	18	20	32	26	12	35	33	25	-48
Other pop.									
• in UMN	*0*	*0*	-1	1	*0*	0	*0*	*0*	*0*
• in UMD	1	*-1*	0	*0*	*0*	*0*	*0*	*0*	*0*
• elsewhere N	*0*	*0*	*0*	*0*	*0*	*0*	*0*	*0*	*0*
• elsewhere D	*0*	*0*	*0*	*0*	*0*	*0*	*0*	*0*	*0*

Notes: R: rural. U: urban. F: flat lowlands. H: hilly/intermediate terrain. M: mountainous terrain. D: damaged in the 2007 fires. N: not damaged in the 2007 fires. Shades denote reductions in the coefficients of Tables 5 and 6 (from one year to the next). Italics denote effects associated with p-values > 1% in Tables 5 and 6.

Evidently, the populous and densely populated (costly-to-live-in) areas are associated with higher tax-filer numbers and incomes compared to other places, while Patras, the main towns and smaller communities grew during 2001-04 and 2005-09 in terms of income-filer numbers, and more so (in both periods) the rural areas in the hills that initially possessed a relatively large male workforce involved in financial intermediation and were not damaged in the 2007 wildfires. (Hence, the trend probably pre-existed). In addition, average incomes generally increased over time across the region, though small reductions were generally observed:

1) In 2004 in all places but Patras, the three towns situated along *route A7* (*districts A*), and five neighbouring districts (*districts C*) (This was prior to the fires).
2) In a number of communities that were not damaged in the fires: (a) in 2004, 2008 (negligibly) and 2010 in lowland rural areas that initially possessed a relatively large male workforce involved in electricity-gas-water supply, (b) in 2009 (negligibly) in hilly urban areas that initially possessed a relatively large male workforce involved in construction, and (c) in 2004 (negligibly) in mountainous urban areas that initially possessed a relatively large workforce.
3) In a number of communities that were damaged in the 2007 fires: (a) negligibly in 2003-4 (before the fires) in mountainous urban areas that initially possessed relatively large workforces, and (b) in 2010 in lowland urban areas that initially possessed a relatively large female workforce involved in household activities.

In short, there is no apparent evidence of change in income-filer numbers or average income figures reported in the damaged and undamaged areas that may be attributed to the fires.

8. Conclusions

The article engages in a series of cross-sectional econometric analyses to (a) study the evolution of annual personal income in the Peloponnese, in southern Greece, at the disaggregated level (across 158 districts), year after year from 2001 to 2010, and (b) to comprehend the medium-term economic effects of the 2007 wildfires on income-filer figures and average incomes in the burned areas and in the other (unburned) areas in the region operating under the same climatic, commercial or other general conditions. To the best of our knowledge, the incident has not been empirically examined yet. However, the pooled and panel set analyses do not fare well with the data available, while the difference-in-difference approach may be misleading.

The main limitation is that two components of personal income observed over the years (i.e. the number of income-filers and the average level of declared incomes) may only be studied in terms of altitudinal features and other spatial dummies, as well as population and workforce concentration features that do not vary over time but date to 2001 (i.e., are explained in connection with hard-to-change natural settings and the initial population densities and working population compositions of damaged and undamaged areas). Explaining phenomena in terms of earlier data is acceptable. In addition, to isolate the information and effects contained in the categorical variables from the information and effects contained in the other available explanatory variables, the two sets of regressors are made independent from each other. Thus, each factor's impact over time, especially before and after the fires, in both damaged and undamaged communities, is worked out. Interestingly, all regressions are associated with high econometric fitness; and the factors that turn out as important for 2001 are also important for 2002, 2003, and the following years (i.e., both before and after the 2007 fires).

Consequently, the article finds that a small number of spatial features and initial workforce composition elements explain, by and large, the variation observed in the beginning of the period and throughout the decade. In addition, there is no strong evidence of change in income-filer numbers or average income figures reported in damaged and undamaged areas on account of the fires. So, it seems and that either due to the operation of the economy, or the type of damage caused, or government intervention, or the medium-term nature of the data or other reasons, no inordinate drop or rise in average income figures or income-filer numbers is detected in the fires' aftermath, particularly in the communities damaged by the fires. Indeed, for the most part, the patterns observed predate the fires. For instance, throughout the period under consideration, the main urban centres linked via routes A7 and A8, and neighbouring localities, as well as the urban parts of the lowlands with a relatively large initial female population employed in household activities that were damaged in the 2007 fires, generally exhibited higher levels of average income compared to the rest of the region up to 2009.

Though the findings may hold some useful lessons for territorial development policy interventions, we do not press the analysis beyond 2010. By most macroeconomic accounts, by mid-2009, the international, economic and financial crisis reached Greece, adversely affecting two of the country's largest earners, namely, shipping and tourism (inherently sensitive to the contraction of international activity); and once austerity measures were adopted, the whole country entered into deeper recession, upsetting both threshold exemptions and the economic behavior and reactions of those involved.

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Appendix 1. Panel data analyses with fixed and random effects

As the available natural setting, gender, and sectoral information regarding the 158 districts does not vary over time, neither a fixed effects model nor a random effects model of average income and income-filer numbers with a forest-fire dummy is computable within a panel data analysis framework. So, in order to tell the difference before and after the fires, in both damaged and undamaged areas, the data is split into two, and one analysis is performed for the years before the fires, while the other analysis is performed for the years after the fires. All explanatory information is consolidated in two regressors and a trivariate version of a random effects model is estimated. The Breusch-Pagan Lagrange multiplier test justifies the use of random effects in all four parts of the analysis.

The coefficients obtained in the process are provided in Table A1, and suggest that in the post-fire period the estimated effects associated with the damaged areas are higher than the effects associated with the undamaged areas as compared to the pre-fire period. Though not a definitive proof of a reaction or a shift attributed to the fires, the findings are not incompatible with the argument.

Table A1. Random effects panel analyses with robust standard errors across 158 districts of the Peloponnese during 2001-10

	Figures declared before the fire (2001-06)				Figures declared after the fire (2007-10)			
	Number of income filers		Average income		Number of income filers		Average income	
Explanatory variables	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values	coeffi-cients	p-values
Constant	-320.28	0.00	9,090.09	0.00	-347.32	0.00	12,380.67	0.00
Initial working pop. in D areas	1.30	0.00	0.24	0.00	1.53	0.00	0.27	0.00
Initial working pop. in N areas	1.31	0.00	0.11	0.00	1.37	0.00	0.13	0.00
Overall R^2	*99.2%*		*16.5%*		*99.4%*		*13.6%*	
Wald X^2	*16,968.13*		*52,65*		*18,515.74*		*42.26*	
Prob. > Wald X^2	*0.00 %*		*0.00 %*		*0.00%*		*0.00%*	

Notes: Districts are heterogeneous: each containing parts (areas) damaged (D) and/or undamaged (N) in the 2007 fires. All parts were considered. Though reported, the Wald test is of limited use given that it is based on the population rather than a sample.

Appendix 2. Pooled data analyses focusing on the autonomous and time-variant effects of the fires in damaged and undamaged areas

The average local incomes, income-filer numbers and their annual changes are empirically estimated in the difference-in-difference setting described by Card and Krueger (1994), in order to ascertain the patterns before and after the fires in both damaged and undamaged districts across the Peloponnese, while controlling for natural setting, district-specific, gender and sectoral effects.

The recovered coefficients are provided in Table A2, and suggest the presence, initially (in 2001), of:

- A lesser autonomous and a larger time-variant effect on income filing population figures in the districts damaged during the 2007 fires compared to other districts (see col.1, lines 33-37). From 2007 onwards, a lesser autonomous effect in damaged communities (areas) compared to other communities in both damaged and undamaged districts, and a lesser time-variant effect in damaged districts compared to other districts (col.1, lines 38-43).
- A lesser autonomous and a larger -though decreasing- time-variant effect on average incomes in the districts damaged during the fires compared to other districts (col.2, lines 33-37). From 2007 onwards, a larger (lesser) autonomous effect in damaged communities compared to other communities in the same districts (undamaged districts), and a lesser time-variant effect in the damaged districts compared to other districts (col.2, lines 38-43).
- A larger autonomous and a lesser time-variant effect on the annual change of income filing population figures in the districts damaged during the fires compared to other districts (col.3, lines 33-37). From 2007 onwards, a larger autonomous effect in damaged communities compared to other communities in the same districts (col.3, lines 38-43).
- A lesser autonomous and a larger -though decreasing- time-variant effect on the annual change of average income figures in districts damaged during the 2007 fires compared to other districts (col. 4, lines 33-37). From 2007 onwards, a lesser autonomous effect in damaged communities compared to other communities in the same districts, and a lesser time-variant effect in damaged districts compared to other districts (col.4, lines 38-43).

However, the importance of the fires is tempered by the presence of superior econometric fits in regressions regarding annual changes in the number or income filers (NIF) and average incomes (AI), under slightly earlier regime switches: that is, if the categorical variable associated with the presumed regime switch were moved to 2006 (for NIF) or 2005 (for both NIF and AI). See Table A3.

Table A2. Econometric regressions with robust standard errors on the evolution of the income filing population and of average income in the 158 districts of the Peloponnese during 2001-10

	Explanatory variables	Number of income filers (NIF)		Average income (AI) in euro		Annual change in NIF		Annual change in AI	
		coefficients	p-values	coef.	p-values	coef.	p-values	coef.	p-values
1	Constant	2,595	0.00	6,465	0.00	-61	0.33	917	0.00
	Districts								
2	City of Patras	97,266	0.00			1,049	0.07		
3	Town of Kalamata	31,087	0.00			589	0.11		
4	Other main towns (population centers)	14,214	0.00			194	0.02		
5	Isthmia (in Corinthia)			4,697	0.00			297	0.19
6	Areopolis, Kotronas (in Laconia), Dorion (Messenia)			2,706	0.00			69	0.58
7	Twenty three high income localities[a]			1,861	0.00			64	0.11
8	Other places (reference)								
	Initial working population (WP)								
	Distribution across space (%)								
9	In urban and flatland areas (reference)								
10	In rural areas	-23	0.00	-7	0.00	-0	0.04	-0	0.66
11	In hilly areas	-0	0.71	-1	0.38	-0	0.84	0	0.95
12	In mountainous areas	-8	0.00	-9	0.00	-0	0.20	0	0.71
	Distribution by gender (%)								
13	Females (reference)								
14	Males	5	0.58	-36	0.00	-0	0.53	-3	0.56
	Distribution across sectors (%)								
15	In agriculture-hunting-forestry (reference)								
16	In fishing	45	0.00	97	0.00	0	0.87	15	0.12
17	In mining-quarrying	-328	0.07	639	0.00	-0	0.98	-6	0.93
18	In manufacturing	-55	0.00	140	0.00	-1	0.69	9	0.11
19	In electricity-gas-water supply	17	0.36	137	0.00	-0	0.90	14	0.06
20	In construction	54	0.00	42	0.00	1	0.45	-2	0.64
21	In trade, etc.	157	0.00	-6	0.70	1	0.43	-8	0.26
22	In hotels-restaurants	-46	0.00	54	0.00	-0	0.85	0	0.99
23	In transport-storage-communication	-4	0.86	58	0.00	-0	0.86	-3	0.76
24	In financial intermediation	260	0.00	136	0.00	4	0.57	28	0.32
25	In real estate-renting-business activities	-37	0.29	218	0.00	0	0.94	3	0.85
26	In public admin., etc.	-130	0.00	65	0.00	-2	0.33	-1	0.93
27	In education	69	0.02	237	0.00	1	0.76	14	0.28
28	In health-social work	-58	0.06	71	0.01	3	0.73	6	0.75
29	In other community-social-personal serv.	-96	0.01	133	0.00	-1	0.73	2	0.87
30	In household activities	-255	0.00	245	0.00	-5	0.33	-1	0.95
31	In extra-territorial bodies	3,913	0.05	2,802	0.06	21	0.92	137	0.86
32	In unknown activities	33	0.00	36	0.00	0	0.90	2	0.64
	Districts damaged in the fire								
33	Autonomous component	-85	0.63	-273	0.08	84	0.00	-114	0.08
34	Time from 2001 onwards (in years)	130	0.00	488	0.00	141	0.00	-406	0.00
35	Time squared					-29[b]	0.00	89[c]	0.00
	Other districts								
36	Time from 2001 onwards (in years)	40	0.12	279	0.00	147	0.00	-481	0.00
37	Time squared			16	0.01	-27	0.00	98	0.00
	From 2007 onwards in the damaged areas (DA)								
38	Autonomous component: District dummy × share of WP in DA	-3	0.67	16	0.07	2	0.00	-6	0.00
39	Time component	-2	0.58	-17	0.04	3	0.00	-19	0.00
40	Time component squared			5[d]	0.01			5[e]	0.00
	From 2007 onwards in the other areas								
41	Autonomous component	0	0.98	22	0.00	2	0.00	-5	0.00
42	Time component	-1	0.27	-20	0.00	2	0.00	-21	0.00
43	Time component squared			5[f]	0.00	-0	0.00	5[g]	0.00

Table A2 (Continued)

	Number of income filers (NIF)	Average income (AI) in euro	Annual change in NIF	AI
Observations	1,580	1,580	1,422	1,422
Model fit (R^2)	97.7%	90.2%	22.7%	62.7%

Notes:

[a] Drepanon, Nea Epidaurus (in Argolis), Dervenion, Hiliomodion, Kaliani, Klimention, Pitsa, Vrahation, Xilokastron (in Corinthia), Kalavrita, Kertezi, Sageika (in Achaea), Varda (in Achaea and Ilis), Lehena (in Ilis), Dimitsana, Kosmas, Megalopolis, Tripolis, Vitina (in Arcadia), Dorion, Kiparissia, Psarion (in Messenia), Gerolimen, Pirgos-by-Diros (in Laconia).
[b] Maximum 2.4 years. [c] Minimum at 2.3 years. [d] Minimum at 12.8 years. [e] Minimum at 2.0 years. [f] Minimum at 3.5 years. [g] Minimum at 1.9 years. The maxima and minima are calculated on the basis of the first and second order conditions of the respective functions with respect to time. In each function, the linear or non linear specification for time was considered and decided based on the best fit (adj R^2). Ramsey's regression equation specification error test suggests an absence of relevant of variables at the 1% error probability level in all regressions.

Table A3. The R^2s recovered from the expressions employed in Table A2 under hypothetical, slightly earlier regime switches

Year of presumed regime switch	Number of income filers (NIF)	Average income (AI) in euro	Annual change in NIF	AI
2005	97.7%	89.9%	31.4%	67.5%
2006	97.7%	89.9%	25.4%	60.4%
2007	97.7%	90.2%	22.7%	62.7%

ASECU South-Eastern Europe Journal of Economics 1 (2019) 33-55 SEEJE

CREDIT, HOUSE PRICES AND THE MACROECONOMY IN CYPRUS

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Abstract

This paper examines whether there is a link between house prices, credit and macroeconomic conditions in Cyprus, using a vector error correction model (VECM) and quarterly data from 2005Q4 to 2016Q4. Overall, the results suggest that a link does exist and that house prices have a bi-directional relationship with loans and unemployment rates. Macroeconomic conditions matter for Cyprus economy as an unexpected shock in unemployment has been found to have a persistent impact on all model variables. Interest rates have also been found to have an effect on wages and house prices.

JEL Classification: E44, E50, R20
Keywords: Credit, House Prices, VECM, Cyprus

All views expressed in this paper are strictly personal and do not necessarily reflect the views of the Central Bank of Cyprus or the Eurosystem. We are indebted to Marios Polemidiotis, George Georgiou and the Economic Analysis and Research Group at the Central Bank of Cyprus for their valuable comments and suggestions. All remaining errors and omissions are ours.

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1. Introduction

After the turn of the century, most developed economies experienced strong rates of money and credit growth associated with house price increases[1]. The link between the two became more pronounced by 2008, when the global financial crisis was fully blown. As the crisis evolved, the interaction between credit and house prices was already considered to be the root cause (Duca *et al.*, 2010).

In the case of Cyprus, a similar phenomenon had been observed during the period preceding the country's accession to the European Union in 2004. Fuelled by real estate demand and strong domestic credit growth, property prices and construction activity soared. During the 2006-2008 period, credit growth in Cyprus surged further and intensified following the irrevocable fixing of the Cyprus pound against the euro in July 2007 at its ERM II central parity. The main factors behind this increase were the lower foreign exchange risk perceived and the anticipated convergence of interest rates towards euro area rates. The two factors generated a climate of euphoria in the real estate sector both for Cypriot and foreign buyers, which boosted demand and led to an escalation in property prices. In addition, an influx of foreign deposits, attracted by higher interest rates, compared with the rest of the euro area, further boosted credit growth, which, in its turn, was channelled to housing projects. Indeed, the sharp rise in property prices observed in Cyprus from 2006 onwards was among the highest across euro area countries (Argiridou-Dimitriou *et al*., Ch.6, 2012).

These developments raise an important question: are the developments observed in house prices, loans, unemployment and other macroeconomic variables caused by a common driving force, such as the economic cycle, or are they due to direct linkages between said variables? Furthermore, if a direct link exists, does it run from house prices to loans and to other macroeconomic variables, does causality flow the other way or is there evidence of the connection moving in more than one directions?

Building on the findings of relevant literature (see also next section), this paper examines, for the first time, the possible existence of such a link in Cyprus. In particular, by employing cointegration analysis and vector error correction models (VECM), we allow for the assessment of credit dynamics, house prices and macroeconomic developments on the island. Through the use of quarterly time series data from 2005Q4 to 2016Q4, findings show evidence of a bi-directional relationship between loans and house prices and between unemployment rates and house prices. In other words, changes in loan growth are foreseen to affect house prices, whilst changes in house prices will also have an impact on loan growth.

1. In this context, credit and loans are used interchangeably and refer to Monetary Financial Institutions (MFIs) lending to non-MFIs.

Furthermore, wages affect unemployment rates and vice versa. Macroeconomic conditions are extremely important for Cyprus economy, as an unexpected shock in unemployment has a persistent impact on all the variables employed in the model. Interest rates are also found to have an effect on wages and house prices, but a lower effect on other variables.

The remainder of the paper is structured as follows: Section 2 provides a short overview of relevant literature, with Section 3 describing the set of data employed. Section 4 describes the empirical methodology used; Section 5 presents the empirical results, impulse responses and variance decomposition, while Section 6 concludes.

2. Literature Overview

Existing literature has presented evidence that higher house prices induce homeowners to spend and borrow more, via a wealth effect, hence providing a link between house prices and credit growth (see, *inter alia*, Bostic *et al*., 2009; Dvornak and Kohler, 2007; Gimeno and Martinez-Carrascal, 2010; Gan, 2010). As recent literature has shown, increased spending promoted by rising house prices can have a significant effect on economic growth, even at the regional level (Miller *et al.*, 2010; Davis and Palumbo, 2001). Furthermore, as Iacoviello and Minetti (2008) point out, monetary policy transmission is also affected by the housing finance system through the bank balance sheet channel.

In general, other researchers have taken up the task of exploring the linkage between house prices, credit and the macroeconomy for a number of countries, with most studies using macro-level data. In the study most closely related to this one, Goodhart and Hofmann (2008) examine linkages between money, credit, house prices and economic activity in 17 industrialised countries. Using quarterly data from 1970-2006 and a fixed-effects panel VAR, they find that shocks to house prices, credit and money have a significant impact on economic activity and aggregate price inflation. Shocks to GDP, the overall price level (CPI) and interest rates also have an effect on house prices, money and credit. The effects are stronger when prices are booming. Overall, the authors found evidence of a significant multidirectional link between house prices, monetary aggregates and the macro-economy.

The importance of the relationship is further confirmed by Arestis *et al*. (2014), who endogenised the development of bank credit by paying special attention to those variables that are related to the real estate market. In particular, they assume that the main source of demand for credit emerges from households' desire to acquire real assets and they confirm this relationship using cointegration techniques in the cases of nine OECD countries. Building on these results, Anundsen *et al*., (2016) use the link between credit and the housing market to explore how developments in the two affect the likelihood of a financial crisis. Using quarterly panel data for 16 OECD countries that experienced financial crises, they showed that private sector credit and housing market exuberance have a significant positive impact on the probability of a crisis.

The link between credit, the housing market and the macroeconomy has also attracted the interest for country-specific studies. For example, Muellbauer and Murphy (2008) explored the multiple interactions of housing markets with the rest of the economy in the UK. They found that income, housing stock, demography, credit availability, interest rates and lagged appreciation drive house prices. Housing collateral and down payment constraints were also found to have a significant role in house price variations. With regards to the US and the euro area, Musso *et al.* (2010) estimate a structural VAR and find that monetary policy shocks are more significant in the US, whereas credit supply has a substantial impact in the euro area.

In Singapore, Ng and Chow (2004) study the impact of public policy decisions restraining bank credit, using an error correction model (ECM) specification. Their results show that both private and public sector residential properties are affected by the same mortgage rate dynamics. Turk (2015) examined the interactions between house prices and household debt over the long run and the short run, using an ECM specification for Sweden. The results suggest that household borrowing affects house prices in the short run and house prices impact household debt over the long run.

In one of the few studies that have used micro-level data to investigate the housing-credit nexus, Kelly *et al*. (2015) use micro loan-level data on Irish mortgages from 2003 to 2010 to provide estimates of the relationship between credit, house prices and macro-prudential policy. In addition, the authors found that the link between credit availability and house prices allows them to evaluate the impact of the macro-prudential limits in the mortgage market. Their findings show that macro-prudential policies have an important role in preventing a housing market boom. Their results echo those of Favara and Imbs (2015) who present evidence of the impact of credit on house prices using micro-level data.

In line with the literature overviewed, the remainder of this paper presents evidence of the link between house prices, credit and the macro-economy in Cyprus. Whilst the existence of correlation between credit, house prices and the macroeconomy in Cyprus is undeniable, one cannot easily detect the direction of causality, if any. To this end, we begin by providing an overview of data and macroeconomic developments in Cyprus over the past decade.

3. Data

This section provides an overview of the data used in our estimations, based on existing literature and data availability. Specifically, our analysis focuses on the following five variables: loans to households for home purchase, Residential Property Price Index (RPPI), unemployment rates, wages and interest rates charged on loans for home purchase.

Theoretically, house prices and housing loans have a positive economic relationship: an increase in house prices will lead to an increase in loans for home purchase, as the ratio of the loan to the value of collateral will be higher (Adams and Fus, 2010). Similarly, an increased demand for housing loans supported by more favourable loan supply factors (i.e. more favourable credit terms and conditions) should cause an increase in demand for homes/real estate, eventually leading to an increase in house prices (Hempell and Kok, 2010).

Andrews (2010) argues that the relationship between interest rates on housing loans and house prices is negative and depends on the level of competition among banks. When interest rates are rising, the cost of borrowing increases and potential buyers are discouraged. As a result, housing demand and, hence, demand for housing loans decreases. When, on the other hand, interest rates go down, the cost of housing decreases and demand for housing loans increases (Apergis and Rezitis, 2003; Igan *et al*., 2011). According to Frederic (2007), interest rates can affect house prices both in a direct and an indirect way. The direct approach refers to the impact the expectations for future house price movements and credit supply on the cost of capital. The indirect impact is seen through changes in housing wealth and the effect of credit channels on consumption and demand. From a reverse point of view, Jud and Winkler and Painter (2002) and Redfearn (2002) argue that the impact of house prices on the level of interest rates is insignificant, in contrast to Zan and Wang (2012), Goodhart and Hoffman (2008) and other researchers that consider it as one of the most important macroeconomic determinants of housing decision making.

Unemployment is another factor considered important. Smith and Tesarek (1991) found that a decrease in housing activity leads to an increase in unemployment rates. Schnure (2005) estimated that a unit increase in unemployment rates leads to a 1% decrease in house prices.

Finally, another potential candidate for the housing-credit nexus is the level of wages in the economy, which is also expected to have a positive impact on the two variables (Davidoff, 2006). Based on existing literature, wages are used as a proxy for income expectations and are considered to have a significant positive correlation with loans to households for home purchases, as well as house prices (Valverde and Rodriguez Fernandez (2010). This makes intuitive sense, as an increase in affordability driven by an increase in wages could encourage borrowing and, given a temporary fixed supply due to the time it takes to construct new housing units, it would result in an increase in house prices, *ceteris paribus*,. This increase in demand will then feed into higher residential property prices (Goodhart and Hofmann, 2008).

The inclusion of both unemployment rates and wages is made in order to fully capture changes in the macroeconomic environment. As explained below, developments in both variables have been important in Cyprus in the past decade. In essence, the inclusion of these two variables allows us to capture both any changes in the macroeconomic environment through unemployment rates (which are directly linked to output via Okun's law) and any changes in income expectations of those who remained employed[2].

Data on credit, using real housing loans as a proxy, mortgage rates and the residential property price index (RPPI) were obtained from the Central Bank of Cyprus website, while data for real wages and unemployment rates were obtained from the Cyprus Statistical Service. Specifically, the loan variable refers to the notional stock of housing loans, RPPI refers to the Central Bank of Cyprus residential property price index, unemployment rates refer to the Labour Force Survey rates, wages refer to total wages and salaries (employees) and mortgage rates refer to the MFI new business interest rates for house purchase with an initial rate of fixation of up to 1 year[3]. House prices, wages and loans were deflated using the harmonised index of consumer prices, also obtained from the Cyprus Statistical Service. The sample period is limited to 2005Q4 - 2016Q4, as there are no data available concerning loans, mortgage rates and the RPPI prior to this date.

Table 1. Descriptive Statistics: Full Sample 2005Q4-2016Q4

	Loans (%)	Wages (%)	RPPI (%)	Unemployment Rate	Interest Rates
Mean	8.34	1.55	0.57	9.87	4.94
Median	4.42	2.26	-3.66	10.01	5.16
Max	33.20	11.92	25.34	17.61	6.79
Min	-5.19	-13.82	-9.45	3.14	2.98
Std. Dev.	11.92	6.98	10.64	4.87	1.01
Obs	40	40	40	40	40

The results for the full sample (Table 1) appear to mask the business cycle phase in Cyprus. The year-on-year growth of the RPPI remained flat, on average, during each quarter of the sample, as the increase in the RPPI during the boom period was offset by the decrease observed during the downturn[4]. This is supported by the wide standard deviation of the time series. On the contrary, despite decreases in

2. No case of multicollinearity exists as regards wages and the unemployment rate, as the respective correlation coefficient stands at -0.22. This result can be attributed to the relative wage stickiness in Cyprus economy prior to the crisis, an outcome believed to be caused by the fact that wage growth was mainly determined by contractual agreements and cost of living adjustments, which were largely unrelated to changes in the unemployment rate.
3. The notional stock of housing loans is constructed by imposing housing loan net transactions (obtained from the European Central Bank's Statistical Data Warehouse) on the initial stock of housing loans. Notional stocks are adjusted for amounts that do not arise from "actual" transactions. Such amounts are reclassifications/other adjustments, revaluation adjustments and exchange rate adjustments.
4. Additional studies on the behaviour of house prices in Cyprus can be found in Pashardes and Savva (2009), Sitivanides (2015) and Savva and Michail (2017).

lending during the downturn, housing loans were found to have increased by an average of 8% per quarter during this 10-year period with a maximum rate of 33% and a minimum of -5%. Wages also appear to have increased over the period, despite the major ups and downs observed. As for interest rates on new housing loans, it averaged close to 5% throughout the sample period, despite the significant declines observed since 2013. The unemployment rate stood, on average, at 9.7% during the period, albeit again with very wide fluctuations, as reflected in standard deviation values.

As already suggested, the overall descriptive statistics are not very helpful in understanding the changes which took place in the economic environment of the Cypriot economy during the past decade. As such, Tables 2 and 3 present descriptive statistics for two sub-periods, the first one being 2005Q4--2011Q4 (Table 2) and the second 2012Q1- 2016Q4 (Table 3). The results from the sub-periods are indicative of the boom and bust periods in the Cyprus economy.

Table 2. Descriptive Statistics: Sub- Sample 2005Q4-2011Q4

	Loans (%)	Wages (%)	RPPI (%)	Unemployment Rate	Interest Rates
Mean	18.29	6.53	6.39	5.41	5.60
Median	14.78	5.98	-0.77	5.34	5.60
Max	33.20	11.92	25.34	8.94	6.78
Min	4.76	0.60	-6.50	3.14	4.45
Std. Dev.	8.80	3.66	12.40	1.70	0.63
Obs	20	20	20	20	20

Table 3. Descriptive Statistics: Sub- Sample 2012Q1-2016Q4

	Loans (%)	Wages (%)	RPPI (%)	Unemployment Rate	Interest Rates
Mean	-1.61	-3.42	-5.25	14.33	4.28
Median	-1.73	-1.36	-5.34	14.62	4.44
Max	4.09	3.21	-0.94	17.61	5.41
Min	-5.19	-13.82	-9.45	11.09	2.98
Std. Dev.	2.42	5.88	2.72	1.96	0.90
Obs	20	20	20	20	20

As Table 2 indicates, the 2005- 2011 period was marked by rapid credit growth, with a mean increase of 18.3%, along with associated wage and house price increases. As the economy boomed, wages and the RPPI recorded exceptional growth rates, despite higher interest rates. Unemployment was low during this period, standing on average at 5.4%. After the bust, these variables notched negative growth rates, with house prices experiencing the largest average drop, standing at -5.25% (Table 3). During the same period, the average unemployment rate nearly tripled to 14.33%.

In general, the data overview underlines the fact that Cyprus underwent significant changes in its economy during the past decade. The boom in real estate fuelled by strong credit growth, which was, in turn, motivated by an influx of foreign funds, was followed by a path characterised by contracted income, as well as a sharp adjustment in house prices and a subsequent surge in the unemployment rate. The results in the following section aim to shed some light on the structural relationships behind these developments.

4. Methodology

In order to investigate the presence and direction of causality among the economic variables analysed in the previous section, we employ the vector error correction model (VECM) developed by Johansen and Juselius (1990). We define the model as follows:

$$\Delta y_t = C + Ay_{t-1} + \sum_{i=1}^{n} B_i \Delta y_{t-i} + Dx_t + \varepsilon_t$$

where y_t is a vector of (log) endogenous variables, C is the vector of estimated constants, matrix A contains long-run coefficient estimates, matrix B contains short-run coefficient estimates, x_t includes any exogenous variables and D is the matrix of coefficient estimates for said exogenous variables. In the case of Cyprus, y_t contains housing loans (notional stocks), wages, the RPPI Index, the LFS unemployment rate and the mortgage rate. x_t contains a period dummy variable in order to capture the effect of the strong credit expansion in Cyprus, similar to the approach of Goodhart and Hofmann (2008) and Anundsen *et al.* (2016). The period dummy variable is used as an exogenous variable and takes the value of one from 2005Q4 to 2011Q4 and the value of zero, thereafter. All variables employed, excluding the unemployment and mortgage rate, were transformed into natural logarithms.

Table 4. Unit Roots and Stationary Tests

	Phillips-Perron		**Augmented Dickey-Fuller**	
Variables	Levels	Difference	Levels	Difference
Loans	-3.04*	-3.33*	-2.42	-0.63
Wages	-2.27	-3.74*	-2.32	-2.16
RPPI	-1.72	-2.98	-1.37	-3.10*
Unemployment	-0.86	-8.55*	-1.41	-2.19
Mortgage rate	-0.18	-4.82*	-0.56	-4.82*

Note: Phillips-Perron test and Augmented Dickey-Fuller,
* rejects the null hypothesis 5% level, respectively (critical value: -3.02).

Table 5. Cointegration Tests

	Johansen Cointegration Tests		Johansen *et al.* (2000) test
Variables	Trace	Max Eigenvalue	Test statistic
None	0.64* (0.01)	0.64* (0.01)	126.9*
At most 1	0.48* (0.01)	0.48 (0.08)	74.9
At most 2	0.35 (0.07)	0.35 (0.17)	34.5

* rejects the null hypothesis of no cointegration at 5% level.

Critical values for the Johansen *et al.* (2000) test are 113.5, 84.4 and 59.1, respectively.

To examine whether the variables are suitable for inclusion in a VEC model, we first examined them for stationarity. The results, which can be found in Table 4, are supportive of the view that all variables exhibit I (1) behaviour[5]. The Johansen cointegration tests (Table 5) suggest that two cointegrating relationships exist on the basis of the Trace tests and one cointegrating relationship exists on the basis of the Maximum Eigenvalue tests. To settle on the number of cointegrating relationships, we examined them for economic significance and suitability prior to the estimation, as elaborated in Hendry and Juselius (2000, 2001). During this procedure, the second cointegrating relationship was found to be economically insignificant and was, therefore, removed from the calculation. Supporting the choice of one cointegrating relationship, the Johansen *et al.* (2000) cointegration test, which allows for a break in the series (as per the specified dummy), also found in Table 5, showed that only one cointegrating relationship exists. In fact, robustness checks were also conducted, in which two cointegrating equations were included in the VECM estimation, with no qualitative changes recorded. Hence, we proceed to estimate the VECM, using quarterly data from 2005Q4 to 2016Q4 with two lags, based on the information criteria (AIC and SIC) and one cointegrating equation.

5. The unit root process was also tested via the Kwiatkowski *et al.* (1992) and Elliot *et al.* (1996) unit root tests, which reached similar conclusions. Results are available upon request.

Table 6. Cointegrating equation

Variables	Cointegrating Relationship Estimates
Loans	1.00
Wages	-1.95*** (0.14)
House Prices	-0.53*** (0.10)
Unemployment	-0.06*** (0.00)
Mortgage Rate	0.05*** (0.01)
Constant	8.04

*** denotes significance at the 1% level

The order specifies notional stock of loans, first followed by wages, residential property price index, unemployment rate and, finally, the mortgage rate. In essence, the order implies that banks and consumers base their lending decisions on the past performance of the remaining variables, with bank lending affecting all other variables contemporaneously. The cointegrating equation estimates can be found in Table 6 below and show that the relationship between loans and the remaining variables is significant in the long-run. Despite the use of orthogonalised impulse responses, the model is robust to changes in order. This is also confirmed through the use of the Generalised Impulse Response approach of Pesaran and Shin (1998), which, although not reported here, produces almost identical results. These are available on request.

5. Results

5.1. Granger causality

The existence of cointegrating relationships between loans, RPPI and macro conditions suggests that there ought to be Granger causality in at least one direction, on the basis of the Granger representation theorem. Following the estimation of the VECM model, the model's causal relationship was validated using the VECM Granger causality/block exogeneity Wald test. In Table 7, we provide joint $\chi 2$ Wald statistics for the aforementioned variables.

Table 7. Granger Causality Cesults

Dependent Variable:	Loans	Wages	RPPI	Unemployment Rate	Mortgage Rate
Loans	-	6.75 (0.03)	21.00 (0.00)	13.40 (0.00)	0.87 (0.65)
Wages	0.41 (0.82)	-	3.30 (0.19)	0.79 (0.67)	3.08 (0.21)
RPPI	3.09 (0.21)	8.54 (0.01)	-	2.36 (0.31)	0.21 (0.90)
Unemployment Rate	4.53 (0.10)	2.19 (0.33)	2.72 (0.26)	-	3.21 (0.20)
Mortgage Rate	0.02 (0.99)	9.14 (0.01)	8.92 (0.01)	0.45 (0.80)	-
All	21.79 (0.00)	20.83 (0.01)	28.56 (0.00)	16.83 (0.03)	8.48 (0.36)

It is noticeable from the results that a direct Granger causality between the five variables may not always exist. In particular, the mortgage rate is found to be an exception, as no causal relationship can be detected, neither for the interaction of the mortgage rate with the other variables independently, nor for the system as a whole. This could be attributed to the fact that, historically, changes in mortgage rates in Cyprus had mostly been driven by supply rather than demand factors, and perhaps also attributed to persistence due to a long history of a maximum lending ceiling, before the liberalisation of interest rates in 2001.

Nonetheless, a direct causal relationship exists between wages, the RPPI and unemployment rates to loans, as these reject the hypothesis of non-causality. Similarly, wages are found to cause the RPPI, while loans are found to weakly affect unemployment. Overall, long-run causality, evidenced in the last row of Table 7, is present (i.e., statistically significant) in all equations, with the exception of the mortgage rate. Even in cases when a direct causal relationship between variables cannot be verified, a causal relationship may exist for joint interaction of the variables examined, i.e. via indirect effects on the variables at hand. This is further analysed in the following section, using impulse responses.

5.2. Impulse Responses

Figure 1 shows the results from a standard deviation shock to housing loans on the other variables included in the VEC model, using an Impulse Response horizon of 20 periods. As the figure illustrates, the response of wages to the shock is initially large but gradually stabilises in the 20-period horizon.

Similarly, the unemployment rate and the lending interest rate on housing loans do not yield substantial responses to the shock and appear to be insignificant throughout the shock horizon. More specifically, the shock on loans has a long-run impact of approximately -0.2% on the unemployment rate, with the mortgage rate response also maximising at about -0.2% and dying out in the long run. On the other hand, the increase in loans has a positive impact on the RPPI. This impact is persistent, notching a constant increase throughout the horizon, stabilising at about 5% in the long run. This is in line with relevant literature findings, pointing to a link between loans and house prices (e.g. Favara and Imbs, 2015).

Figure 1. Variable Responses to a Standard Deviation Shock to Loans

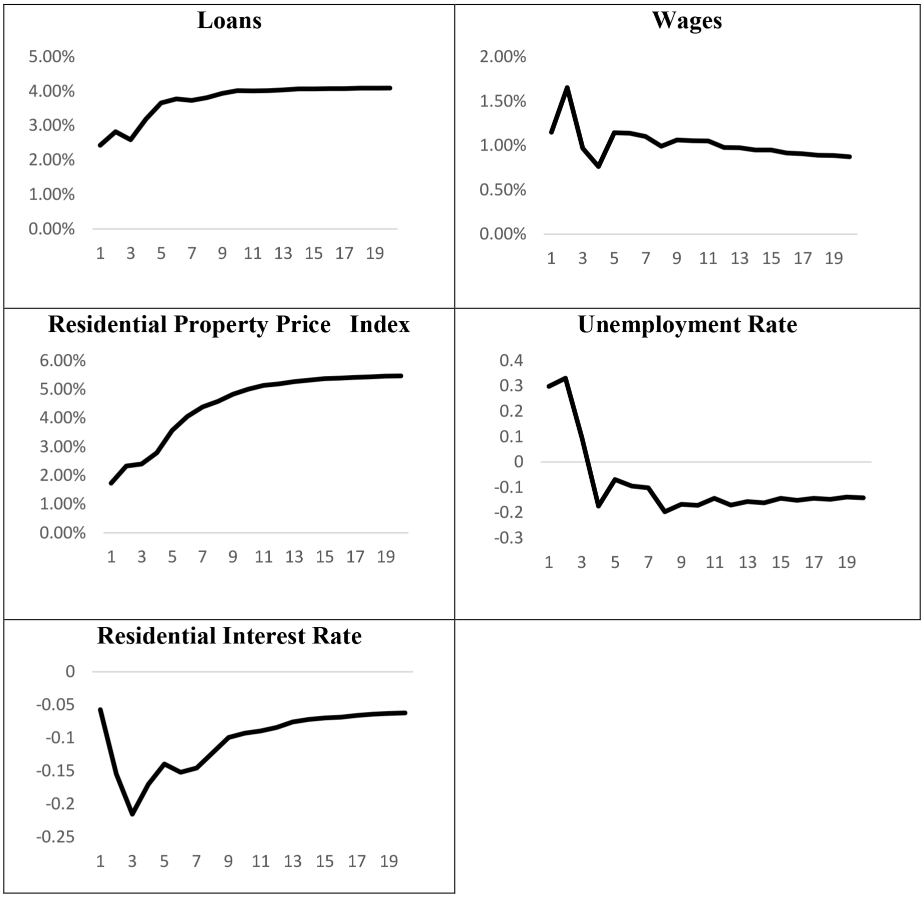
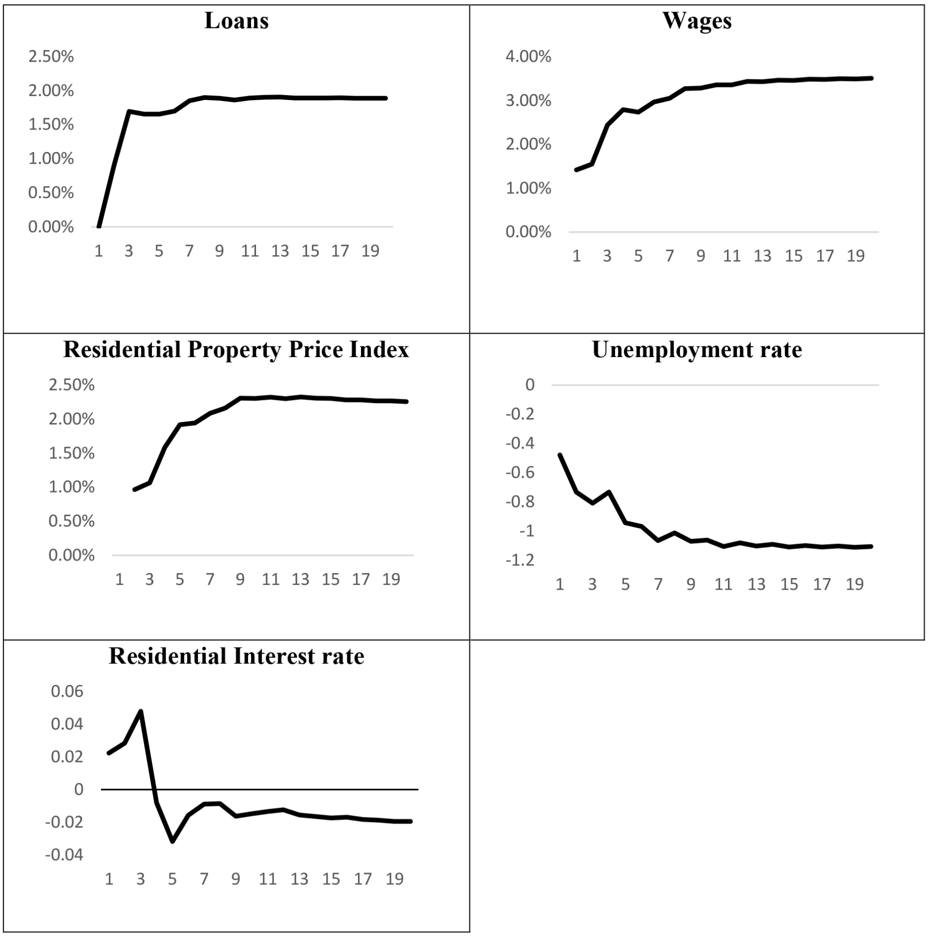
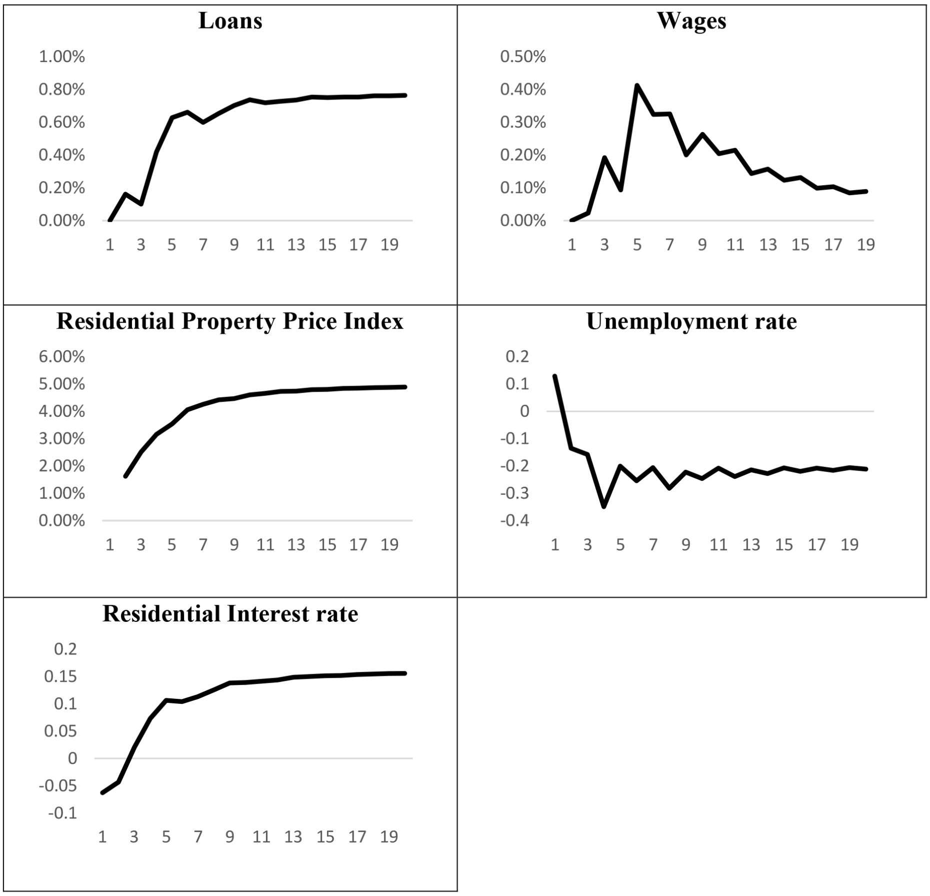


Figure 2. Variable Responses to a Standard Deviation Shock to Wages



The results from a standard deviation shock to wages on the other variables in the model can be found in Figure 2. This shock has a positive impact on loans and the RPPI (after the initial period), as all shocks are persistent in the long run. On the contrary, the response of the mortgage rate is very low, less than 0.05%. An increase in wages yields a persistent decrease in the unemployment rate, reaching approximately 1.1% in the long run.

Figure 3. Variable Responses to a Standard Deviation Shock to the RPPI



In Figure 3 we examine the impact of a standard deviation shock to the RPPI. The increase in residential property prices has a strong positive and permanent impact on housing loans, reaching approximately 0.8% in the long run. The response increases relatively fast, with the slope stabilising after approximately nine periods. At the same time, the shock to property prices leads to an expected increase in residential interest rates, perhaps in an effort by the banks to increase their margins in view of higher demand for loans. A permanent decline in the unemployment rate is also observed, albeit to a smaller extent, reaching its long-run level fast. Wages also increase sharply in the short run, with the effect slowly waning off in the long run.

Figure 4. Variable Responses to a Standard Deviation Shock to Unemployme

Figure 4 shows the responses after a shock to the unemployment rate. Impulse responses suggest that a shock to the unemployment rate has a strong impact on all variables, excluding the mortgage rate, which register a minimal response. Wages record the largest short run response, stabilising at about -3% in the long run. This is in line with the economic theory whereby an increase in unemployment generates downward pressure on wages, given the greater labour market slack. The RPPI reacts much less in the short run, reaching a maximum of -1%, with the effect dying out in the long-run. Loans also register a small, albeit persistent, decrease.

Lastly, the results from a standard deviation shock to the mortgage rate are shown in Figure 5. As expected, loans are strongly affected, registering a reduction close to 1.2% in the long run. Wages react similarly, with a 1.1% decline in the long run, whilst the unemployment rate is also affected by the shock, increasing by approximately 0.6, percentage point by the end of the horizon. A similar impact is registered by the RPPI, which exhibits a sharp decline on account of the assumed proportionately higher contraction in demand for housing relative to supply.

Figure 5. Variable Responses to a Standard Deviation Shock Mortgage Rate

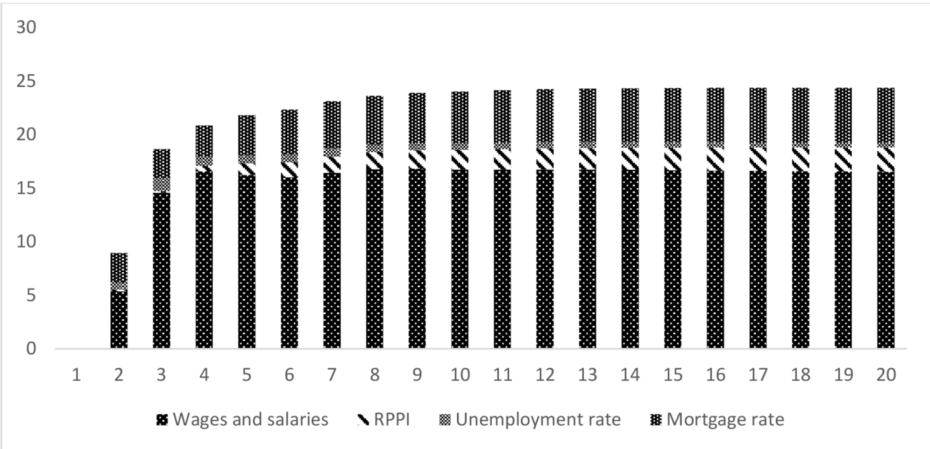
Overall, there is enough evidence to suggest that a bi-directional relationship between loans and RPPI exists in Cyprus. In other words, an increase in the outstanding amount of total domestic loans will bring about an increase in residential property prices and vice versa. At the same time, an increase in residential property prices, driven, for example, by a rise in foreign demand for housing, will eventually lead to an increase in the level of lending to domestic residents, as higher prices necessitate higher lending. A similar conclusion can be reached regarding the relationship between unemployment and residential property prices, with the former declining after an increase in the latter and vice versa. Simply put, higher (lower) unemployment results in lower (higher) residential property prices, given the loss (gain) of aggregate purchasing power.

Another factor that seems to affect both unemployment and residential property prices is the mortgage rate. When the mortgage rate increases, the cost of repayment increases, as does the cost of funding, which leads to lower disposable income. As a consequence, the unemployment rate increases and property prices decrease. There is also evidence of a strong negative impact on wages. Furthermore, a negative relationship is affirmed between wages and the unemployment rate. It is also notable that a standard deviation shock to the unemployment rate has very significant effects on all variables included in the VECM estimated. An overview of the contribution of each variable to the other variables' variance can be found in the following section.

5.3. Variance Decomposition

Figures 6 to 10 present the percentage of variable variances that can be explained by the variance of the remaining variables used in the model. Figure 6 shows that only 25% of the variance of loans in a 20-quarter horizon can be explained by the remaining variables. The lion's share of that 25% is explained by wages, underlying the close relationship between the two. The RRPI and the mortgage rate also explain part of the variance, while, in contrast, unemployment explains the least part.

Figure 6. Variance Decomposition of Loans



In the case of wages, 40-45% of the variance can be explained by other variables (Figure 7). The unemployment rate explains the largest part of the variance in the long-run, even though their effect is minimal upon impact. In contrast, loans have a larger impact in the short run, but the effect wears off in the long run. The Mortgage rate also plays some part in explaining wage variance, while the RPPI has an insignificant role.

Figure 7. Variance Decomposition of Wages

Figure 8. Variance Decomposition of RPPI

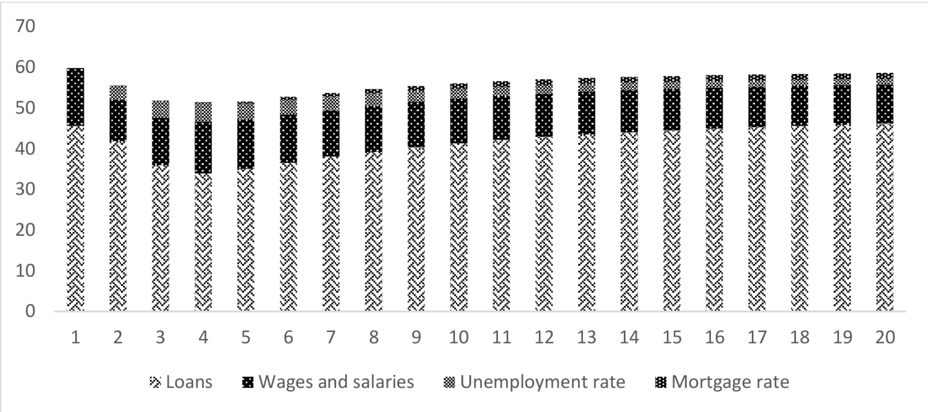
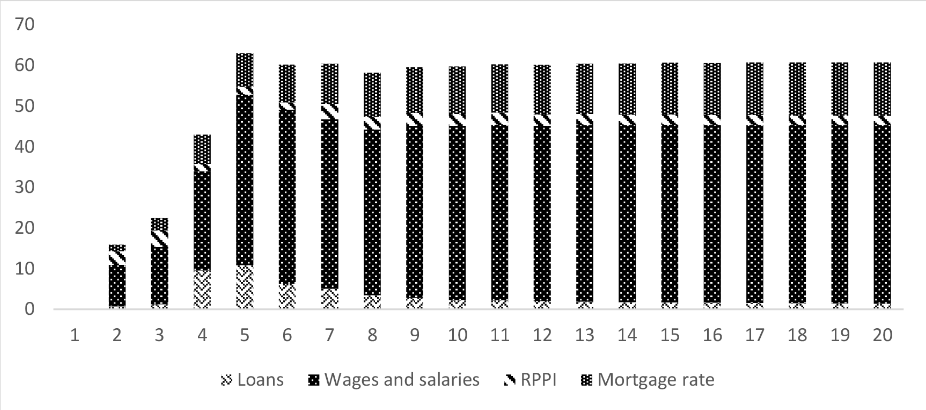


Figure 9. Variance Decomposition of Unemployment Rate



Loans have a strong impact on the variance decomposition of RPPI (Figure 8), maintaining their impact throughout the 20-period horizon. Wages also have a large and persistent effect, while unemployment and the mortgage rate have a minimal impact. Interestingly, more than 60% of the variance of the RPPI can be explained by the VECM variables. Unemployment, as seen in Figure 9, is mostly affected by wages and, to a smaller extent, by the mortgage rate. The RPPI has a minimal effect and loans have some short-run effect, which is eliminated in the long run. Finally, with regards to the mortgage rate (Figure 10) loans explain the largest part of the variance. However, their effect decreases over time. On the contrary, the RPPI impact, albeit small in the short run, increases over time. The unemployment rate has a stable impact throughout the horizon, while wages only have a minimal impact. Furthermore, the mortgage rate is the variable that is least explained by the remaining VECM variables, a result to be expected, given that interest rates usually change due to discretionary policy choices.

Figure 10. Variance Decomposition of Mortgage Rate

Overall, the results of this section are supportive of the findings in section 4.1, as loans and the RPPI have a strong bi-directional relationship. Naturally, the extent of this relationship is not equal, as loans appear to have a more intense effect on the RPPI rather than vice versa. The mortgage rate has a muted response to all variables, except for the unemployment rate and loans. Wages are, as before, heavily affected by the unemployment rate, while a reciprocal relationship is also noted. Interestingly, the RPPI has an effect on the mortgage rate, suggesting that, at times of higher housing demand, it would be possible for banks to raise interest rates in an attempt to secure more profits. In summary, results confirm that the multi-directional link between housing prices, credit and the macro-economy in Cyprus holds.

5.4. Policy Implications

The link between credit, the housing market and macro-economy is one of crucial importance to economic and financial stability. Understanding the link between these variables could also provide a clearer picture of the interrelationship of monetary policy with macroprudential policy. Consequently, these results can be employed as a reference point for bank stress testing and scenario analysis, as they contain very useful information about how the economy reacts to shocks to macro variables, such as the unemployment rate and residential property prices. As suggested by Vazquez *et al.* (2012), this is the first step in creating stress testing scenarios.

Our findings highlight the importance of credit and house price developments for safeguarding financial system stability; they also stress that policymakers should keep a close look at developments in both the real and financial economy when evaluating macroprudential policies. However, this may prove difficult to achieve, especially within a monetary union, given that monetary policy is homogeneous and member-states could exhibit significant regional differences. One such case is the euro area, since the ECB's monetary policy is based on euro area aggregates. If significant regional differences in house prices and credit dynamics exist, the common monetary policy may, at times, prove to be ineffective.

The conclusions reached by this study support the view expressed by Goodhart and Hofmann (2004, 2007, 2008) that one way to overcome the problem of a homogeneous monetary policy in heterogeneous countries within the euro area would be to consider a secondary financial instrument, such as the LTV ratio, at the regional level, to directly address the link between credit and house prices. In addition, we also note that when demand for housing loans increases, banks appear to raise their interest rates in an attempt to secure higher profits, further destabilising the system by increasing the overall debt burden due to higher repayments.

6. Conclusions

The aim of this paper was twofold; firstly, to understand how domestic loans and residential property prices in Cyprus interact with each other, and, secondly, to determine how these variables impact the overall domestic economy. In this respect, a VECM was employed to assess the relationship of loans, residential property prices and domestic macroeconomic conditions. In fact, the outstanding amount of total domestic loans, wages, the unemployment rate, the residential property price index and the mortgage rate were used as dependent variables for the period 2005Q4-2016Q4 and a dummy variable was included as an exogenous variable to account for the period when strong credit expansion was recorded in Cyprus.

The empirical results offer a number of interesting insights. First, a distinct link between credit, housing prices and the macro-economy was found to exist for Cyprus, consistent with other relevant literature. At the same time, a significant multidirectional link between the variables studied also holds in the calculation. Loans appear to influence house prices and vice versa. The results further suggest that shocks to house prices and loans have significant repercussions for economic activity, with the macro-economy also found to have a strong effect on the variables in question. More specifically, a shock to unemployment rates was found to have significant permanent effects on house prices and wages, but less so on loans and mortgage rates.

The variance decomposition of shocks indicates that the response of house prices to a loan shock is of a similar magnitude to that caused by the opposite shock. Unemployment is found to have a similar response, in terms of magnitude, on wages and residential property prices. On the other hand, most of the mortgage rate shock variance is due to its own dynamics. This can be explained by the fact that, prior to 2013, lending rates in Cyprus had been, in their most part, insulated from changes in the macro environment.

All in all, the results of this study confirm the existence of a link between house prices, loans and the macro-economy in Cyprus. However, it should be emphasized that this is by no means definitive. First and foremost, the sample size is limited by data availability, an issue which will hopefully be dealt with in the future. As such, all estimates and subsequent results are, therefore, limited to that extent. Secondly, given that these are currently the only results available for Cyprus, future studies by other researchers will further assist in understanding this link. Finally, a very interesting complication of the results presented relates to the issue of non-performing loans (NPLs) and their impact on domestic economy. Since the issue is beyond the scope of this paper, we leave this very promising field open for future research.

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ASECU South-Eastern Europe Journal of Economics 1 (2019) 57-89 SEEJE

STABILITY OF LABOUR SHARES: EVIDENCE FROM OECD ECONOMIES

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Abstract
In light of ongoing concern about rising inequality in developed economies, this paper revisits the old standing issue of the stability of labour shares. The paper focuses specifically on the empirical aspects of the problem and considers statistical properties of the labour shares in OECD economies in the 1960-2014 period, using a battery of time series models and unit root tests. We account for structural changes in labour shares using Lagrange Multiplier (LM) unit root tests with up to two structural breaks, address the problem of heterogeneous level shifts using LM panel unit root tests, and examine four types of statistical patterns (trend stationarity, mean reversion, random walk with and without drift) using the Augmented Dickey-Fuller (ADF) test. Empirical results indicate diverse patterns in labour share movements, the most preponderant being a downward deterministic trend with break(s). Upward trends are observed in a limited set of economies (Belgium, Luxembourg and the Netherlands). Overall, the stability of the labour share hypothesis appears to find only weak support. Exploratory analysis demonstrates that most of the structural breaks are economically significant and relate to the recent economic and political history of individual economies. The nature of labour share dynamics, as a country-specific and (to a large extent) policy and political phenomenon, is emphasized.

JEL Classification: C22, D33, N10, P17
Keywords: Labour Share, Unit Root, Trend, Factor Distribution of Income

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1. Introduction

The issue of functional distribution of national income was a central thread in discussions of 19th century classical political economy. Whilst frequently superseded in modern economics by other related topics (such as personal income distribution) and, thus, receding into the background, it has recently become topical again, following works by Blaug (1996), Atkinson (2009), Glyn (2009) and Piketty (2014). Hailed as 'the principal problem of political economy' (Atkinson, 1996: 3), functional distribution matters for several reasons.

Firstly, if the theory of the instability of labour shares is validated, the next step would be to establish the factors that caused instability and to consider implications of instability for other models and theories (for example, for production function models, such as the Cobb-Douglas model, that conventionally assumed fixed factor shares). Secondly, an unstable (or more specifically, falling) labour share may affect personal income distribution (Ryan, 1996). Thirdly, if stability of labour share is proven to hold, the re-distribution arguments (such as the demands of unions and workers for higher or 'fairer' labour share and higher wages) would be weakened: any attempts to increase wages and labour shares above 'natural' or equilibrium levels would cause distortions in the labour market, including higher unemployment.

This theoretical debate as to (im-)possibility of stable factor shares and the related empirical analyses (reviewed in detail in the next section) have been vibrant over years, albeit without firm conclusions being reached. In recent years, the focus of the analyses has been shifting to the determinants of and directions in labour share; the statistical analysis of the labour share patterns is likely to be instrumental to this end and can potentially assist the resolution of the "(un-)stable labour share" controversy.

Thus, the purpose of this paper is to provide, using new data made available in recent years, statistical evidence of stability of—or, in contrast, trends in—labour shares and, thereby, complement previous theoretical debate. In addition, without making broader generalisations regarding the determinants and drivers of labour share, the paper interprets in a qualitative manner trends and structural breaks in the labour share series in individual OECD economies in light of their economic history experience and political economic developments during the post-WWII period. The analysis of labour share patterns undertaken in this paper indicates that movements of the labour shares were to a large extend country- and period-specific phenomena, despite similar stages of economic development, growing economic and political integration across OECD economies, globalisation and increasing movement of capital (Arpaia *et al*., 2009). Thereby, consideration of these specific circumstances related to labour share movements would potentially assist the analysis of factor income distribution within OECD and complement conventional cross-sectional analysis that attempts to identify common drivers of labour shares across economies. Recent empirical evidence appears to justify such an approach: as stated by Blanchard (2000), the variation of labour shares by broader economic factors (such as capital-labour substitution or real-wage-productivity divergence) explains no more than 10-40% of the variation, the remainder being likely to be related to country-specific policy and institutional factors. The sample period is limited to the past five decades and, thus, the focus of the paper is confined to medium-term movements in labour shares (i.e. changes over three to four decades, and across business cycles), as well as short-term fluctuations (changes during turning points in the business cycle or during a stage of a business cycle).

The plan of the paper is organised as follows. Section 2 reviews empirical literature pertaining to labour share measurement and labour share determinants. Section 3 discusses methodological issues relating to empirical testing. Section 4 provides empirical results and attempts to establish their statistical and economic significance. Section 5 summarises the paper.

2. Literature Review

In neoclassical economic theory, labour and capital shares are assumed to be stable, and the whole analysis of factor income distribution is confined to and subsumed within the analysis of production functions with constant elasticity of output with respect to labour and labour-augmenting technical progress (Zuleta, 2012). In fact, what was adopted was a conventional assumption of the level of labour share at 2/3 of the GDP. For such assumptions to hold, it is necessary that constant labour share is attributed to constant savings-output ratio, with propensities to save out of wages being offset by propensities to save out of profits (Kaldor, 1956) and that relative price of labour to capital is also stable, based on proportional changes of wage costs in capital-producing and capital-using industries (Lebergott, 1964). As early as the 1950s, it was argued that such a view may be unfounded.

As noted by Solow (1958), the stable labour share in national income or stable ratio of labour to capital income may hold only if movements in the relative prices of labour and capital are exactly offset by counter movements in quantities of factors. In addition, Solow argued that a labour share may be variable due to growth of income of unincorporated enterprises and human capital stock. In a similar vein, Alterman (1964) argued that proportionality of changes in wage costs in capital-producing and capital-using industries can hardly be ensured, as it also requires proportionate changes in the rate of return to capital and capital productivity. Another argument against labour share stability is put forward by Johnson (1954), attributing long-run increases of labour shares to structural changes in the economy (decline of agriculture, where labour share of output is particularly low), the growing prominence of government contribution to GDP, in the form of government employees' compensation, and the decreasing proportion of unincorporated businesses in total labour force. In heterodox economics (in particular, post-Keynesian economics), the view of stable factor shares is likewise disputed: the imperfect competition, as well as varying mark-up power and related varying bargaining power of labour and capital, would cause factor shares to fluctuate (Stockhammer, 2009).

Empirical evidence tends to point towards instability of labour shares. Piketty (2007) argues that the labour share in selected developed economies is stable. However, his empirical results were derived from long samples, covering periods of longer than one hundred years (US, UK and France samples). In the medium-term, the evidence is overwhelmingly in favour of labour share instability, specifically when recent decades are concerned. Krueger (1999) points to the significant variation of labour share in the US over the 1939-1998 period. Rodriguez and Jayadev (2012), using economy-wide and manufacturing sector data, established the decline in labour share at national, regional and global level over the 1950-2005 period, caused by falls in 'intra-sector labour shares as opposed to movements in activity towards sectors with lower labour shares' (p. 1). The evidence of the decline in labour share since the 1980s was provided by: the IMF (2017), which documented the labour share decline of two percentage points on average for a sample of economies between 1991 and 2014; Cho *et al*., (2017), indicating decline in labour shares in OECD economies by an average of one percentage point over the period of 1995-2014; and Karabarbounis and Neiman (2014), who showed that the global corporate labour share declined by 7.8% over the 1975-2012 period. The labour share deterioration tendency in the mid-term is also confirmed by Atkinson (2009), Carter (2007), Bentolila and Saint Paul (2003) and Dunhaupt (2012). In many cases, empirical analyses did not consider factor shares explicitly, but, nonetheless, identified deterioration of the labour share; for instance, divergence of real wage growth from labour productivity growth and, hence, a fall in the labour share (Giammarioli *et al*., 2002).

On the other hand, other historical periods witnessed increases in labour share in many instances: during the industrial revolution in Britain, the rise of factory organisation and the demise of self-employment made the wage share increase (Phelps Brown and Weber, 1953). Similar developments during the 19th century were documented for the US and Germany (Scitovsky, 1964; Jeck, 1968). In other economies, there were also periods when labour share rises were experienced – for instance, in Canada in the 1920s-50s (Goldberg, 1964), Denmark in the early 1960s (Bjerke, 1966), Italy in the 1920-30 decade (Gabutti, 2016), Japan in 1916-25 (Minami, Oro, 1979). In addition, labour share dynamics are not uniform across economies, periods and economic sectors: for instance, according to Giammarioli *et al*., (2002), the decline in labour share was more pronounced in continental Europe, whilst in Anglo-Saxon economies the share remained stable. Harrison (2002) indicated a decreasing labour share trend in developing economies, and an upward trend in developed ones during 1960-1997. Karabarbounis and Neiman (2014) also pointed out that, while labour share decline was experienced in the majority of economies, in 9 out of 59 economies of the sample, the trend was positive over the 1975-2012 period. Period-wise, Rodriguez and Jayadev (2012) showed that across 130 economies, and also within relevant regional groups labour share, decline was mostly a post-1980 phenomenon (with declines becoming even more pronounced in the post-1990 period), while the 1960s and 1970s witnessed a stable or increasing share.

In the short-run, fluctuation in labour share across the business cycle stages was also well documented (Rafallovich *et al.*, 1992; Young, 2004). On a theoretical front, a number of hypotheses were formulated, relating fluctuations of labour share to fluctuation in macroeconomic variables, namely: the 'overhead labour' hypothesis, implying a negative relationship between labour share and capacity utilisation (Bernanke, 2000); the 'labour hoarding' hypothesis, considering procyclical productivity and counter-cyclical wages and labour shares (Caballero, Hammour, 1998); the 'realisation failure' and 'wage-lag' hypotheses, implying a negative relationship between GDP growth and output prices on the one hand, and labour share on the other (Sherman, 1991), and the 'rising strength of labour' hypothesis, assuming a positive relationship between employment and labour share (Boddy and Crotty, 1975). Empirical analysis appeared to confirm these hypotheses: the spikes in labour shares were experienced in the time of major recessions of the mid-1970s and late 2000s (Bruno, Sachs, 1985; Grubb *et al.*, 1982; Chan-Lee and Sutch, 1985; Heap, 1980; Jankowski, 1998, Diwan, 2001). However, similarly to the medium term case, the underlying reasons for labour share spikes appeared to be different. For instance, according to McClam and Andersen (2016: 267), in Austria and Belgium, the spike of labour share was driven by inflation factors, while in Sweden the spikes were due to lagging productivity growth. Likewise, regarding the recent recession of 2008-09, distinct labour share patterns were documented (moderate decline in labour share during the recovery period in many European economies, but drastic decline in the US, Spain and Greece (IMF, 2012).

The absence of national accounts data largely restricted the analysis of factor shares in the earlier periods, in particular, making it impossible to split mixed income into capital and labour incomes and to estimate the income of the self-employed, thereby limiting empirical work to the analysis of wage (rather than labour) share. This problem has been overcome through the construction of the labour share series based on the income-side estimates of GDP. The relevant dataset construction included Rodriguez and Jayadev (2012), Karabarbounis and Neiman (2014), Guerriero (2012) and Penn World Table and Extended Penn World Table projects (Heston *et al.*, 2011; Foley, Marquetti, 2012) for a range of developed and developing economies; Neira Barria (2012) and Tosoni (2014) for Latin American economies;

and Kraemer (2011a) for a set of developed economies. Dataset construction by institutional bodies included the AMECO database by the European Commission and the Structural Analysis Database by the OECD. Most of the datasets spanned several decades, including either the most recent decades or the entire post-WWII period. Historical datasets, from as far back as the 1930s or the 19th century, are provided by Piketty (2014) and Bengtsson and Waldenstrom (2015), mostly for a small set of industrialised nations and some developing economies. In this paper the analysis is conducted based on AMECO database.

The empirical and theoretical literature identified multiple forces determining factor shares in general, and causing decline in labour share, in particular. Given that the analysis of these determinants is not the objective of this paper, we mention them in passing. The forces that potentially caused the decline in labour share (and rise of profit share) in recent decades include: capital accumulation and capital-augmenting technical change (Bentolila, Saint Paul, 2003; Raurich *et al*., 2012); changes in relative prices of investment goods (Blanchard, 1997; Karabarbounis, Neiman, 2014); technological factors associated with the increased use of IT-based capital goods and faster obsolescence of capital goods (Ellis, Smith, 2007); financialisation and the increasing role of financial motives, financial actors and institutions in the operations of the economy (Dunhaupt, 2012); deregulation of labour markets and weakening of labour bargaining power (Blanchard, Giavazzi, 2003; Kristal, 2010), employment rationalisation during cyclical downturns, flexibilisation of labour through contract work and outsourcing, international competitive pressures on wages (Rodrik, 1997), privatisation (Torrini, 2005; Azmat *et al*., 2011); globalisation and greater trade openness (Guscina, 2006; Elsby *et al*., 2013); foreign direct investment and stronger financial capital flows (Furceri, Loungani, 2015), as well as various short-run macroeconomic factors, such as exchange rates and oil price changes (Dombrecht, Moes, 1998). Several factors that could have a positive effect on labour share are mentioned, such as: democratic rule (Rodrik, 1999); offsetting shifts in different industries that keep aggregate labour share stable (Young, 2010), and technological innovation and trade openness (Guerriero, Sen, 2012: 31).

Whilst considerable effort has been made in constructing the labour and capital share series, in identifying relevant driving forces and in analysing trends in factor shares, little or no formal econometric analysis has been conducted to establish stability (or its absence) in factor shares. Although a visual examination of series may suggest that labour share is in decline, a more formal analysis is needed to confirm this hypothesis (in particular, to determine sources of instability, such as presence of deterministic trends or unit root processes with or without breaks).

Formal statistical analysis could help resolve the long-standing controversy as to whether stability of the labour share is an ‘illusion’ or even a ‘mystery’, lacking a theoretical basis (Keynes, 1939: 48; Schumpeter, 1939: 575; Solow, 1958) or, indeed, a ‘stylized fact’ (Kaldor, 1961) or even a law (*Bowley’s Law*, Bowley, 1920).

3. Econometric Methodology

3.1 Data

Labour share data has been obtained from the European Commission AMECO database. The labour share variable (ALDC0 code in AMECO database) was defined as the ratio of compensation of employees for the total economy to the number of employees in all domestic industries, divided by the ratio of GDP at market prices to employment of persons in all domestic industries. The adjusted labour share was, thereby, obtained by imputing the average employees' compensation to the self-employed, based on labour force composition. This way, the adjusted labour share of GDP is calculated (which is greater than wage share) and the systematic downward bias in labour share is eliminated (Gollin, 2002; Ellis, Smith, 2007), as the correct figure, which includes the income of self-employed agents and income of owners of unincorporated businesses, is obtained. Labour share is measured at factor costs, thereby removing the values of depreciation and taxes on production and imports and adding back the values of subsidies. This would give more precise estimates, since these items do not represent returns to production factors (Guerriero, 2012: 6).

This adopted labour share measure is likely to be superior to others used in empirical work. Firstly, it is more robust than the adjusted labour share, calculated by allocating two-thirds of the mixed income from self-employment to labour income (Johnson, 1954), which appears to be an arbitrary procedure that does not account for variation in labour and capital income proportions over time. Secondly, allocation of all mixed income to labour income (Kravis, 1959) overstates labour share, particularly in developed economies, since self-employment generates capital income. Thirdly, the measure adopted does not rely on the assumption of the same labour and capital income proportions for the self-employment sector and unincorporated enterprises as in the rest of the economy and corporate sector (Atkinson, 1983).

The period covered for each economy was set sufficiently long to examine variation in labour share, spanning 1960 to 2014 for all economies in question except Iceland (where the sample included 1970-2014 observations). The paper considers the following developed economies: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Luxembourg, the Netherlands, Norway, Portugal, Spain, Sweden, the UK and the USA.

3.2 Model

As a first step, a log-linear trend model was estimated using the following formula:

$$\ln(LS)_{it} = c_i + \beta_i t + \mu_{it} \text{ ,} \tag{1}$$

where LS denotes labour share for country i in year t , t is the year of observation and μ_{it} is a random disturbance term. The trend value is given by β_i, which represents average annual change in labour share ratio for country i over the period. Specifically, $\beta > 0$ indicates an increase in labour share, whilst $\beta < 0$ reflects its deterioration. The possibility of serial correlation dictates that the model be estimated in AR terms: to this end, the Prais-Winsten procedure is employed. By removing autocorrelation, whilst retaining the first observation, Prais-Winsten transformation improves model efficiency (Doran, 1981; Wang, Jain, 2003: 85).

It is acknowledged (Nelson, 1987) that, if the dependent variable is non-stationary, the OLS estimator may turn out to be inefficient, resulting in spurious trend results (a statistically significant trend when none is, in fact, present). To address this potential problem, we adopt an autoregressive specification of equation (1) that includes trend- and difference-stationarity (Bleaney, Greenaway, 1993; Athukorala, 2000). When re-parametrized in differences and lagged variables, it takes the form of the Augmented Dickey-Fuller (ADF) test regression as follows:

$$\Delta \ln LS_{it} = c + \beta_i t + \sum_{i=1}^{m} E^* \Delta \ln LS_{t-1} + \Phi \ln LS_{t-m} + \mu_t \tag{2}$$

where $\Phi = -\left(I - \sum_{i=1}^{m} E^* \right)$ and long-run trend in labour share is $b = -\beta \Phi^{-1}$. The model incorporates four alternative hypotheses: the presence of deterministic trend ($\beta < 0$, $\Phi < 0$ or $\beta > 0$, $\Phi < 0$), reversion to historical mean ($\beta = 0$, $\Phi < 0$), random walk with drift ($\beta < 0$, $\Phi = 0$ or $\beta > 0$, $\Phi = 0$) and random walk without drift ($\beta = 0$, $\Phi = 0$). Equation (2) is conceptualised as an ideal error-correction model if coefficient Φ (the error-correction term) is significant and belongs to $-1 < \Phi < 0$ (Bleaney, Greenaway, 1993: 351). In this case, change in *LS* is negatively related to its current level, with *LS* being pulled back to deterministic trend or historical mean. In contrast, when $\Phi = 0$, no such reversion occurs and random walk patterns are present.

We also implement more robust unit root tests to confirm the presence (or absence) of trend stationarity, specifically Lee-Strazicich univariate and panel Lagrange Multiplier (LM) unit root tests with up to two structural breaks (Lee, Strazicich, 2003, 2004; Im *et al.*, 2005).

Both univariate and panel versions of the LM test were implemented using Model C, allowing for two shifts in the intercept and trend. Breaks were considered to occur at unknown times and were determined endogenously through a grid search over [0.1T; 0.9T] interval, where T is the number of observations in the sample. The null hypothesis was the presence of unit root with up to two breaks, whilst an alternative hypothesis was trend stationarity with up to two breaks.

The test statistic was estimated using the following equation:

$$\Delta LS_t = d' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum \delta_i \Delta \tilde{S}_{t-i} + \varepsilon_t \tag{3}$$

where $\tilde{S}_t$ is a de-trended series, ε_t is an independently and identically distributed error term, Z_t is a vector of exogenous variables, ϕ is a relevant estimator used in calculating minimum LM statistic. The latter is defined as $LM = \inf_{\lambda} \tau(\lambda)$, where λ is the break location and τ is a ratio of estimator ϕ to its standard deviation. The number of augmenting terms ΔS_t (included to correct for serial correlation) was obtained through a general-to-specific procedure, with the maximum number of augmenting terms k set at $k_{max} = 8$.

The panel LM unit root test is performed to ensure greater robustness of results, given the low power of univariate tests in small samples (Shiller, Perron, 1985). The panel LM test statistic is calculated as an average of univariate LM test statistics for each economy in the panel, as follows:

$$\bar{LM}_{NT} = \frac{1}{N}\sum_{i=1}^{N} LM_i^{\tau} \tag{4}$$

The standardised panel LM test statistics is calculated using expected value and variance of LM_i^{τ}; in effect, $E(L_T)$ and $V(L_T)$ as contained in Im *et al.* (2005). Due to the possibility of heterogeneous autocorrelation errors, these values are selected based on the weighted average of k, determined by the univariate LM test for individual economies.

The standardised panel LM test statistics are, thus, given as:

$$\Gamma_{LM} = \frac{\sqrt{N}[\bar{LM}_{NT} - E(L_T)]}{\sqrt{V(L_T)}} \tag{5}$$

The univariate models (log-linear trend, ADF and LM tests) were implemented sequentially. The trend and ADF models were estimated initially with no structural dummies and, if diagnostic problems appeared (heteroscedasticity, serial correlation and non-normality of residuals), they were re-estimated with dummies and/or additional lag terms. The structural breaks and respective dummies in ADF and trend models were determined through a combination of procedures (residuals from ADF regressions, recursive residuals, N-step forecasts and the Quandt-Andrews test).

The univariate LM test was first implemented with two structural breaks. If only one break was significant (in effect, only one trend dummy variable D_t was significant), the LM test with one break was performed (irrespective of the acceptance or rejection of the null hypothesis). If no breaks were significant, the LM unit root test with no breaks was implemented (Schmidt, Phillips, 1992).

We consider the possibility that three univariate tests (log-linear trend model, ADF and univariate LM tests) and visual inspection may be delivering conflicting results. For this reason, an eclectic procedure is adopted. It is well-known (Kendall,

1953) that *ad hoc* visual inspection without a sensible statistical model is prone to delivering spurious results and patterns; hence, visual inspection is performed, in conjunction with formal tests and based on an analysis of the economic significance of labour share changes. With regard to the log-linear trend model, several authors (Granger, Newbold, 1974; Nelson, 1987) indicate likelihood of spurious trends, whilst others argue that trend models are valid and robust (Canjels, Watson, 1997; Kakwani, 1997), as long as asymptotically valid inference is possible and efficient estimators are available.

The Dickey-Fuller methodology suffers several shortcomings, specifically: the false non-rejection of the null hypothesis of unit root when structural breaks are not considered; low power against an alternative hypothesis of stationarity when large autoregressive root is present, and the tendency to over-reject the null when series contain large negative MA root. At the same time, given that conventional unit root tests results are not definitive sources of information about the series but rather results of an exploratory procedure (Mahadeva, Robinson, 2004: 12) and that adopting a general form of ADF test allows testing multiple hypotheses and detecting a variety of statistical patterns, the results are informative.

Lagrange Multiplier tests are superior to ADF and to standard unit root tests (Perron, 1989), as well as many of the unit root tests with structural breaks. Contrary to Perron tests, LM tests determine the timing of the breaks endogenously. As opposed to Zivot-Andrews and Lumsdaine-Papell tests, LM tests allow for unit root behaviour with breaks under null hypothesis, and, hence, can convincingly accept/reject unit root null (Christiano, 1992; Lee, Strazicich, 2003). Thereby, we consider them the principal analytical instrument with which to make inference.

Six alternative outcomes are possible: (1) if all three tests point to a trend in the series (with or without breaks), it is concluded that labour share is not stable and earlier balanced growth assumptions are less justified; (2) A similar conclusion is reached (albeit in a weaker form), if univariate LM tests suggest trend (with one or two breaks) that one of the other procedures adopted points to the same; (3) If univariate LM tests reject the trend hypothesis and only one of the other tests indicates the trend, whilst another one does not, we conclude that no trend (with or without breaks) was present; (4) Likewise, if univariate LM tests indicate the trend hypothesis, but two other tests do not, the conclusion is that there is no trend (with or without breaks). (5) If all three tests point to non-deterministic behaviour, the trend is not present. (6) If the results of ADF and log-linear trend models override the results of the LM tests (the former suggest trends, whilst the latter do not), no conclusion is reached and further testing is required.

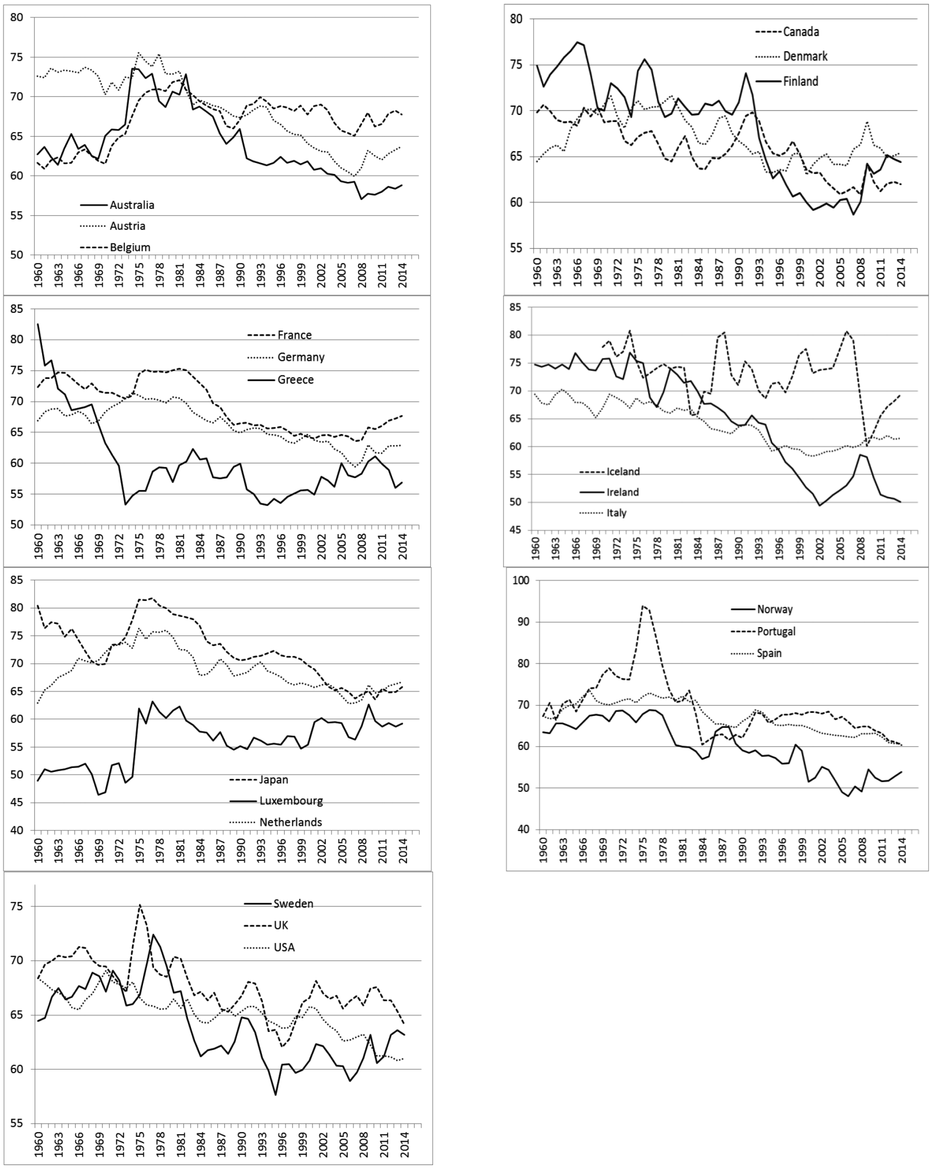
With outcomes (3), (4) and (5), the series tend to revert to the historical mean (particularly when ADF points to mean reversion); hence, labour share is considered stable in line with the predictions by Kaldor (1961) and Bowley (1920). Alternatively, labour share is seen to follow random walk, with or without drift, and no definite conclusions regarding its future direction are possible.

4. Empirical Analysis

4.1 Test results

The visual representation of the labour share series (Figure 1) suggests that, in most economies, the level of labour shares at the end of the sample period was lower than at the beginning of the period. Belgium, Denmark, Luxembourg and the Netherlands stand as exceptions, showing positive changes in labour share.

Figure 1. Labour share (%) in OECD economies, 1960-2014



In other economies, a decline in labour shares was observed, either without major breaks in the series (the case of the USA), or with changes in the intercept of the series (Greece), with temporary increases (Australia, Portugal), or possibly stepwise decline (Norway). In several instances, labour shares appeared to show no distinct patterns or tendencies (Iceland). In terms of magnitude of changes, the largest or most precipitous declines were observed in Greece and Ireland (with labour shares in 2014 standing at 68.8% and 67.0% of their levels in 1960, respectively) and the largest increases in Luxembourg (20.9% increase over the period studied). It is acknowledged that inspection of time plots or estimation of log-linear trend models may lead to spurious trend results and invalid inference, and, therefore, ADF regression is considered.

The results of log-linear trend model with the Prais-Winsten transformation are presented in Table 1.

Table 1. Log-linear trend model estimates

Country	Trend	Cumulative change (%)	p-value	rho	Break	Model
Austria	-0.0032	-17.49	0.00	0.85		TS
Australia	-0.0017	-9.29	0.22	0.91	1974	
Belgium	0.0016	8.85	0.14	0.94		
Canada	-0.0023	-12.34	0.00	0.73		TS
Denmark	-0.0007	-3.62	0.35	0.83		
France	-0.0016	-8.47	0.14	0.96	1975	
Finland	-0.0038	-20.51	0.00	0.84		TS
Germany	-0.0018	-9.69	0.01	0.89		TS
Greece	-0.0058	-31.14	0.01	0.91		TS
Iceland	-0.0020	-8.87	0.12	0.63	2009 or 1983	
Ireland	-0.0081	-43.63	0.00	0.85		TS
Italy	-0.0027	-14.84	0.00	0.86		TS
Japan	-0.0037	-19.72	0.00	0.91		TS
Luxembourg	0.0032	17.49	0.01	0.80	1975	TS
Netherlands	0.0004	2.21	0.78	0.95		
Norway	-0.0051	-27.75	0.00	0.70		TS
Portugal	-0.0027	-14.67	0.17	0.88	1975	
Spain	-0.0024	-12.89	0.00	0.89		TS
Sweden	-0.0016	-8.73	0.07	0.84		TS
UK	-0.0014	-7.60	0.00	0.72	1974-5	TS
USA	-0.0018	-9.91	0.00	0.79		TS

Note: TS represents stationarity around deterministic trend.

Negative trends were present in 18 out of 21 economies, and positive trends were identified in three economies (Belgium, Luxembourg and the Netherlands). Out of 18 negative trends, 12 were found to be statistically significant (at 1% or 5% levels, with the exception of Sweden, where the downward trend was significant at a 10% critical level). Out of the three positive trends, only two were significant (in Belgium and Luxembourg). The largest, statistically significant declines in labour shares occurred in Ireland, Greece, Norway and Finland (-43.63%, -31.14%, -7.75% and -20.51%, respectively) and the smallest in the UK (-7.60%). The larges, statistically significant increase was experienced in Luxembourg (17.49%). These results are supported by a visual inspection of the series.

The ADF model estimates are presented in Table 2.

Table 2. Augmented Dickey-Fuller (ADF) model estimates

Country	δ	p-value	ψ	t-statistics	Break	Trend	R^2	Notes	Model
Austria	-0.0008	0.02	-0.200	-2.60	1975	-0.39	0.21		DT
Australia	-0.0004	0.08	-0.121	-2.27	1974	-0.34	0.35		DT
Belgium	0.0000	0.92	-0.099	-2.11	1974	X	0.17		MR
Canada	-0.0007	0.00	-0.266	-3.29	2009	-0.26	0.26	NW	DT
Denmark	-0.0005	0.01	-0.247	-3.28	2009	-0.22	0.21		DT
France	0.0000	0.91	-0.049	-1.08	1975	X	0.35		RW
Finland	-0.0010	0.03	-0.220	-2.85	1975	-0.43	0.26		DT
Germany	-0.0004	0.10	-0.091	-1.35	2009	-0.39	0.26	$\Delta lnLS_{t-2}$	ST
Greece	-0.0003	0.42	-0.165	-3.05	1973	X	0.30		MR
Iceland	-0.0006	0.33	-0.497	-4.20	2009	X	0.37		MR
Ireland	-0.0020	0.00	-0.219	-3.04		-0.94	0.17		DT
Italy	-0.0006	0.05	-0.196	-2.53	1975	-0.30	0.15		DT
Japan	-0.0002	0.37	-0.062	-1.16	1971	X	0.15		RW
Luxembourg	0.0008	0.03	-0.189	-2.62	1975,1977	0.40	0.64	$\Delta lnLS_{t-2}$	DT
Netherlands	-0.0005	0.01	-0.186	-3.48	1975	-0.27	0.27		DT
Norway	-0.0025	0.00	-0.411	-3.88		-0.62	0.19		DT
Portugal	-0.0010	0.00	-0.282	-5.27	1975, 1984	-0.34	0.52		DT
Spain	-0.0008	0.00	-0.225	-3.64		-0.38	0.25		DT
Sweden	-0.0007	0.02	-0.283	-3.37	1977	-0.25	0.18		DT
UK	-0.0005	0.01	-0.347	-4.39	1974-5	-0.14	0.48		DT
USA	-0.0005	0.01	-0.266	-2.94	1983	-0.18	0.19		DT

Notes: DT, MR, ST and RW represent deterministic trends, reversion to historical mean, stochastic trends and random walk. X indicates that the trend coefficient is not statistically significant. NW indicates Newey-West standard errors due to presence of heteroscedasticity and autocorrelation. $\Delta lnLS_{t-2}$ is an additional lag term to overcome autocorrelation. The t-statistics critical values for a small sample (n = 55) are 2.668, 2.004 and 1.673 at 1%, 5% and 10% significance levels, respectively. ADF test critical values for n = 55 are -4.15, -3.50 and -3.18 at 1%, 5% and 10% significance levels, respectively.

The coefficient of error-correction term (Φ) is negative and, hence, the model is valid. The models passed the usual diagnostic tests (normality, autocorrelation, heteroscedasticity, joint significance of variables). In the case of Canada, the Newey-West estimator was used to overcome autocorrelation and heteroscedasticity and respective standard errors were obtained. In the case of Germany and Luxembourg, the additional lag of the difference variable was introduced to address autocorrelation.

Deterministic trends are likely to be present in 15 economies ($\beta \neq 0, \Phi \neq 0$), if conventional t-statistics critical values (2.004 at 5% significance level for a n = 55 sample) are used to determine the significance of Ψ, or in four economies, if the Dickey-Fuller critical value is used (-3.50 at 5% critical level). In the latter case, the deterministic trend is present in Norway, Portugal, Spain and the UK. With the exception of Luxembourg, all identified trends are negative. The largest decline in labour shares along deterministic trends was experienced in Ireland (-0.94% p.a.) and Norway (-0.62% p.a.). Reversion of the series to historical mean was witnessed in Belgium, Greece and Iceland. Labour share appeared to follow random walk in France and Japan, and a stochastic trend in Germany.

In the majority of cases, the correctly specified model was obtained if dummy variables (of impulse of shift form) representing structural breaks in series were included. The majority of breaks appeared to occur in the mid-1970s (12 breaks), 2009 (four breaks) and early 1980s (two breaks). Importantly, the majority of breaks in the labour share series correspond to rise in the level of the series. Structural breaks in Greece (1973), Italy (1975), Portugal (1984), Spain (1984), Sweden (1977) and USA (1983) stood as exceptions.

The LM unit root tests with breaks demonstrate mixed results (Table 3). For the labour share variable, structural breaks were present in all economies in question (one break in Greece, Iceland, Italy, the Netherlands and Sweden; two breaks in the remainder of the sample) and at least one of the dummy variables representing a change in level or trend was significant at a 5% level. Schmidt-Phillips unit root tests were, therefore, not performed.

Trend stationarity with break(s) was witnessed in all economies except Canada, Germany, Greece, Ireland, Italy and the Netherlands, implying that labour shares were not stable over the study period. The location of the breakpoints was less precise than with the ADF test, with less correspondence to actual economic developments (this, as shown below, being the major shortcoming of the LM test). Nonetheless, out of 37 breakpoints, ten were located in the 1970s, another ten in the early 1980s and one in the late 2000s.

Table 3. Univariate Lagrange Multiplier (LM) test results

Country	LM test (2 breaks)					LM test (1 break)				Model
	Break significance			Break dates		Break significance			Break date	
Austria	-7.621*	[7]	B1D1B2D2	1973	2002					TSB
Australia	-6.038	[8]	B1D1B2D2	1972	1985					TSB
Belgium	-5.299**	[6]	B1D1D2	1972	1984					TSB
Canada	-4.669	[1]	D1D2	1988	2000					URB
Denmark	-5.671	[6]	B1D1D2	1974	2000					TSB
France	-6.290	[7]	B1D1D2	1983	1998					TSB
Finland	-6.634*	[6]	D1D2	1988	1999					TSB
Germany	-5.223	[3]	D1B2D2	1981	2009					URB
Greece						-3.564	[8]	D1	1975	URB
Iceland						-4.708	[3]	D1	2006	TSB
Ireland	-5.175	[8]	B1D1D2	1982	1994					URB
Italy						-3.853	[8]	D1	1992	URB
Japan	-5.474**	[6]	D1D2	1978	1983					TSB
Luxembourg	-7.612*	[1]	B1D1B2D2	1976	1990					TSB
Netherlands						-3.328	[5]	B1D1	1978	URB
Norway	-5.743*	[7]	D1B2D2	1978	1999					TSB
Portugal	-5.672	[1]	D1D2	1977	1991					TSB
Spain	-5.705	[6]	B1D1D2	1982	1995					TSB
Sweden						-4.617	[1]	D1	1981	TSB
UK	-5.920	[1]	D1D2	1980	1998					TSB
USA	-6.560*	[8]	B1D1D2	1981	1998					TSB

Notes: TSB indicates trend stationarity with break(s), URB represents unit root with break(s). B1, D1, B2, D2 indicate significant (at 5% level) intercept and trend dummy variables (for the first and second breakpoints, respectively). Lags selected by general-to-specific procedure are shown in square brackets. In Model C, with one break at 5%, critical values range from -4.45 to -4.51. In Model C (two breaks) critical values are: -6.16 to -6.45 (1% significance level); -5.59 to -5.74 (5% significance level); -5.27 to -5.33 (10% significance level, depending on the location of the breakpoint. In Model C (one break) critical values are: -5.05 to -5.15 (1% significance level); -4.45 to -4.51 (5% significance level); -4.17 to -4.21 (10% significance level, depending on the location of the breakpoint. Series are trend stationary with breaks at 5% significance level unless otherwise indicated; symbol (*) indicates significance at 1% level and symbol (**) significance at 10% level.

The panel LM unit root test (Table 4) was firstly conducted on a full sample of 20 economies, excluding Iceland (for which earlier observations were not available). Secondly, to ensure robustness of results and to account for the possibility of rejection of unit root null, due to only one of the series being stationary (Taylor, Sarno, 1998), the test was implemented on a curtailed basis, consisting only of economies for which the univariate LM test did not reject the unit root null hypothesis (five such economies in the case of the LM test with two breaks and 15 economies in the LM test with single break).

The results of the panel LM unit root test (run on both full and smaller samples) confirm univariate test results. Univariate LM tests pointed to 15 instances where unit root null was rejected. The panel LM unit root test, likewise, indicates very strong rejection of unit root, suggesting, firstly, the higher power of LM tests in a panel framework, and, secondly, the high likelihood of trends in labour shares across economies (as opposed to stability or random walk).

Table 4. Panel Lagrange Multiplier (LM) test results

Panels	LM unit root test with 1 break	LM unit root test with 2 breaks
Panel of 20 economies	-13.804	-27.990
'Unit root' panel	-12.447	-21.323

Notes: The sample for panel LM unit root testing includes all economies, except Iceland. Critical values for the panel LM unit root test at 1%, 5% and 10% significance levels are -2.326, -1.645 and -1.282, respectively.

4.2 Interpretation of results

Systematic analysis of the determinants of labour shares and the identification of common drivers and general regularities require proper econometric analysis in a multivariate setting. However, the focus and the argument of this paper is that most labour share changes in individual economies were country-specific and driven by a set of unique factors. Thereby, even similar economies were frequently exhibiting diverse labour share patterns. We analyse these country-specific developments on a case-by-case basis, looking at the magnitude of changes in labour shares, and by examining the timing of structural breaks and reversals and discontinuities in labour share trends. Given the limited space, we mention only major country-specific factors, while comprehensive case study of labour shares in individual economies can become a fruitful research project on its own.

1). *Australia.* In Australia, a sharp increase in labour share in the mid-1970s was associated with wage-push inflation and 'real wage overhang' (Riach, Richards, 1979; Stegman, 1980: 302). The labour share decline in the 1980s was due to Price and Incomes Accords that brought moderation in wages in exchange for increased benefits outside labour remuneration (Cockerell, Russell, 1995). Further on, the labour share continued to decline steadily throughout the 1990s and 2000s. As put by Australian Bureau of Statistics (2018), the share declined in 11 out of 16 major industries and sectors. Manufacturing experienced an increase in labour share, but the relative contribution of manufacturing to the GDP has declined substantially. In contrast, labour share in finance and insurance and agriculture (which are major sectors in Australian economy) has declined, as these industries were becoming more capital intensive. Expansion of mining (which was increasingly important for

Australian economic development and growth) reduced its labour share by 1.5 percentage points over two decades. Parham (2013) provides similar evidence: labour share fell by over 4 percentage points since 2000. Commodity boom and drastic increase in producer prices (well ahead of the increase in consumer prices) and associated improvement in terms of trade resulted in increase for both labour and capital income. The latter, however, grew faster than the former, thereby leading to a fall in labour share; the real purchasing power of labour did not decline and it offset the fall in labour share. The overall increase in capital stock and capital intensity in the total economy substantially changed economic structure: even with moderation of terms of trade, the reversion of labour share to initial levels will, thereby, be unlikely (indeed the reduction in commodity prices may spur cost cutting by minerals producers and, hence, further fall in the labour share, Stanford, 2017: 6).

2). *Austria*. Possible breaks in labour share were identified for 1973, 1975 and 2002. In line with McClam and Andersen (2016), we argue that the break in 1973 was likely due to the deterioration of Austrian terms of trade ("first oil shock") that heavily affected Austria due to its status of being an oil importer. The labour share declined moderately in the 1980s, stabilized in the early 1990s and fell precipitously in the 2000s. This is likely due to the nature of wage policies in Austria that are strongly geared towards changes in macroeconomic conditions. Real wages are flexible with respect to external and internal macroeconomic shocks and are subservient to a consensual goal of maintaining low unemployment (Hofer *et al*., 2014: 4-5). Until the mid-1990s, the real wage growth was commensurate with productivity growth, resulting in a rather stable labour share. Onwards, the real wages lagged behind productivity, resulting in a declining labour share, while maintaining the competitiveness of the economy and exports. In addition, given the prominent place of Austria in European economic relations with Eastern Europe, trade with this region and outward FDI had strong negative effects on the Austrian employment and labour markets (Onaran, 2008). The effects concerned both high- and low-skill workers and industries, principally originated from the expansion of foreign affiliates of Austrian companies, and, in regard to wage share, were most pronounced in the industry (as opposed to the total economy). In line with this paper's findings, Onaran shows that the expansion of FDI to Eastern Europe resulted in a 25.2% decline in real wages and 18.1% decline in wage share during 1996-2005. Similar negative effects were experienced when final and intermediate imports from Eastern Europe were concerned.

3). *Belgium*. Breaks in the labour share were identified for 1972, 1974 and 1984. The former two roughly correspond to external shocks faced by Belgium in the early 1970s, specifically by the drastic fall in productivity and terms of trade: as an early industrialised country with a large base manufacturing sector (steel, chemicals and machine building), Belgium experienced significant negative effects during the time of oil shocks (Biatour, Kegels, 2015). Labour share dynamics in the 1970s and 1980s constituted an outlier case, compared to other European economies: the labour share has grown substantially since the early 1970s and has never declined to original levels. On the one hand, this was a result of consistently slow productivity growth in manufacturing, as evidenced by a low degree of ICT penetration and lack of innovative leaders in manufacturing (Biatour *et al.*, 2011). On the other hand, this was due to highly regulated labour markets, with a high level of union membership (that did not decline significantly in the past decades) and the coverage of collective bargaining that remained the highest within the OECD (Marx, Van Cant, 2018). The path of labour policy reform has been protracted over the years (Cox, 2007), and this has contributed to a stable and rising labour share (as well as low income inequality and solid income growth of the middle class) at the expense of competitiveness.

4). *Canada.* In Canada, the labour share fluctuations were similar to those of the US, not a surprising fact given the tight integration of these economies (this particularly concerns the stagnation of real wages and the wage-productivity gap, Harrison, 2009). The major difference was a substantial increase in labour share in the mid- to late 1980s. Morel (2006: 4) argues that this was likely a statistical artefact: with self-employment rising from 14% to 17% of total employment, the increase in proprietors' income boosted the labour share. The fall in the early 1990s can be attributed to competitive pressures on the economy and wages due to higher openness and integration with the US under NAFTA (Campbell, 2001). Labour share dynamics in the late 1990s and 2000s mirrors Australia and other resource-exporting economies – expansion of 'low labour share' sectors (mining and oil and gas extraction), resulting in commodity price increases, a wedge between producer and consumer prices and expansion of profits.

5). *Denmark.* In Denmark, the labour share experienced an increase in the 1960s and early 1970s, stability during the 1970s, a decline in the 1980s and 1990s, and stabilization in the 2000s (with respective structural breaks in 1974 and 2000). The labour share increase of the 1960s is attributed to sectoral reallocations: relative decline of agriculture, rise of wages in manufacturing and retail and wholesale trade (Bjerke, 1966). The stable share in the 1970s was a result of pressures for higher wages on the part of organised labour, similar to many other developed economies at that time. A decline in the late 1980s-early 1990s is likely attributed to the reorganisation of a system of collective bargaining: merger and centralisation of employers' organisations (Nieminen, 1997). The reorganisation included introduction of wage regulation mechanisms, limiting the number of times when collective bargaining can take place (Nieminen, 1997), ceilings on pay increases, as well as the move towards more flexible pay systems (minimum-wage agreements, minimum-pay agreements and agreements without minimum rates), where the actual pay is fixed at the firm level (Abildgren, 2008). Arguably, the very moderate decline in labour share over 1960-2014 (and stability in the 2000s) is a result of Denmark making substantial headway in restructuring towards high-value-added sectors or industries, such as transportation services, biotechnology and high-tech agriculture (OECD, 2007).

6). *Finland*. Labour share in Finland experienced a dramatic decline in 1992-94, stabilised to a new normal level and experienced some recovery in the 2000s (Sauramo, 2005). These changes are not solely attributed to the rapid restructuring of the Finnish economy after the breakdown of COMECON and the changed trade arrangements with Eastern Europe (relevant structural breaks were identified for 1975, 1988 and 1999, but not for 1991-92). Indeed, the structural reforms, similar to those conducted in Ireland, have taken place since the mid-1980s (and, hence, the break in 1988) and included de-regulation of the economy, and, in particular, liberalisation of inward FDI regime (Hoj, Wise, 2004; Golub, 2003). While deregulation and greater competition were supposed to reduce profit margins and increase labour share, the effect depended on the structure of wage setting institutions, and the path of creative destruction process (in the latter case, productivity increases due to deregulation could exceed wage growth, and, hence, lower labour share). Deregulation of the economy was assisted by flexible labour market institutions (the absence of firm-level strikes, mild employment protection), while low administrative barriers for business start-ups allowed reallocation of labour to more profitable and productive firms and industries (e.g. information and communication technologies), i.e. assisted creative destruction and new business creation (Kyyra, Maliranta, 2008). The decline in labour share was also driven by cuts in the welfare system, fast recovery in asset values and capital incomes in the post-1991 recession period, and tax system distortions, diverting labour income into capital income. (OECD, 2010: 108).

7). *France*. Labour share in France largely followed the patterns experienced by other European economies, with increases in the mid-1970s (respective structural break in 1975). However, the 1983 break is country-specific: reforms in labour legislation introduced by F. Mitterand's government substantially weakened collective labour, and resulted in greater flexibility of employment and the rise of part-time and contract work arrangements; ironically, such change was introduced and implemented by a socialist government (Sachs, Wyplosz, 1986: 263, 267). In addition, the decline in union membership and animosity between trade unions also played its role (Giammarioli *et al*., 2002: 16-17; Goetschy, Rozenblatt, 1992).

8). *Germany*. In Germany, two major instances of labour share decline are identified. Firstly, the labour share fell in the early 1980s (break in 1981), during the period of reconsideration of Keynesian policies and less reliance of the ruling coalition governments on the support of trade unions (Giammariolli *et al*., 2002: 14-15; Tutan, Campbell, 2005). Secondly, the labour share fell during the early 2000s during the implementation of Hartz reforms, a major change in the labour market policy (Guschanski, Onaran, 2016: 15). It should be noted that German reunification did not appear to have any significant effect on the labour share, i.e. no structural breaks occurred in the early 1990s and no trend reversals took place, in contrast to the drastic decline in union membership following reunification (Ebbbinghaus *et al*., 2000). Indeed, there was some moderation of labour share in the early 1990s, when a sharp rise in wages was experienced (particularly in East Germany, as part of wage equalisation policies, Hoffman, 2000).

9). *Greece*. In Greece, the labour share experienced some increase for a brief period in the mid-1960s, due to the strengthening of trade unions, as one of the manifestations of the democratisation process. However, the share fell drastically in the late 1960s and early 1970s, as a result of anti-labour and pro-business policies of the "regime of the colonels", 1967-74 (Ioakimoglou, Milios, 1993: 96-97). The breaks in 1973 and 1975 represent the end of this downward trend: the share rose moderately in the post-1974 period, thereby supporting the view that democratization has positive effects on wages and the labour share (Rodrik, 1999). The share, however, never fully recovered. The political-economic regime that got entrenched in the 1980s was not conducive to efficiency and economic growth: the slow-growing economy with distortions in product and factor markets and high unemployment rates made it increasingly difficult to redistribute income among the labour force (Alogoskoufis *et al*., 1995).

10). *Iceland*. Of all economies in the sample, Iceland exhibited the most volatile labour share, as well as its most drastic decline during the 2008-09 global financial crisis. This pattern illustrates a set of unique characteristics of the Icelandic economy. Iceland is a small open economy dominated by few resource industries (fishing and aluminium production); as a result, the country depends more heavily than other OECD economies on international commodity prices and is affected by investment cycles in the aluminium industry (Feldbaum-Vidra, 2005). The labour markets in Iceland are flexible: labour supply responds dynamically to the economic cycle and nominal wages are adjusted to cushion external shocks (through reduction of working hours or shifting to part-time work), with low levels of unemployment typically being a favourable result (OECD, 2013: 13; Andersen *et al*., 2011). The debt crisis that took place late in 2008 hit the economy particularly hard and required bigger than usual downward adjustments in wages and the labour share.

11). *Ireland*. In Ireland, the labour share fall proceeded from an originally higher level. The fall was, thus, due to initially low profitability; it also related to distorting tax and accounting practices, such as under-reporting or rental incomes and profits under the older tax regime, and the dramatic increase in profit share following relocation of the headquarters by multinationals, due to the favourable tax regime, and associated transfer pricing practices that started in the early 1990s (Sweeney, 2013: 112, 116; Sharpe, Uguccioni, 2017: 37). The breaks identified for 1994 reflect this development.

12). *Italy*. In Italy, one of the breaks was identified in 1992, representing an accelerated decline in the labour share. This roughly corresponded to the structural reforms of the labour market that took place in 1993 (Pontoriero, 2017). The reform intended to alleviate structural and macroeconomic problems that pervaded the Italian economy in the 1970s, i.e., high inflation, accompanied by devaluation of the Italian Lira, as well as automatic wage indexations in line with inflation, leading to downward pressures on competitiveness. As a result of the reform, the automatic wage indexation was abolished, and tighter connection of wage purchasing power change to productivity change was established, both leading to the decrease in the labour share. It is worth noting that this paper's empirical findings do not identify similarly deep effect by further labour market reforms (1997 Treu law, and 2003 Biagi reform) on the labour share as that caused by the 1993 agreement. In other respects, Italy stands as an exceptional case in that its labour share decline was reversed in the early 2000s and continued rising until present times (Torrini, 2016). Torrini attributes this development to the reduction of mark-ups over marginal costs and the loss of competitiveness of the Italian economy. Furthermore, the rising weight of housing services and the increasing value of imputed rents were also responsible for the labour share fall in the mid-1970s.

13). *Japan*. In Japan, the labour kept falling during the times of rapid capital accumulation and economic growth (the 1960s), starting from a relatively low economic base, accompanied by conservative fiscal and monetary policies within a corporatist and centralised economic management that deliberately weakened labour vis-à-vis capital. The fall of the labour share in Japan that continued until the early 1970s has been well documented by Pempel (1978) citing capital share tripling between 1953 and 1974. Pempel indicates that in the aftermath of the first oil shock, labour unions managed to secure large wage increases (as represented by the structural breaks in 1971 and 1978). Regarding overall decline during the post-war period, Shalev (1990: 71-72) explains sluggish labour in terms of a series of moves by the government to strengthen divisions in the labour movement, creating conditions to nurture loyalty to enterprises, co-optation of trade unions by the firms, deferment of wage increases through seniority-based mechanisms or generous pension packages and the like.

14). *Luxembourg*. The labour share in Luxembourg exhibited similar patterns to Belgium throughout the period (given tight integration of the economies, comprising economic union). The breaks 1975 and 1977 represented the conventional response of the labour to the economic turmoil of the mid-1970s. The further dynamics was different, however: Luxembourg managed to introduce wage moderation policies in the early 1980s (and suspend automatic wage indexations in 1982), thereby restoring competitiveness. It also made substantial progress in economic restructuring towards a service economy dominated by the financial sector (in turn assisted by the rise of Eurodollar markets), and rationalization of the steel production (Zahlen, 2007).
15). *The Netherlands*. One of the structural breaks was identified for 1978. This roughly corresponds to a decline in the labour share due to rapidly increasing unemployment rates in the late 1970s, and the Wassenaar Agreement of 1982 that aimed to reform welfare institutions and to introduce wage restraints in exchange for greater investment, job creation and shorter working hours (Hartog, 1999). The effect of the Wassenaar Agreement was lasting, with wages growing modestly on a par with or below productivity growth rates and the wage share falling dramatically in the 1980s (Salverda, 1998).
16). *Norway*. In Norway, the decline in labour share was attributed to the offshore expansion of the oil industry and related policies to counteract the negative effects of such expansion (OECD, 2012: 148). Indeed if the industry was excluded from calculations, the labour share would turn out to be stable over the study period. As argued by Larsen (2006) and Cappelen *et al*., (2000), Norway managed to avoid the negative effects of the "Dutch disease" and resource curse and to manage well the gains that the development of the oil and gas industry brought. Specifically, the country managed to maintain the competitiveness of the non-resource economy, to diversify exports, and to prevent uncontrolled wage increases (hence, raising the labour share). This was achieved by means of wage controls and income coordination programmes that allowed wages to rise in line with productivity.
17). *Portugal*. In Portugal, the sharp labour share spike of the mid-1970s was driven by a combination of political and external economic factors (Vilares, 1986: 184-185). The spike in 1974-5 can be related to the pro-labour policies of the left-wing government that came to power after the "carnation revolution" of 1974, resulting in sharp increases in real wages (15.8% and 12.6% in 1974 and 1975, respectively) and redistribution of income in favour of labour. Accompanying factors were the drastic decline in GDP, the loss of colonies and disintegration of the colonial empire, massive emigration out of the colonies, as well as the increase in oil prices that affected Portugal, as an open economy, more substantially. Since the mid-1980s the labour share has practically remained

unchanged (with minor upward fluctuations): this reflects the most stringent job protection rules within the OECD, favourable collective bargaining and strike laws, workers' participation and oversight in enterprises, dating back to the 1976 Constitution that aimed to construct a "socialist society" (Bover *et al*., 2000; Cardoso, Branco, 2017: 6).

18). *Spain*. In Spain (Prados de la Encosura, Roses, 2009: 1082; Roman, 2002: 97), the rise in the labour share took place between the early 1960s and the end of Franco's regime in 1975, resulting in the decrease of the capital share of income and profit rate. This is due to real wages growing faster than productivity, starting from a low base (which, in turn, is explained by the deliberate efforts of Spanish economic planners of the 1940-50s to compensate for the lack of foreign investment and the low capital base with profit reinvestment). The moderate wage share (with real wages growing slower than labour productivity) played a role in Spain after 1977, following the demise of Gen. Franco's regime (the demise of the old corporatist wage bargaining system, and the implementation of wage restraints and anti-inflationary policies as part of the Moncloa Pact and the social pacts of 1978-86). These effects are documented by Fina *et al*., (1989: 114-116). In addition, the fall in the labour share (as evidenced by the structural break in 1982) is explained by the adjustments to labour policies that were enacted as part of the accession of Spain to European Community (Giammarioli *et al*., 2002: 18).

19). *Sweden*. In Sweden, the period from 1960 until the late 1970s (the "golden age" of the Swedish welfare state model), witnessed expansion of the welfare state and growing strength of the trade unions, resulting in the labour share growing faster than labour productivity, and, respectively, in labour share increase (Bengtsson, 2014: 298; Bengtsson, 2013). The major labour share spike occurred in 1977-78 (as shown by the structural break identified): this was the result of the failed attempts by the social democratic government to slow down wage growth in exchange for tax reductions, i.e., the so-called 'Haga agreements' (Ahlen, 1989). A series of devaluations implemented in 1976 and 1982 reversed the previous upward trend and resulted in the restoration of the profit share and of the competitiveness of the Swedish corporate sector (this is represented by the 1981 structural break). Real wages continued to fall during 1980-85, accompanied by a decentralisation of the wage-bargaining system. Further reforms implemented in 1996-7 were more conducive to wage moderation and decrease in wage pressures: specifically, the new mechanism of bargaining was based on the export sector setting the norms/limits for wage increases (Bengtsson, 2014: 305).

20). *The UK*. In the UK, the labour share spike in 1972-75 stands out (represented by the structural break in 1974-5). This corresponds to the period of industrial strife and radicalisation of trade unions and labour politics in general (Brown, 2004). The election of E. Heath's conservative government marked an assault on organised labour, as manifested by imprisoning union leaders and passing the Industrial Relations Act (1971) that mandated compulsory ballots before strikes and established the Industrial Relations Court to handle administrative and civil cases against unions. The national coal strike (January 1972) that followed resulted in a complete victory of the unions and in wage increases (in mining by 17-24%). Inflation and nominal wage growth levels peaked in 1975. The labour governments of H. Wilson and J. Callaghan (1974-79) pursued a less confrontational approach: the proposed "social contract" of 1975-77 intended to stymie the wage growth through voluntary wage restrains achieved in consultation with trade unions (Ryan, 1996). Partly successful, the "social contract" policies resulted in stabilisation of real wages for a period of three years; however, major discontent with such policies in times of accelerating inflations led to disintegration of the contract and the general strike of 1978-9 during the so-called "winter of discontent" (Fiorio, 2013: 36). The consecutive conservative governments of Thatcher and J. Major secured the profit share restoration and firm wage restraints during the 1980s and early 1990s through a broad-brush economic deregulation. The partial restoration of the labour share noted in the second half of the 1990s attributed to real wages growing in excess of productivity (Batini *et al*., 2000).

21). *The USA*. In the USA, the labour share exhibited a downward deterministic trend with possible breaks in 1981, 1982 and 1998. The share steadily declining starting from the early 1970s and acceleration in the decline took place in the early 2000s. Elsby *et al*., (2013: 29) attribute the break in the early 1980s to the growth in average labour productivity exceeding hourly compensation growth, while Fleck *et al*., (2011) identify this deviation as early as the 1970s. Importantly, Elsby *et al*., (2013) maintain that the driving forces of the labour share decline were not associated with capital-labour substitution or capital deepening. Following Piketty (2014: 309-10), CEA (2013), Schorr (1991) and Burgmann (2016), the downward trend and, specifically, the structural break in the early 1980s can be attributed to policy and political-economic factors: stagnation of minimum wages during R. Reagan and H. Bush administrations, weakening of organised labour in the 1980s, tax distortions that favoured the corporate sector, financial deregulation (as manifested in the rise of interest share of income in the early 1980s, Dagum, 1988: 215), intensification of work and increase in working hours. With regard to the accelerated decline in the 2000s, the demise of the manufacturing sector is likely to have been a factor: as noted by the Council of Economic Advisers (CEA, 2013), half of the decline in the labour share in the 2000s was due to the decline in manufacturing. In addition, as argued by Parham (2013: 16, 46), during the 2000s, distributional changes were taking place against the backdrop of economic growth slowdown: although both labour and capital income presented deceleration, the former slowed down more than the latter.

5. Conclusion

The principal empirical finding of this paper is that diverse labour share patterns were present and definitive conclusions could be drawn only for a smaller set of economies. Firstly, all three univariate tests suggested the presence of a deterministic trend with two breaks in Austria, Finland, Luxembourg, Norway, Spain, the UK and the USA and a deterministic trend with a single break in Sweden. Secondly, two tests indicated a deterministic trend with two breaks in Australia, Denmark, Japan and Portugal. Thirdly, in the case of Canada, Ireland, Italy and the Netherlands, no definitive conclusions have been possible and additional tests may be needed. Fourthly, in the remainder of the sample, no deterministic trends with breaks were discovered. The ADF indicated possible non-deterministic patterns: mean reversion in Belgium, Greece and Iceland; random walk without drift in France; and a stochastic trend in Germany. Importantly, the labour share direction was not uniform: while in most economies the level of the labour share was lower in 2014 than in 1960, three economies managed to increase their labour shares over the years (Belgium, Luxembourg and the Netherlands).

Overall, rather weak evidence was provided for the stability of the labour share as a 'stylised fact' of economic growth or even as a law of growth. Given the empirical evidence, it appears to be more appropriate to conceptualise stability of the labour share as a working hypothesis with respective implications for production function models, distribution theories and economic policy. Panel LM unit root tests confirmed this finding: seen as a panel, the labour share is more likely to exhibit deterministic trends than to revert to the mean. Regarding the economic significance of labour share patterns, the exploratory analysis of AMECO labour share data and empirical findings suggested that the identified breaks and trends were generally in line with the events and developments in the economic history of industrialised economies between the 1960 and 2010s.

The paper demonstrated that labour share dynamics in short and medium term is likely to be a result of a complex interplay of economic, structural and political forces. Each country likely had its own unique combination of factors that affected its labour share, with possible offsetting or synergistic effects present, making generalisations and identification of a single principal factor behind labour share fluctuations difficult. Certain common tendencies are evident (such as the general decline of labour shares in the majority of economies, and labour share spikes during the mid-1970s). However, the relative strength of the underlying causal factors behind these tendencies (e.g. industry offshoring, privatisation, technical change, trade openness, or decline in the bargaining power of the labour) tended to vary in individual economies. In many cases the shifts in labour shares, the timing of structural breaks and directions of the trends were not in line with these causal factors. Likewise, a similar type of factor could have a differential effect on labour shares. Methodologically, this may suggest that case studies and analyses of country-specific policies and institutional factors may be an appropriate complement to econometric models of labour share determinants.

There are several avenues for future research into labour share stability. Firstly, a more systematic approach may be adopted to distinguish labour and non-labour incomes. Instead of multiplying unadjusted wage share by an adjusting factor (the ratio of the number of persons employed to the number of employees) in an *ad hoc* manner, a preferable approach may be to use national accounts and other supplementary data to separate compensation of employees from gross operating surplus (and other capital and property incomes, such as incomes from homeownership, holding financial assets, capital-funded pensions) for individual economies. To this end, the database of capital shares constructed by Bengtsson and Waldenstrom (2015) for a sample of 19 developed and developing economies may be a useful source. The database, whilst giving robust figures for capital share, covers a limited set of countries, contains time gaps and, more importantly, does not allow comparison of labour and capital shares across economies (one of the reasons is the calculation of labour shares based on either gross or net value added in individual economies). Secondly, this paper attempted only a cursory approach to examining the driving forces of labour shares and identifying breaks. Once established that a labour share was stable or trending, a formal decomposition analysis might be conducted, akin to the one performed by Kraemer (2011b). This is particularly the case for economies subject to multiple political and economic influences in a short period of time (for example, Spain and Portugal in the 1970s and the 1980s). Thirdly, in cases where definitive conclusions were not possible, additional tests could be recommended. Future research might use other conventional unit root tests (Kwiatkowski–Phillips–Schmidt–Shin/KPSS, Phillips-Perron/PP or Elliott-Rothenberg-Stock/ERS), as well as more advanced unit root tests with structural breaks and non-linear unit root tests (such as those developed by Harvey, Mills, 2004; Kapetanios, 2005, in the univariate context; or Westerlund, 2006, for panel data). Finally, future research might concern the effects of labour share changes on other economic variables; for example, investigating the relationship between factor income distribution and rising income inequality or the effect of falling labour shares on aggregate demand or investment activity.

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ASECU South-Eastern Europe Journal of Economics 1 (2019) 91-127 SEEJE

WAGE DYNAMICS AND BULGARIA CO-MOVEMENT AND CAUSALITY

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Abstract
Motivated by recent debates on the possible role of wages as an income policy tool, in this study we examine the dynamic inter-relationship between wages in Bulgaria, mainly in the context of its EU accession. Relative to the WDN studies on the other EU member states, the novelty in this paper is the inclusion of the minimum wage as a possible conditional determinant of the other two wages. We demonstrate that minimum wage increases do not cause changes in average wages in either the government or the private sector. Using variety of econometric tests, we also demonstrate the leadership of private sector wage over public compensation and recommend the implementation of policy measures aimed at labor productivity growth.

JEL Classification: C32, E62, J3, J4
Keywords: Public Sector Wages, Private Sector Wages, Minimum Wages, Causality

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Section 1: Introduction and Motivation

A major concern for economics is to understand how labour markets work because roughly two-thirds of total income is classified as labour income. One such aggregate aspect is the relationship between real public and private sector wages and the causal links running between them. We focus on dynamic inter-relationships, sectoral spill-overs and transmission mechanisms. The issue of wage leadership is further relevant for both policy makers and central banks from analytical and monetary policy perspectives. So far, the literature has mostly neglected the dynamic correlation between the two wages and no studies have included minimum wage together with the analysis of public and private wage dynamics. This is where the contribution of this paper lies.

This paper analyses the co-movement and causality of public and private sector wages as well as the role of the minimum wage in Bulgarian economy for the period of 2000-2016. We chose this period, on the one hand, for the relative stability of the Bulgarian lev and, on the other, because of the 2007 Bulgarian accession to the EU. The first part of the study examines the causal effects between public and private sector wages in the context of existing literature. Most studies of EU countries, such as Lamo *et al.* (2008) point to a strong bi-directional causality and co-integration of both wages with the private sector, established as a leader in wage determination. Our study adds to the existing knowledge by reporting one-directional causal effects running from private over to public sector wages. In line with EU findings, we conclude that the private sector is an established leader in wage determination and that its leadership is stronger than in some EU countries. Despite the presence of a large public sector as a major employer, both in Bulgaria and in the EU, it is the private sector that establishes itself as a leader in wage determination in the case of Bulgaria. We further derive the co-movement over the business cycle and establish that there is a tendency for long-term co-integration. The second part of the study focuses on the role of the minimum wage as a tool for income policy. We find no long-term causal effects of the minimum wage on private and public sector wages. This suggests that minimum wage plays no role in improving welfare, as raising it does not, in turn, raise average public and private sector wages. This finding is in line with the framework in the minimum wage study by Economides and Moutos (2016)[1].

In our empirical analysis we use de-trended quarterly data to obtain a VAR model in a time-series context. Furthermore, we derive all empirical results, in both nominal and real terms, with CPI preferred as a main deflating tool. Results show that real

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1. Economides and Moutos (2016), discussed in Section 2, find that it is impossible for any level of minimum wage to increase incomes of employed workers and, this way, the authors prove the inefficiency of minimum wage as an income policy tool.

public wages react to increases in private sector wages with a lag of half a year (two quarters). In addition, nominal public and private sector wages exhibit a common trend and tendency for long-term co-integration. We use Granger-causality Wald tests to show that the private sector is the leading one in wage determination and that minimum wage plays no role in the latter. Furthermore, applying impulse response functions leads to the same conclusion as do the correlation coefficients between the wages and the Wald tests. The effect of real private wages on public and minimum ones is graphically presented to show that economic adjustments to shocks tend to follow two periods (quarters) of time. Half a year after a wage increase in the private (leading) sector, a public wage increase follows until long-run equilibrium is reached. We further discuss some possible sources of the shocks that establish private wage as the leader. Next, we delve into public and private wage determinants by considering several labour market models, such as perfect competition, monopsony and unions, search and matching frictions, efficiency wages and minimum wages.

The rest of the paper is organised as follows: section 2 reviews the literature, followed by section 3, which presents description of the data, some stylised facts on Bulgarian wages and the process of wage determination. Section 4 explains the methodology employed in analysing the data. We also include some limitations to our study, a policy recommendation in line with our findings, a venue for possible future research and conclusions under section 5.

Section 2: Literature Review

As pointed out in Lamo, *et al.* (2008), the literature on wage dynamics, spill-overs and leadership proves to be quite scarce. Furthermore, its focus falls on the relationship between public and private sector wages with the effect of the minimum wage, if such, studied separately. The few existing models, such as in Demekas' and Kontolemis' one (2000), generally assume a static relationship between public and private wages, where the public influences the private wage through the labour supply principle, i.e. increases in the public wage leave no choice for private businesses but to increase wages in the respective sector, as well.

This direction of causality is reversed in the so-called Scandinavian model, as discussed in Jacobson *et al.* (1994), which finds that the sector more open to international competition is the established wage leader. In the case of Sweden, the leader is the private sector, as it is found to have higher productivity growth. The competitive market theory, according to which wage increases run from the competitive to the protected sector, has been further elaborated by Lindquist *et al.* (2004), who find that the Swedish private sector Granger-causes the public one. Therefore, it can be concluded that, in the case for Sweden, Norway and Finland, it is the market forces that drive wage mechanisms in establishing causal links. However, apart from the two models above, there is no unified theoretical model, since findings tend to be heterogeneous across individual countries.

Not only theoretical models but also empirical results for most countries differ, so we primarily focus on literature concerning the Eastern European region and EU member states. It should be noted that one of the reasons for this difference in wage dynamics may be attributed to the different institutional and wage bargaining processes in place, as well as to the variety of approaches used in analysing the data available. We chose to follow an ECB study of wage interactions over the period 1960-2006 for the Euro area, Euro area countries and a number of other OECD countries (Lamo *et al.*, 2008), as our main references. This ECB study is was undertaken by the Wage Dynamics Network (WDN), consisting of economists from the European Central Bank (ECB) and the National Central Banks (NCBs) of the EU Member States. The study uses a VAR model to find a strong contemporaneous correlation between private and public wages over the business cycle, as well as a tendency towards long-run co-movement. Furthermore, Lamo *et al.* (2008) find that causal links between the two nominal wages suggest that feedback occurs in a direct manner, i.e. through prices. Despite institutional differences across countries, strong correlation and long-term co-integration are reported for the majority of the cases.

In another study undertaken as part of the WDN, Afonso and Gomes (2010) analyse the interactions between public and private sector wages of OECD countries for the period 1973-2000. Their study presents a two-sector system, public and private, where the two wages are estimated by two different wage functions. The study further tests the validity of variables as instruments affecting wage determination. The econometric method used is a three-stage least-squares, an estimation of the two wages and their determinants via a two-equation system. The study reports that public sector wage growth is mainly driven by private sector wage increases and the government fiscal condition. Besides, public wage growth is found to positively affect private wage increases. These results are driven by the validity of instruments used, such as total factor productivity, unemployment and urbanisation rates, growth rates of the average hours worked per employee and fiscal conditions. However, there is no minimum wage present in the model.

Christou (2007) obtains bi-variate VAR estimation on dynamic public and private wage behaviour for the period 1993-2007. The study finds bi-directional causality running between public and private wages. Moreover, the Romanian economy can be regarded as similar to the Bulgarian one, since both countries gained EU accession in 2008 and both share comparable institutional and economic settings. Despite the fact that in Romania government wages are higher, on average, and their sector share is growing faster (in line with most EU countries), private wages are found to equally influence public ones. Christou (2007) ignores minimum wages from the estimated VAR system.

Another important paper in the wage dynamics literature relevant to our study is the study by Demekas and Kontolemis (1999), who find evidence of public wage leadership in VAR analysis for Greece (1971-1993). Demekas and Kontolemis (1999) also use a two-sector theoretical model to find that employment and wage decisions in the public sector are fundamentally different from those in the private one due to the presence of political economy factors (employees are also voters and can be patronised by the government) in the government sector. Moreover, the study reports that increases in government wages lead to both increases in private-sector wages and higher unemployment. Empirical results indicate that the government sector's decisions as an employer are important for understanding aggregate market settings and conclude that this effect of the public sector should not be taken as an absolute fact.

In a very recent study, Vasilev (2015) uses data about Germany for the period 1970–2007 to study the importance of public sector unions within an RBC model, relevant for a number of EU member states. This is relevant to our research, as Vasilev (2015) studies wage dynamics using a micro-founded general equilibrium model. The study also finds that both government wages and public employment share increase at the expense of the private sector. Furthermore, the correlation found between public and private wages in Germany is less than perfect (0.5) but positive, providing some support for the moderate leadership of private sector wage over the public one. In a following paper on German data (1970-2007) in the context of EU-12 countries, Vasilev (2016) models the government sector as unproductive, or wasteful, with public and private wages jointly determined as endogenous variables, once again. Furthermore, public wage determination is only slightly affected by the process of rent-seeking, but still mainly determined by the government's balanced budget and households' supply of labour in the public sector. Overall, Vasilev (2015, 2016) studies public and private wage dynamics in a bi-directional relationship, but does not include minimum wage as a possible income policy tool. This is where we try to contribute with this study.

Another relevant study is that by D'Adamo (2011), who uses a VAR specification to analyze spill-over effects in wage determination for ten Eastern European countries over the last decade. Since results are largely heterogeneous, across different countries, we focus on his VAR models for the two Bulgarian wages. The study adopts the theoretical framework of the Scandinavian model, where the internationally traded sector is the leader for wage determination. D'Adamo (2011) finds that, for Bulgaria, the Industry (Traded) and Services (Non-Traded) sectors are wage leaders and that a weak version of the Scandinavian model applies for the country with the traded (private) sector established as a leader. D'Adamo's (2011) results are also in line with our findings, where private wage exhibits even stronger causal leadership over the public one.

Similar to findings in the literature on minimum wage, there is plenty of discussion on its efficiency as an income policy tool, but there is no systematic approach[2]. The exception is a very recent study by Economides and Moutos (2016), who incorporate minimum wage in a dynamic general model applicable to any country. In their model, workers and capitalists are the two main agents for minimum wage determination. Their study considers the case of perfect competition among firms, while public wages are missing from the model. The government is taken as an agent that imposes the minimum wage in addition to levying taxes. Economides and Moutos (2016) find that it is impossible for any level of the minimum wage to increase incomes of employed workers and that minimum wage is, therefore, inefficient as an income policy tool. The reason behind this finding is the fact that minimum wage introduces inefficiency, since an artificially imposed wage ceiling reduces a firm's profits. The cost of this inefficiency cannot be transferred to anyone else but capitalists and this would result in decreasing returns to scale. Moreover, these analyses are in line with economic theory, which dictates that employers would choose to have fewer workers when a binding minimum wage is imposed on the economy.

Lastly, we explore the literature on Bulgarian wages. There are several surveys conducted as part of the WDN (Wage Dynamics Network), which examine wage rigidity and the main features of the wage-setting process for firms in Bulgaria. Vladova (2012), Lozev *et al.* (2011), Loukanova (2011) and Paskaleva (2016) report a relatively weak wage-price link in the case of Bulgarian wages, suggesting that labour cost growth is not fully in line with productivity growth. Besides, Loukanova (2011) finds that the minimum wage is not a push-up for the average one and that, in fact, it affects only wage values close to it. Lozev *et al.* (2011) and Paskaleva (2016) report that wage changes occur only once a year as compared to the eight-month price duration. The latest survey of Bulgarian wages, discussed in detail in the next section under stylised facts, finds that firms with minimum wage prevalence claim that economic uncertainty, high payroll taxes and changes in labour laws lead to lower employment of workers.

Section 3: Data description and Stylised Facts

3.1 Data Description

We use data from the NSI (October 2016) database on Bulgarian CPI and wages. As data on nominal wages are reported on a monthly basis, we converted them into quarterly ones for easier modelling. We used NSI data on average earnings

2. For example, see: Burkhauser R. and J. Sabia (2007), Economides G. and T. Moutos (2014, Neumark D. and W. L. Wascher (2008).

for government, private and minimum wages approximated as compensation per employee. Compensation for all private sector employees is defined as compensation of all employees minus compensation of government employees in the Bulgarian economy. Compensation per private sector employee is then computed by taking the private compensation of employees, divided by private sector employees minus government employment minus self-employment, as in Lamo *et al.* (2008). The rest of this section provides a detailed explanation on how wage determination mechanisms work in Bulgaria.

Regarding wage measurement, we have taken compensation per employee both in real and nominal terms for the period 2000-2016. We use CPI as a main price deflator to obtain real wages, with all specifications conducted in a time-series context. The driving force behind deflating wages is to exclude possible shocks and minimise the possibility for spurious outcomes when modelling wage relations. Figure 1, below, shows that wages seem to share a linear trend, which may also indicate the presence of non-stationarity, co-movement and possible long-term co-integration.

As evidenced from Figure 1 on the next page, the 2008 economic crisis led to private wage decreases and higher unemployment rates. During this period the gap between private and public wages was the highest and lasted until the end of 2010, when the economy experienced some positive growth. At the beginning of 2011, the private wage was again closely co-moving with the public one. After a few years of adjustment, the two wages have nearly converged (2016). To show that our results are not an artefact of inflation, real wages in Figure 2 below are shown to have similar co-movement and trend over the period studied.

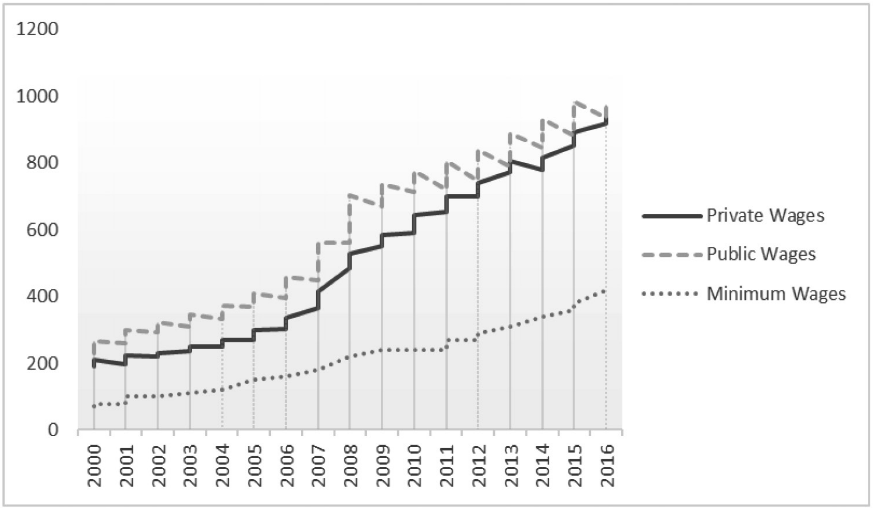

Figure 1. Nominal Wages Movement

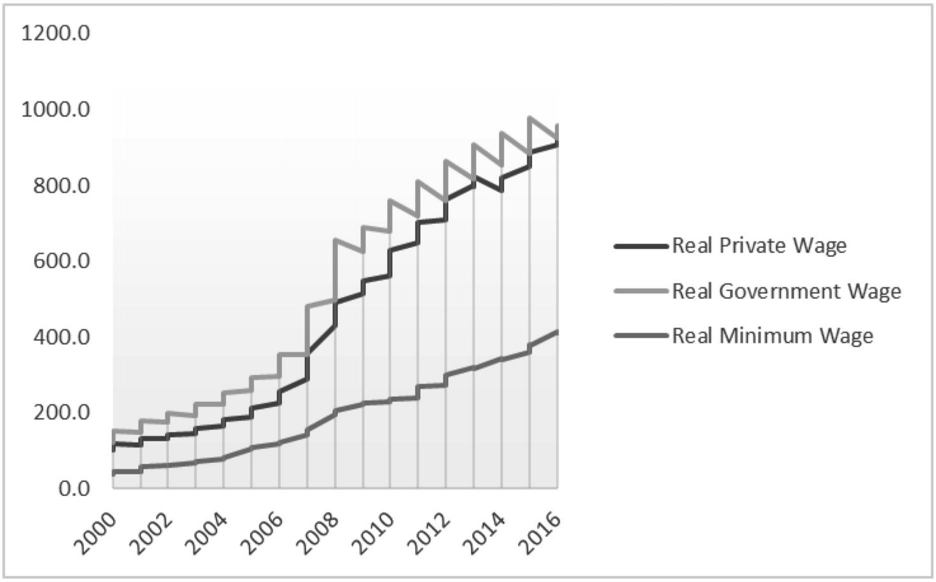

Figure 2. Real Wages Movement

The increase in government sector wages during the crisis suggests poor adaptability of wages to economic performance, as also pointed out by Loukanova (2011). Moreover, as employment decreased, many low-skilled workers were laid off (approximately 20% according to Loukanova (2011)) which lead to increases in average wage. The latter hints that the government budget deficit was not handled very well during that period, as it indicates lower economic efficiency and productivity. Furthermore, productivity increased in the period after the crisis (2011), but wage growth was slower and real wages actually decreased.

3.2 Public and Private Wage Determination

First, as shown in Figure 1 above, public sector wages are consistently higher than private ones for the period studied (2000-2016). This is particularly important in the context of EU government spending on labour, since the latter has been increasing in many European countries over the past years, as pointed out by Afonso (2008). Moreover, as documented by Vasilev (2015), the increase in the public wage bill, as evidenced in the post-World War II period together with the increase in public sector employment have led to a higher share captured by the public output for selected EU member states.

As for Bulgaria, there are several forces behind wage determination. The main factors for public and private wage formation are the legislation system and collective bargaining forces acting in accordance with budget constraints. Wage setting is determined through the so-called Tripartite Cooperation bargaining, comprising government representatives, business employers and unions, as documented by Nenovsky and Koleva (2001). Further, collective wage bargaining can be differentiated at two levels: sector and company. Companies set their wages according to profit constraints, which, in turn, limit the rigidity of labour unions in the private sector. Since firms are profit-oriented, unions cannot ask for excessively high wages, or else the firm would not be able to afford such costs and would eventually close down. The other possible reason for lower private sector wages may lie in monopsonies existing in certain sectors. Many firms in Bulgaria were privatised and, as a result, they can now be viewed as a single regional employer, which would allow for setting lower wages and creating an envelope wage practice.

Moreover, collective bargaining in Bulgaria is mostly characteristic at the firm rather than the industry level, with most companies following their own price setting. This makes the exact determination of private wages more difficult, contributing to the weak price-wage link. Firms that faced worsened economic conditions implemented reduced flexible wage components and wage cuts in the private sector. The 2009-2013 period was marked by deterioration in demand and worsened customer ability to pay (Lozev *et al.*, 2011), which are factors connected to wage rigidity. Survey results show that wage changes occur in approximately 27 months for the above period, a figure relatively low for the EU (Paskaleva, 2016). The main reason behind it is that firms reported a preference for wage freezes rather than wage changes, which can be taken as yet another explanation for the weak wage-price link in Bulgaria. All of these reasons can be viewed as influences on real wages, as well as on their slower growth and change frequency, for the last couple of years (Paskaleva, 2016).

Rose (1985) notes that public employees depend on private sector employees, since taxes generated in private sector activities account for the revenues paying the salaries of most public employees. Direct comparison between the two sectors is not possible, for various reasons. First, the output of the two is not comparable, since it cannot be measured in monetary terms. Next, productivity cannot be fully captured, as some public goods are more labour-intensive than others and the working time of employees in the public sector differs from that in the private sector. Furthermore, according to Rose (1985), having higher public wages can be viewed as somewhat counterintuitive because there should be a trade-off between higher level of job security, the benefits provided by the government and the level of wages. However, this is not the case, since public employees receive better compensation than private ones, despite the fact that private wage earners face higher taxes and lower job security. Having higher public wages than economic principles dictate can be an indicator of the presence of political economy. It should be noted that government employees are also voters and enjoy higher labour union protection than their private sector counterparts.

One way to look at wage settings, as pointed by Vasilev (2015), is that private sector wage is determined within a competitive market framework, while public sector wages could be viewed as a solution to a bargaining process between unions and the government. This means that, when there are government funds available, labour unions in the public sector would ask for higher wages, as they have greater bargaining power than their private sector counterparts. This is due to the lack of profit motive in government administration, together with the fact that the government may start running budget deficits, as opposed to profit-maximising firms in the private sector. As a result, the government wage tends to be higher than the public sector wage.

3.3 Minimum Wage Determination

There has been a mandatory minimum wage in Bulgaria since 1990. Its scale is determined as a nominal value per month and hour and has been further calibrated according to the poverty line level introduced in 2007. As documented by Loukanova (2011), the minimum wage was introduced due to the strong bargaining power of trade unions, which argued in favour of protection of low-income workers and of establishing a basic standard of living.

In practice, however, there are no long-term effects of minimum wage as its increases are not in line with the slow rise observed in labour productivity. Every subsequent minimum wage increase can be shown to raise the bar for marginal worker productivity, which, in turn, forces firms to seek more productive workers, as pointed out by Loukanova (2011). Therefore, it can be argued that the minimum wage practice may actually hurt low-skilled workers despite its initial purpose to provide protection for them. Currently, the minimum wage is 420 Bulgarian leva (EUR 214.74) and is expected to reach approximately EUR 230 in 2017. The government has been steadily increasing the minimum wage over the past few years (see Figure 1 above) which, however, has not been effective in narrowing the gap between minimum and average wage. Therefore, if wages are driven by labour productivity in a competitive setting, minimum wage increases would have no effect in raising the overall standard of living.

Section 4: Methodology

4.1 Unit Root Tests

In this section we focus on the co-movement and causality effects of the public, private and minimum nominal and real wages. We use a vector autoregressive (VAR) model of the three wages, suited for analysis of short-, medium- and long-run correlations at different forecast horizons. The VAR model is a tool to study dynamic inter-relationships between variables. As in Lamo, *et al.* (2008), we use den Haan's methodology (2000), which can be applied to both stationary and non-stationary variables. We use the non-stationary method in obtaining long-term co-movement specifications and the stationary (de-trended) series for the Granger causality tests and impulse-response functions. Moreover, Dickey-Fuller (1979) and Phillips-Perron (1988) unit root tests confirm the need for de-trending and the existence of a single unit root. We, also, further use Breusch-Godfrey (1978) and Durbin-Watson (1951) tests for serial correlation, as discussed in section 4.3.

In our case, the de-trending process applied takes out the deterministic trend component and filters data for seasonal disturbances and cyclical adjustments. The empirical literature on the issue, such as Lamo *et al.*, (2008) applies different types of de-trending methods and filters for the sake of obtaining a non-spurious econometric specification. We, however, focus primarily on using the Hodrick-Prescott (1997) filter, complemented by removal of the seasonal component to de-trend variables. The underlying assumption is that data are integrated at order one (I_1), i.e. the variables contain a unit root as well as a seasonal component which fluctuates around a deterministic trend, both of which become inert when the series are de-trended. As for the forecast horizon, we use a period of ten quarters to obtain impulse-response functions in the short-run.

We test for unit root at 5% significance using the Augmented Dickey-Fuller test with four lags. As evidenced from Table 1 below, nominal private wages display a unit root – the high Mackinnon value of 0.4423 indicates the series are not stationary. There are also a trend and a drift present. Therefore, as Figure 1 in the previous section suggests, we need to seasonally adjust the series in order to account for the difference in private sector salaries during different seasons. To smooth the series, we seasonally adjust them and apply a Hodrick-Prescott filter for the cyclical component, resulting in stationary variables. This procedure is equivalent to first-differencing, which we apply as an alternative method to account for non-stationarity. The results show that the series are first-difference stationary, or I_1.

After removing the seasonal component from the nominal public wage, the results point at the presence of a unit root and a drift (initially, as shown in Table 1, there is no unit root prior to seasonal adjustment of the data); therefore, we also apply the Hodrick-Prescott filter to smoothen data. Regarding the minimum wage, its significant unit root is accounted for in the same manner of de-trending and, as expected, there is no trend present. The reason for this is that changes in the minimum wage occur more rarely and thus are not subject to seasonal disturbances as often as the private and public sector wages are. In order to account for possible spurious results, we also apply the Phillips-Perron unit root test and obtain the same or similar results. After removing the unit root component, the variables display a Mackinnon p-value close to zero, no trend and a constant factor. We also apply first-differencing and obtain a Mackinnon value of 0.000 and, therefore, account for the presence of the single unit root in each variable. The second part of the table displays wages in real terms, which bring about similar outcomes[3]. The same procedure of accounting for the unit root is implemented. The difference here is that the minimum wage displays a significant trend, which is smoothened by the Hodrick-Prescott filter, as standard procedure suggests.

Table 1. Unit Root Tests in Levels

Variable	Mackinnon p-value	Trend p-value	Const. p-value	Order of Integration (I)
Nom. Private Wage	0.4423	0.017	0.041	1
Nom. Public Wage	0.0297	0.001	0.001	1
Nom. Min. Wage	0.9776	0.389	0.382	1
Real Private Wage	0.3903	0.017	0.856	1
Real Public Wage	0.1377	0.006	0.409	1
Real Min. Wage	0.5947	0.029	0.789	1

4.2 Co-integration

Following standard practice, as in Lamo *et al.* (2008), we measure long-term co-movement using the cross-correlation functions for the three non-stationary wages. We use de-trended (stationary) series for obtaining the correlation coefficients and the short-run co-movement and causality results. In short, the only tests for which we do not use de-trended variables are the co-movement and co-integration tests, since establishing long-run links requires the variables to be non-stationary, as documented in den Haan (2000). The reverse methodology is

3. For unit root tests in logs, refer to Appendix A.

applied in the Granger-causality series of Wald tests, correlation relationships and impulse-response functions, where data are required to be stationary and, therefore, de-trended. Our findings in Table 1 on the previous page and Table 2 below are in line with stationarity tests performed by D'Adamo (2011) for Bulgaria and several East European countries (particularly Romania), which also confirm the presence of a unit root and exactly one long-term co-integrating equation for public and private wages. Test estimations show that the number of optimal lags to be used is four, as expected by the quarterly nature of the data. We used Akaike, Schwarz and Hannan-Quinn (1979) information criteria in determining the number of optimal lags.

Table 2. Selection Criteria for Optimal Lag Number

Selection-order criteria			
Lag	**AIQ**	**HQIC**	**SBIC**
0	31.9995	32.04	32.1025
1	24.6771	24.8388	25.0888
2	24.3033	24.5862	25.0238
3	23.7173	24.1214	24.7466
4	**23.21***	**23.7353***	**24.548***

Table 3 below summarises the results obtained from a Johansen co-integration test. There is a single co-integration equation between public and private sector wages, as confirmed by D'Adamo (2011). We also run co-integration tests on minimum wages with the other two and obtain no co-integrating relationship in either case.

Table 3. Johansen test for Co-integration

Johansen tests for cointegration		
Maximum Rank	**Trace Statistic**	**5% Critical Value**
0	17.8539	15.41
1	**1.0984***	3.76
2	-	-

4.3 Correlation

In this subsection, we focus on the correlation of the three wages. Table 4 summarises the results at both level and log forms. Each row displays the correlation coefficients between the three variables, obtained by using a model of differenced and seasonally adjusted wages at time t and t-k (k stands for the number of lags). Since our model is derived by using quarterly data and because rank tests show that the number of optimal lags is four, we trace the correlation relations for four periods. The first output table reports nominal compensation per employee and the second displays real compensation as deflated by the CPI.

Table 4. Nominal Compensation Correlation

Correlations of nominal detrended public and private wages per employ ee in levels and logs									
k(lags)	-4	-3	-2	-1	0	1	2	3	4
Level Forms	-0.14	0.02	0.05	-0.02	0.12	-0.03	**0.54***	0.26	0.06
Log Forms	-0.26	0.06	0.16	0.08	0.24	-0.06	**0.45***	0.29	-0.02

Following common practice on wage dynamics, as in Lamo *et al.* (2008), we base our analysis on the evidence that two variables are said to co-move in the same direction if the absolute maximum value of the coefficient estimated is positive. Furthermore, the variables move in opposite directions if the coefficient of the same de-trended series is negative and they do not co-move if the coefficient is close to zero. Again in line with Lamo *et al.* (2008), we take values between 0.30-0.39 as an evidence of weak to moderate correlation and values above 0.40 as evidence of strong correlation in absolute terms.

Table 5. Real Compensation Correlation

Correlations of real de-trended public and private wages per employee in levels and logs									
k(lags)	-4	-3	-2	-1	0	1	2	3	4
Level Forms	-0.31	-0.04	0.22	0.04	**0.43***	0.10	**0.33***	0.20	-0.14
Log Forms	**-0.33***	0.04	0.18	0.14	**0.54***	0.17	0.41	0.23	-0.15

In both cases, the variables are filtered using Hodrick-Prescott filter (1997) to account for seasonal and cyclical components. Table 4 presents the coefficients in nominal terms; an asterisk (*) marks the highest correlation coefficient, which is observed to be at a lag of 2 periods both in level (0.5445) and in log (0.4492) terms. We regard these values as an indicator of a strong correlation between public and private wages in nominal terms.

Table 5 presents data for nominal wages deflated by CPI, where we derive similar outcomes. The strongest correlation occurs at zero lag, whereas a moderate one can be observed at lag 2 for level forms and at lag -4 for logs. The negative correlation implied by the latter is, however, insufficient evidence against the general positive co-movement of wages, since its value is inclined more towards the weak toward a moderate wage relation and, furthermore, it occurs at only one value. Therefore, we regard the stronger and more persistent positive correlation as a sign of co-movement and causal relationship between wages. We explain these causality links in the following sub-section.

4.4 Empirical Results: VAR Model. Causality

One of the most preferred methods for establishing causal relationship in empirical analysis and literature is that proposed by Granger (1969), which states that a variable X is said to Granger-cause another variable (Y) if it provides statistically significant information about Y. We consider a result to be statistically significant at the 5% level of significance for all tests. In this section we use Granger's (1969) definition of causality for establishing causal links between public, private and minimum wages. Next, we use impulse-response functions to compare results and further evaluate causality links. Following the ECB study on wage dynamics by Lamo *et al.* (2008), we use VAR or vector autoregressive systems and Wald tests for public and private wage causality and extend the study to additionally incorporate minimum wages. The wage variables included in the VAR model are de-trended and filtered for cyclical and seasonal components to account for the existing trend and for possible spurious results.

The following equation captures public, private and minimum wages of the VAR model. C is a vector of constant factors and A is a 3x3 matrix which contains all VAR coefficients of variables of lag from 1 to p. W^p, W^g, W^{min} denote nominal private, government and minimum wages and ε_t are all possible influences outside the model.

Table 6 on the next page displays statistical outcomes of simple probability tests on nominal wages[4]. The table should be read as follows: the excluded variable is the estimator which causes the equation variable at 5% level of significance. Only the real private wage causing the public one has a significant p-value coefficient; all other Wald test probabilities are found to be insignificant. Additionally, the real minimum wage is shown to have no effect on either real public or private wages. Its highly insignificant coefficient values raise the issue of possible policy recommendations discussed in the next section.

As evidenced from the Wald tests, the dominant pattern for all possible testing adjustments is for private wages to lead public sector developments over the business cycle. Nominal wages show similar results to those displayed for both log and level forms. When prices are accounted for, there is a 0.07 probability for the level

4. Correlation data and coefficients on real wages can be found in Appendix B.

forms of real private wages causing public ones. Given that we take a 5% level of significance, the case of real wages in levels is the only one where we can reject the hypothesis that private wages cause public ones. As for the tests at logs, the causality running from private to public compensation is, again, highly significant (0.002), as shown in Table 6 below.

4.5 Impulse-Response Functions

In this section we analyse the Impulse-Response Functions obtained after running a VAR model. An IRF indicates the impact of an unanticipated one-unit change in the 'impulse' variable and the effect it has (if any) on the 'response' variable. In general, IRF functions are used for determining whether one variable is capable of forecasting another over a specified time horizon. Furthermore, IRFs capture the reaction of a dynamic system in response to some external change; in our context they capture how adjustments in one wage affect the other wages. In line with the ECB study on wage dynamics by Lamo *et al.* (2008), we took the results of the differenced, i.e. stationary, level series as the ones most suitable for analysing. We also include real-wage IRFs, where variables are de-trended by CPI to determine the long-run effect of actual price influence[5]. We also take prices explicitly to obtain the impulse of CPI to nominal wages and the feedback that occurs.

Table 6. Real Wage Granger Causality in Logs

Granger Causality Wald Tests		
Equation	Excluded	Probability
Private Wage	Public Wage	0.546
Public Wage	Private Wage	0.002
Private Wage	Minimum Wage	0.562
Minimum Wage	Private Wage	0.599
Public Wage	Minimum Wage	0.562
Minimum Wage	Public Wage	0.760

Next, after fitting a VAR model, we estimate the Forecast Error Variance Decomposition (FEVD). FEVDs are used to determine how much of the forecast error variance of each variable is explained by exogenous shocks to other variables in the VAR. In short, FEVDs measure the relative importance of each shock or innovation that influences the respective wage. After trying different time horizons, we chose 12 forecast periods as sufficient to explain the shocks affecting wages.

5. Results on nominal wage IRFs and real and nominal FEVD tables can be found in Appendix C.

The last panel of Figure 3 on the next page shows the IRF of real private over public wages. This is the most significant graph in our study, since it confirms the results previously obtained from the Wald tests, namely that public wages respond to changes in private wages with a lag of approximately two (to two and a half) quarters. The first plot shows that a shock of real private wage to itself raises it, but quickly dies out and over time reaches zero. As for the shock of private on minimum wage, it can be disregarded as having no statistically significant effect. The last panel focuses on the impact the private wage shock has on the public wage; the shock starts from zero and reaches its peak in the second period, i.e. two quarters after the initial change. This means that public wages are affected by changes in the private wage with a delay of half a year.

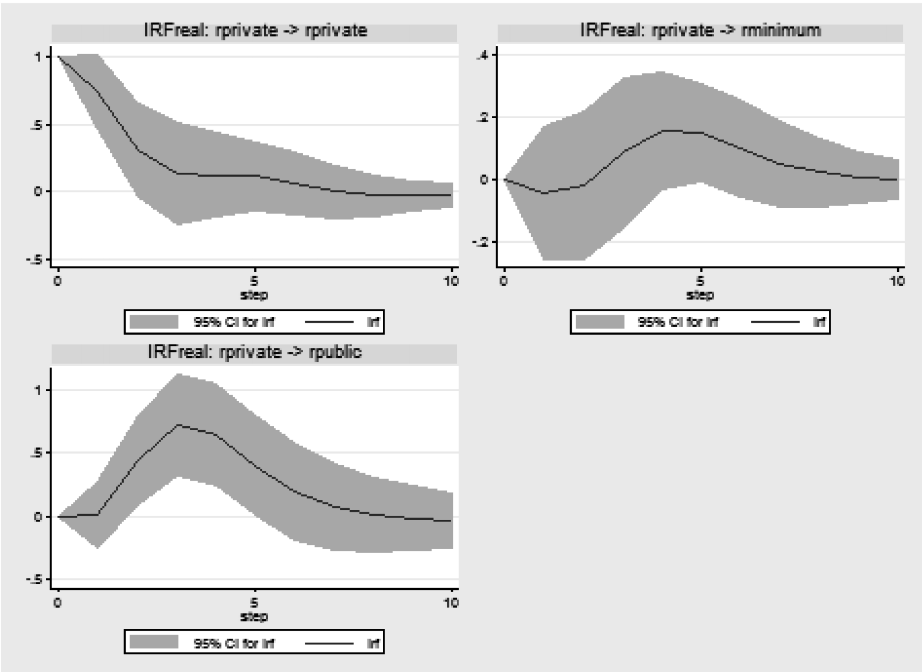

Figure 3. Private Real Wages

The source of this shock, epsilon private, can be attributed to the increase in innovations and productivity (TFP) associated with the private sector. This explanation is consistent with the main theoretical models in labour literature[6], where the wage rate is determined as an outcome of the bargaining process between workers and the firm. In this case, wage is considered to be proportionate not only to labour productivity, but also to the marginal rate of substitution and, more specifically, the shock could be driven by factors other than technological innovation, such as change in taste or leisure preferences or shocks to alternative income (changes in productivity in the non-market sector).

Similar reasoning for the case of real public wage in Figure 4 below shows that, as expected, at step zero, public wage increases by one unit and then slowly decreases to zero. In addition, a public wage shock has no statistically significant effect on minimum wages. Furthermore, the last plot in the figure shows that the response of the private wage has a higher coefficient but remains negligible. This shock can be attributed to policies such as an unexpectedly high tax revenue or drop in other costs, e.g., lower demand for public services, etc. Lastly, government policies of changing public wages do not have any noticeable effect on increases in private wages. The response in public sector wages is consistent with setups where wages are decided based on availability of government funds.

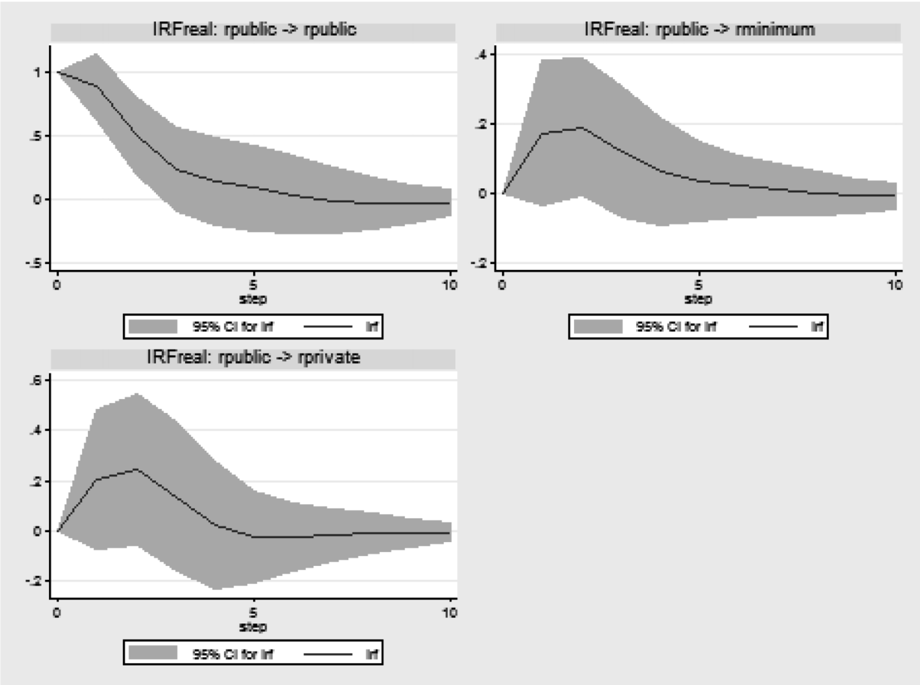

Figure 4. Real Public Wage

6. Discussions on possible models as explanations to epsilon public, private and minimum can be found in Appendix D.

Next, we study the effect of minimum wages on the other two wages. As shown in Figure 5 below, real minimum wage shocks have no long term effects on any of the three variables. This result shows that real minimum wage is still lagging behind productivity growth. Nominal minimum wages also display similar results to changes. Therefore, it can be concluded that the recent government policy of increasing the minimum wage in order to narrow the gap between minimum and average Bulgarian wages is of no substantial effect; the weak or overall lacking response elicited from the two sector wages is negligible over the one-year period shown. Furthermore, epsilon minimum, or the source of the shock, captures the effect of innovation of minimum wage to minimum wage, i.e., government policies aimed at wage increases. This may be used as an instrument to lower the percentage of grey economy, but results show that it is not effective as an income policy tool. A possible explanation is the prevalence of 'envelope' money (people declare minimum wage as an official income, but at the same time do not declare all income), which diminishes the possible effect of the minimum wage.

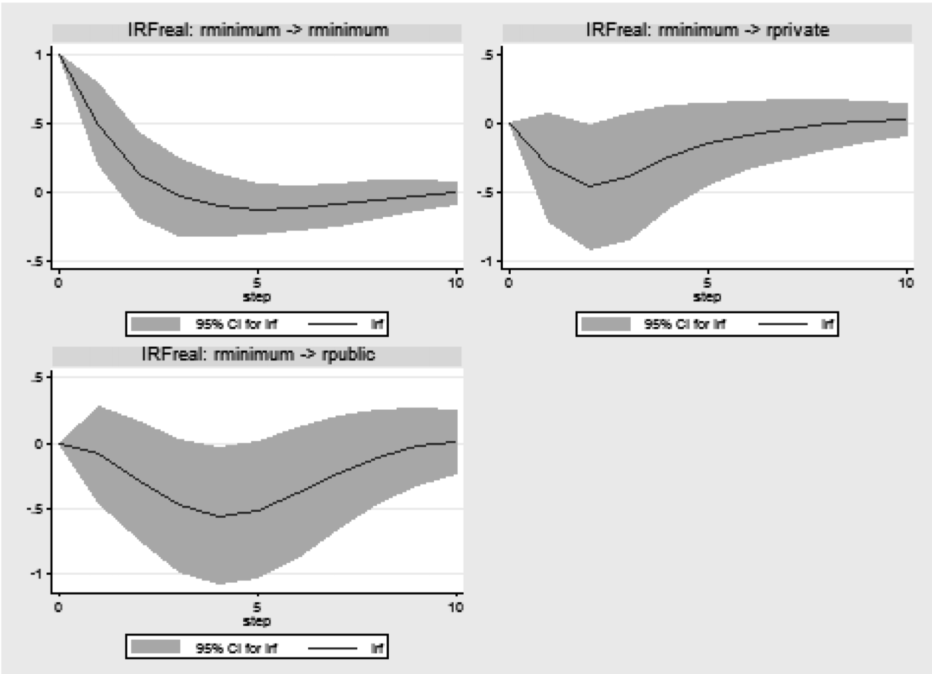

Figure 5. Real Minimum Wage

When we take the shock of CPI to nominal private, public and minimum wages, there is little feedback from them over a one-year period, as evidenced from Figure 6 below. The price shock to nominal private wages gets a statistically insignificant response for the first quarter and eventually dies out in the next period. The feedback from public and minimum wage is even more negligible, which indicates that CPI has no effect in the short run (4 quarters). Price-wage link in Bulgaria is, therefore, weaker when compared to other EU countries and wage changes in response to inflation are not as quick to occur, as documented by Lozev *et al.* (2011).

Wage stickiness may play an important role in price adjustment, which implies that wages in Bulgaria are changed mainly for reasons not linked to inflation (e.g., length of service). Furthermore, Vladova (2012) finds evidence from surveys indicating that only 27% of the firms in Bulgaria take into consideration the connection between prices and wages, as compared to 40% in the EU. Therefore, it can be concluded that price-driven wage changes are not common in the case of Bulgaria.

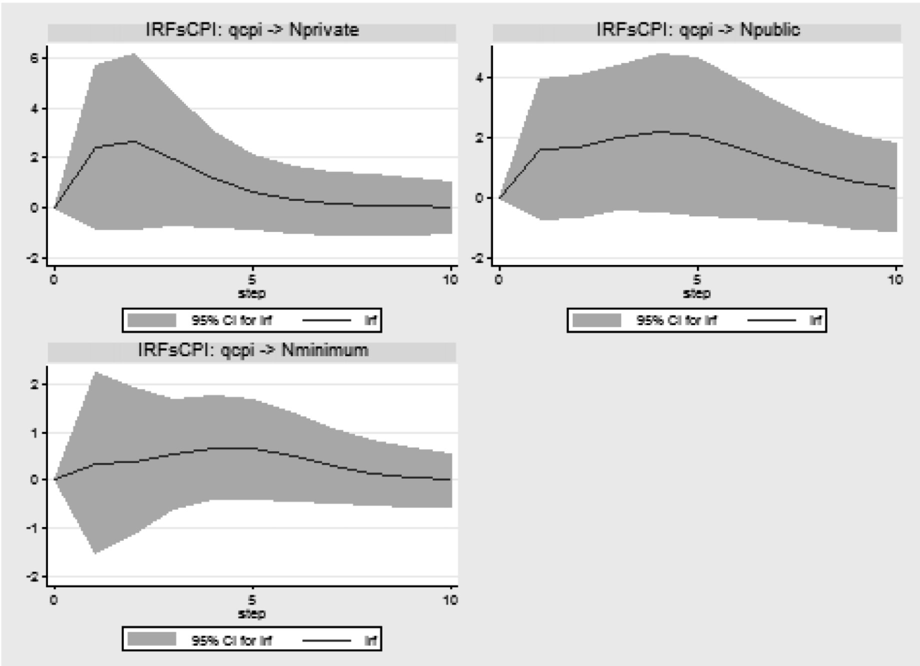

Figure 6. CPI on Nominal Wages

A possible explanation behind the shocks for the private, public and minimum wages is that they could be driven by monetary policy shocks. These shocks are a *de facto* increase in monetary base, or an increase in the circulation and reserve money, which can beneficially affect wages. An example of a one-time monetary shock that might lead to public wage increase is the 2016 budget surplus of 1.6 billion (Reuters). Moreover, Bulgaria is targeting a 1.4% deficit of GDP in 2017, which is expected to induce higher spending, possibly positively affecting labour costs as well (Reuters). Foreign inflows in Euro could also increase money supply (M2) when converted to BGN, in line with the 1.955 fixed exchange rate. Another example of foreign inflows are foreigners' time deposits, such as foreigners depositing their money or receiving their pensions through Bulgarian banks, etc. A factor beneficial to wages might be lowering the main interest rate, which was lastly recorded as 0% and has been close to zero since 2010 (BNB). The lower interest rate also attracts international companies and higher investment, which could be taken as a positive shock to nominal wages.

An oil price shock could be one factor affecting changes in private wages. The Bulgarian industry is energy-intensive and lower oil prices, as in 2016, can result in a positive shock to private wages. Since total costs would fall, capital and labour would be affected through the demand channel. Increased demand for cheaper oil would lead to lower labour costs and positively affect wages. A decrease in oil prices may, in addition, influence public wages by affecting wages in public utility companies. Another positive private wage shock is a possible trade shock, based on the export-led growth resulting from Bulgaria's accession to the EU in 2007. Many private businesses took advantage of the open EU borders and expanded their production beyond country-level, which generally results in a positive impact on private sector wages. In addition, the lower corporate tax rate (10% as of 2008) could also induce a positive wage shock, increasing the value of after-tax "surplus", divided between labour and profit income.

Minimum wage shocks could be the result of a positive shock to laziness, or the decreased preference to work. A shock to laziness, or relative preference for leisure, would trigger a substitution away from official work and towards either collecting benefits or transfers, or working in the grey economy. Further, receiving minimum wage also comes with children benefits, which can be a factor attractive to mothers. Other types of transfers that can induce this substitution effect are food vouchers, unemployment benefits, housing subsidies, transfers in kind (heating vouchers, electricity or other types of vouchers), etc. Another factor could be a shock to investment technology, such as higher inventory or lower installation costs. With more capital used, labour is more productive, so wages also increase. Next, a shock to capacity of input utilisation, i.e. capital and labour or elasticity of substitution between labour and consumption can also have an effect on wages. Alternative sources of income or consumption, such as home production, may increase consumption or lead to higher income. These shocks are all some possible explanations for our IRF results documented above.

Section 5: Conclusion, Limitations of the Study and Future Research

This paper studies the relationship between public, private and minimum wages in Bulgaria. It reports a strong correlation between public and private wages and finds further evidence of co-movement and long-term co-integration. Next, it demonstrates that causality runs from private to public wages and presents the implications from this finding. The study also focuses on policy recommendations based on these empirical results and reviews several labour models as possible explanations of the findings. What is new is the inclusion of minimum wage in the model and this is a contribution to existing literature. We report no causal relationship from minimum wages to public and private ones. This finding is important from a policy perspective since it poses the question whether minimum wage is relevant as a policy making tool. We conduct series of Granger-causality Wald tests on quarterly data and construct impulse-response functions to find the relationship and causal links between wages. In order to avoid spurious results, we use seasonally and cyclically adjusted variables after de-trending them. We consider both level and log forms to conduct our study and further use CPI as a main deflator in obtaining real wages. Following Lamo *et al.* (2008), we include CPI in the VAR with nominal wages to study possible price-wage linkages.

Next, we will consider some limitations of our study. This paper's added value is the inclusion of a minimum wage as an income policy tool in the private and public sector wage dynamics. We discuss some possible explanations for our findings in the Appendix, but it is outside the scope of our study to consider a fully specified model. Our focus is wage leadership and causality in terms of one variable forecasting another. Therefore, what we do not include is a theoretical model that would provide a deeper understanding of wage determination mechanisms.

This limitation opens venues for future research, as it would be extremely interesting to delve into the disciplined theoretical approach of Bulgarian wage dynamics. The current study can be extended to a micro-founded model to explore how wages change over time, rather than taking them in a Walrasian static equilibrium context. An interesting path to follow would be to specify a dynamic stochastic general equilibrium (DSGE) model based on optimisation and rational behaviour.

Finally, in the light of these results, we recommend implementing policies aimed at total factor productivity increases in the private sector to stimulate its growth and, consequently, growth in the public sector. Furthermore, this would strengthen the otherwise weak link between wages and labour productivity. Lastly, in the light of our findings, we recommend less reliance on minimum wage as an income policy tool.

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Appendix A

Section 4: Unit Root data on Log Forms

Table 1.2 Unit Root Tests on Wages in Logs

Variable	Mackinnon p-value	Trend p-value	Const p-value	Order of Integration (I)
Nom. Private Wage	0.6704	0.093	0.055	1
Nom. Public Wage	0.2581	0.016	0.009	1
Nom. Min Wage	0.3291	0.023	0.011	1
Real Private Wage	0.7681	0.166	0.074	1
Real Public Wage	0.7388	0.147	0.071	1
Real Min. Wage	0.8262	0.239	0.077	1

Section 4: VAR statistics

The tables on the next two pages should be read as follows:
PW – Private Wage
PU – Public Wage
MIN W. – Minimum Wage
L1 – First Lag   L2 – Second Lag
Coeff. – Coefficient (alpha)
SE – Standard Error

The VAR tables display the regression outcome of each dependent variable on lags of itself and lags of all other dependent variables (wages).

Table 7. VAR for Nominal Wages in Levels

	Coeff.	SE	P-value
Private Wage			
PW L1	0.862	0.130	**0.000**
L2	-0.246	0.136	0.072
PU W L1	0.156	0.178	0.382
L2	-0.053	0.166	0.749
MIN W L1	-0.083	0.236	0.725
L2	-0.027	0.243	0.814
Public Wage			
PW L1	0.047	0.092	0.604
L2	0.263	0.096	**0.007**
PU W L1	0.780	0.126	**0.000**
L2	-0.137	0.117	0.244
MIN W L1	0.153	0.167	0.358
L2	-0.121	0.172	0.481
Min. Wage			
PW L1	-0.008	0.072	0.903
L2	0.133	0.076	0.083
PU W L1	0.091	0.100	0.358
L2	-0.067	0.093	0.466
MIN W L1	0.595	0.132	**0.000**
L2	-0.102	0.136	0.452

Table 8. VAR for Nominal Wages in Logs

	Coeff.	SE	P-value
Private Wage			
PW L1	0.810	0.135	**0.000**
L2	-0.149	0.143	0.299
PU W L1	0.252	0.193	0.192
L2	-0.048	0.183	0.793
MIN W L1	-0.037	0.083	0.657
L2	-0.106	0.083	0.198
Public Wage			
PW L1	0.010	0.083	0.899
L2	0.331	0.088	**0.000**
PU W L1	0.837	0.119	**0.000**
L2	-0.246	0.112	**0.029**
MIN W L1	0.0243	0.051	0.635
L2	-0.091	0.051	0.073
Min. Wage			
PW L1	-0.089	0.224	0.690
L2	0.248	0.236	0.294
PU W L1	0.257	0.319	0.420
L2	0.108	0.302	0.719
MIN W L1	0.540	0.137	**0.000**
L2	-0.181	0.137	0.186

Table 9. VAR for Real Wages in Levels

	Coeff.	SE	P-value
Private Wage			
PW L1	0.920	0.142	**0.000**
L2	-0.307	0.145	**0.035**
PU W L1	0.090	0.184	0.626
L2	-0.009	0.170	0.957
MIN W L1	0.166	0.300	0.580
L2	-0.144	0.315	0.646
Public Wage			
PW L1	0.082	0.115	0.475
L2	0.107	0.118	**0.044**
PU W L1	0.779	0.150	**0.000**
L2	-0.171	0.138	0.218
MIN W L1	0.474	0.243	0.352
L2	-0.311	0.256	0.230
Min. Wage			
PW L1	0.011	0.066	0.865
L2	0.098	0.067	0.148
PU W L1	0.061	0.086	0.478
L2	-0.071	0.079	0.367
MIN W L1	0.691	0.140	**0.000**
L2	-0.121	0.147	0.408

Table 10. VAR for Real Wages in Logs

	Coeff.	SE	P-value
Private Wage			
PW L1	0.937	0.155	**0.000**
L2	-0.240	0.164	0.143
PU W L1	0.218	0.200	0.276
L2	-0.105	0.183	0.564
MIN W L1	-0.016	0.098	0.868
L2	-0.079	0.096	0.411
Public Wage			
PW L1	0.154	0.113	0.173
L2	0.211	0.120	**0.015**
PU W L1	0.844	0.146	**0.000**
L2	-0.395	0.134	**0.003**
MIN W L1	0.050	0.072	0.489
L2	-0.041	0.070	0.560
Min. Wage			
PW L1	-0.135	0.207	0.514
L2	0.222	0.219	0.312
PU W L1	0.117	0.267	0.662
L2	0.147	0.245	0.548
MIN W L1	0.569	0.132	**0.000**
L2	-0.192	0.128	0.135

Section 4: Granger-causality tests

Table 6.1 Nominal wages in level forms

Granger Causality Wald Tests		
Equation	Excluded	Probability
Private Wage	Public Wage	0.598
Public Wage	Private Wage	0.00
Private Wage	Min. Wage	0.835
Min. Wage	Private Wage	0.104
Public Wage	Min. Wage	0.632
Min. Wage	Public Wage	0.655

Table 6.2 Nominal wages in log forms

Granger Causality Wald Tests		
Equation	Excluded	Probability
Private Wage	Public Wage	0.257
Public Wage	Private Wage	0.00
Private Wage	Minimum Wage	0.188
Minimum Wage	Private Wage	0.547
Public Wage	Minimum Wage	0.174
Minimum Wage	Public Wage	0.547

Table 6.3 Real-term wages in level forms

Granger Causality Wald Tests		
Equation	Excluded	Probability
Private Wage	Public Wage	0.834
Public Wage	Private Wage	0.07
Private Wage	Minimum Wage	0.842
Minimum Wage	Private Wage	0.137
Public Wage	Minimum Wage	0.149
Minimum Wage	Public Wage	0.137

Appendix C

Nominal Impulse-Response Functions:

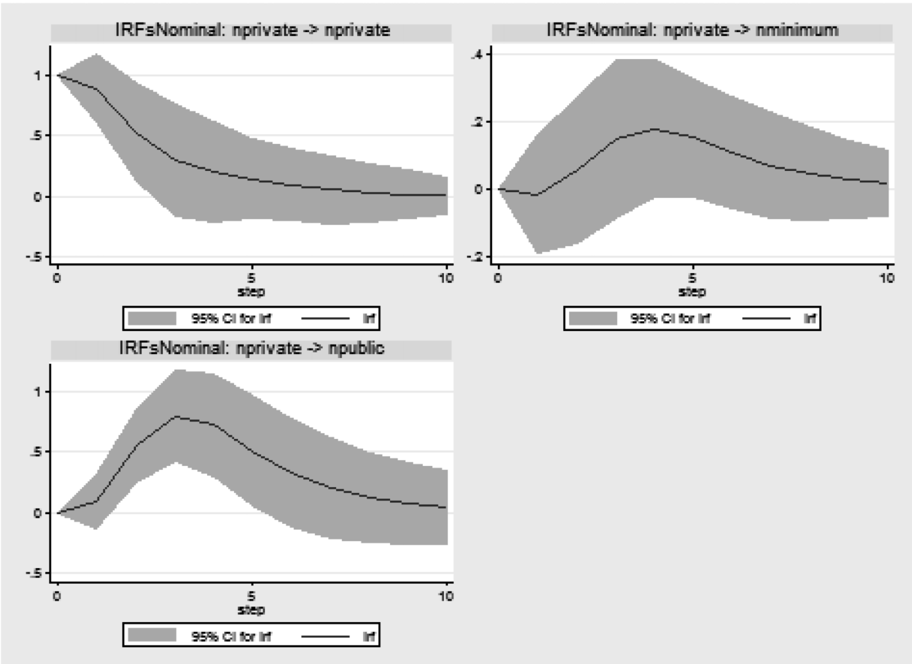

Figure 9. Nominal Private Wages

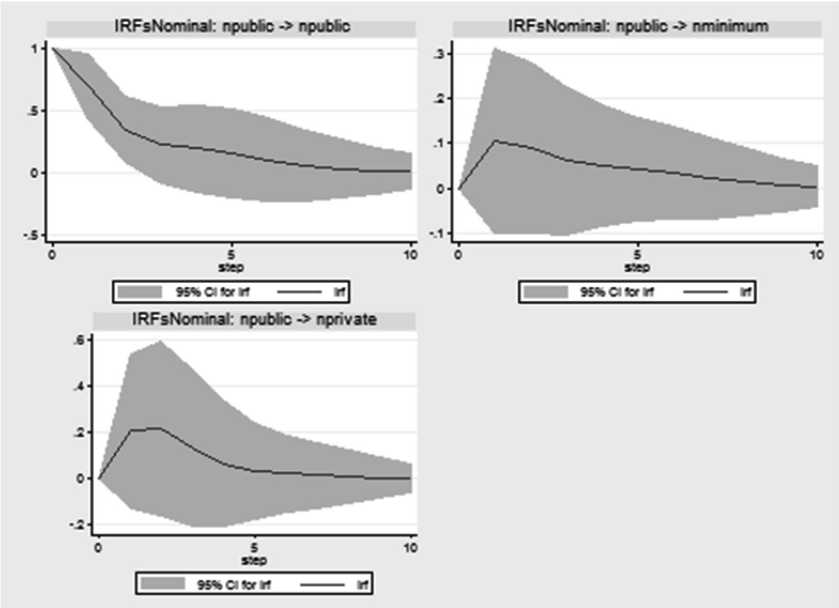

Figure 10. Nominal Public Wages

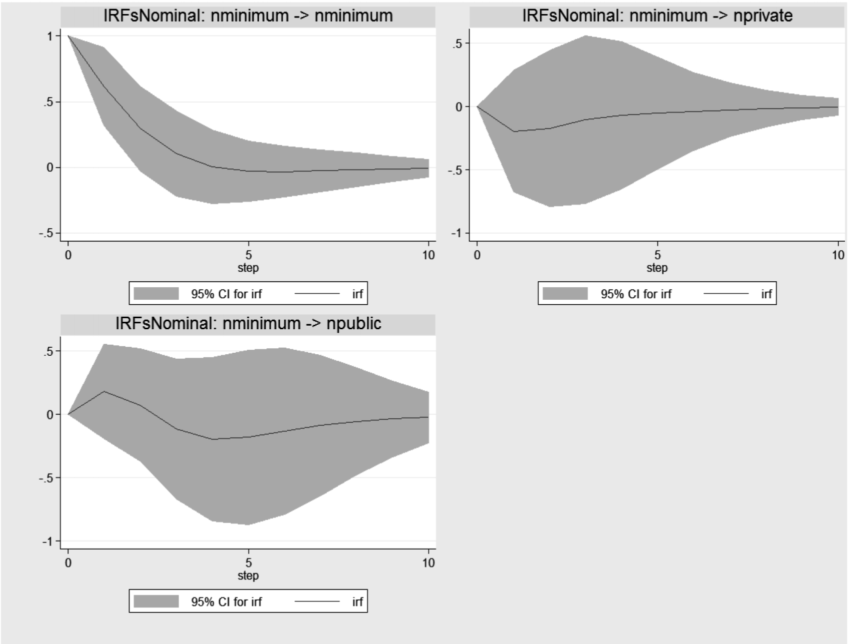

Figure 11. Nominal Minimum Wage

FEVD (Forecast Error Variance Decomposition)

Table 111. FEVD Nominal Private and Real Private Wage

Step	FEVD Private-Private	FEVD Private-Public	FEVD Private-Minimum	Step	FEVD Public-Public	FEVD Public-Private	FEVD Public-Minimum
0	0	0	0	0	0	0	0
1	1	0.095399	0.072028	1	0.904601	0	0.021455
2	0.993514	0.124059	0.074906	2	0.868071	0.005388	0.043451
3	0.984384	0.269065	0.13016	3	0.723478	0.011418	0.045992
4	0.977676	0.426127	0.216028	4	0.568288	0.014796	0.042944
5	0.974103	0.516344	0.275185	5	0.477878	0.016213	0.040661
6	0.972494	0.55642	0.300895	6	0.43619	0.016787	0.039968
7	0.97182	0.572006	0.309109	7	0.418737	0.017041	0.039985
8	0.971544	0.577502	0.311156	8	0.411797	0.01716	0.04012
9	0.97143	0.579305	0.311634	9	0.409112	0.017217	0.040255
10	0.971383	0.579876	0.311645	10	0.408084	0.017244	0.040276
∞	0.971362	0.580063	0.311645	∞	0.407686	0.017256	0.040285

Table112. FEVD Nominal Public and Real Public Wage

Step	FEVD Minimum-Minimum	FEVD Minimum-Private	FEVD Minimum-Public	Step	FEVD Private-Private	FEVD Private-Public	FEVD Private-Minimum
0	0	0	0	0	0	0	0
1	0.906517	0	0	1	1	0.272742	0.148299
2	0.881643	0.001099	0.00787	2	0.994426	0.331508	0.171599
3	0.823848	0.004198	0.007457	3	0.98718	0.407196	0.239792
4	0.741029	0.007527	0.005585	4	0.982506	0.478459	0.314881
5	0.684154	0.009684	0.005778	5	0.98069	0.524724	0.362743
6	0.659137	0.010719	0.00739	6	0.9802	0.547103	0.383199
7	0.650905	0.011139	0.009257	7	0.980076	0.555381	0.389196
8	0.648724	0.011295	0.010701	8	0.980033	0.557674	0.390306
9	0.648224	0.011352	0.011583	9	0.971383	0.579305	0.311566
10	0.648111	0.011373	0.01204	10	0.971362	0.579876	0.311634
∞	0.648079	0.011381	0.012251	∞	0.971353	0.580063	0.311645

Table113. FEVD Nominal Minimum and Real Minimum Wage

Step	FEVD Public-Public	FEVD Public-Private	FEVD Public-Minimum	Step	FEVD Minimum-Minimum	FEVD Minimum-Private	FEVD Minimum-Public
0	0	0	0	0	0	0	0
1	0.727258	0	0.045876	1	0.805825	0	0
2	0.638017	0.003138	0.065901	2	0.7625	0.002435	0.030475
3	0.553634	0.008839	0.061426	3	0.698782	0.003981	0.03917
4	0.483889	0.013067	0.054503	4	0.630615	0.004428	0.037652
5	0.440385	0.014869	0.050479	5	0.586778	0.004441	0.034891
6	0.41963	0.015372	0.048834	6	0.567967	0.004428	0.033267
7	0.412007	0.015477	0.048409	7	0.562394	0.004447	0.032612
8	0.409903	0.015499	0.048378	8	0.561316	0.004468	0.032423
9	0.409112	0.017217	0.04021	9	0.561207	0.00448	0.032395
10	0.408084	0.017244	0.040255	10	0.561207	0.004486	0.032403
∞	0.407686	0.017256	0.040276	∞	0.561205	0.004488	0.032414

Appendix D

In this part of the Appendix we analyse several labour market models in the context of our findings. These models can rationalise the wage dynamics observed and their response to shocks as documented with impulse-response functions. We analyse labour market models for the private and public sector separately.

Private Sector

Perfect Competition

Perfect competition models assume that, in equilibrium, wages equal the marginal product of labour (MPL) on the firm side. Some of the aspects of the model suppose that everyone has perfect information on prices and that firms have limited market power in influencing prices and wages, i.e. firms are small and non-influential. Perfect observability is another assumption of the model. On the consumer side, wage equals the marginal utility MU of labour. Another aspect is the idea of free entry and exit of firms, as well as their ability to adopt technology without any cost incursions. The perfect competition model is particularly useful in its capacity to measure output and effort, something not as easily captured in the cases below. Despite being unrealistic, this restrictive model is a useful starting point when studying labour markets.

Monopsonies

In the general model, monopsonies, or single buyers of labour, have the purpose of maximising profits. In order to do that, employers would offer wages below the competitive equilibrium as compared to that of perfect competition. Paying a lower wage also means that the wage is below MPL, because monopsonists hire according to the marginal cost MC and not the labour supply curve. If a monopsonist hires a marginal worker at a higher wage, s/he needs to increase the wages of all those hired previously. In equilibrium, a monopolist employs fewer people and pays lower wages. This, combined with the limited power of labour unions in the private sector (as pointed out in section 3), is a plausible explanation of the lower private sector wages in Bulgaria. Furthermore, data on Bulgaria show little mobility between regions (NSI 2016). Therefore, this model is credible if we assume that little labour mobility holds; if a monopsony pays less than the market-clearing wage, it is natural that workers would change jobs. Additionally, state-owned monopsonies were privatised in the 1990s, which resulted in forming regional monopsonies capable of keeping wages artificially lower.

Unions

In general, labour unions bargain for higher wages or better working conditions for workers. As discussed previously, in Section 3, private sector labour unions can be viewed as having limited power, since their demands are constrained when bargaining for wages higher than the marginal product in a profit-maximising firm. Unions can achieve wages greater than MPL, for instance, at the expense of decreasing the share of capital income; however, this will still be within the profit constraint in order to avoid bankruptcy. In the presence of monopsony, unions playing the same role as minimum wage mitigate the excessive monpsonistic power, as documented by Boeri and Ours (2008), which means that unions would not allow monopsonies to pay workers below a minimum or a set wage. However, unions have limited potential to explain labour market dynamics in Bulgaria.

Search and Matching

In this model, we consider a labour market with search and matching frictions. Such frictions are primarily the result of having imperfect information or information asymmetry in the labour market. Search and matching frictions generate externalities, since it takes time for jobseekers to be matched to a position. In this non-Walrasian model, equilibrium is determined by the demand side, while the number of people hired is determined by the labour quantity demanded. This model could, in addition, account for the presence of an involuntary level of unemployment. Furthermore, individuals might not always take the prevailing wage, while finding the best match is often a long and costly process, in terms of both time and resources. Another possibility that accounts for the persisting level of unemployment might be that individual workers differ in terms of skills/competences and it may often be difficult to find a perfect match. Therefore, search and matching frictions seem to be quantitatively important in explaining labour market dynamics in Bulgaria.

Efficiency Wages

Efficiency wages in labour economics denote the tendency of some employers to pay more than the market-clearing wage in order to encourage higher productivity. Efficiency wages are often used as a 'gift exchange', where the employer pays higher than the equilibrium wage in order to induce more effort from the worker. Such wages can also be viewed an incentive for semi-skilled and skilled labour. This model could also explain the disparity between public and private wages in the context of real rigidities. Rigidities are prices or wages that do not adjust to the expected equilibrium level in response to changes in other prices or wages. Additionally, if private sector wages are seen as rigid, they may change only in response to technological shocks and account for the percentage of existing involuntary unemployment.

Therefore, this would explain the lower private sector wage, which fails to adjust to wage increases in the public sector. Another explanation for the disparity may be the practice of 'envelope' wages by some employers who prefer to complement the officially documented wage with undeclared cash compensation. Furthermore, real rigidities and efficiency wages imply that a worker's effort depends on the real wage and to maximise profit, firms choose the real wage that would induce the highest effort on the part of workers. However, we also discussed the weak price-wage link, pointed out by Vladova (2016), which means that real wages do not adjust as often as prices change. Finally, efficiency wages can also potentially explain wage dynamics in Bulgaria.

Minimum Wages

The minimum wage model may be applied to attract people from the grey economy, especially when the percentage of the latter is too high in the overall economic activity. A minimum wage can, in this sense, be viewed as an efficiency wage that aims to deter people from shifting to the grey sector by offering an incentive to seek a job in the private sector. Another possible explanation is that a minimum wage is used to increase labour productivity. However, as productivity is difficult and, often, costly to measure and monitor, it is not always clear whether a minimum wage is an effective tool for increasing productivity. Moreover, the higher the minimum wage, the stronger a company's incentive to either fire workers or use 'envelope' wages, which is counter-productive. According to the National Revenue Agency (NRI), one-third of the Bulgarian working population declared working for the minimum wage in 2016. Besides, recent increases in minimum wage are not in line with productivity increases, again according to IME (2015) information.
We can also express the models of private wage with a Nash bargaining equation:
...................................,
where .... is the bargaining weight of the firm multiplied by the marginal product of labour MPL. Marginal utility MU could be viewed as a function of the stochastic taste shift parameter (as a source of epsilon private). MU can also be attributed to taste shocks, such as a change in preferences for higher home production or leisure preferences. It can also work as an outside option, such as people choosing to work in the grey economy or to receive unemployment benefits. When ..... equals one, the model denotes perfect competition and when it equals zero, wage equals marginal disutility of labour or an outside option, such as unemployment benefits or MPL generated in the grey economy. In the search and matching frictions model, gamma would equal 0.5.

In a perfect competition model, wage, marginal product of labour and marginal utility of labour would be equal. In union and search models, however, wage would differ from either MPL or MU. Furthermore, the models of monopsonies and unions could account for the lower wages in the private sector if the reverse causality were true, i.e. if the public sector was the wage leader. However, since private wage changes lead to public wage changes, the model becomes less plausible. Next, going back of the IRFs, we concluded that shocks in CPI do not affect either wage in nominal terms. Additionally, wages in Bulgaria change frequently as compared to the rigid model of efficiency wages. Therefore, the models of both efficiency wages and a minimum wage as an efficiency one cannot fully explain the findings because of the weak price-wage link. However, the efficiency wage and search and matching frictions models still capture business cycles better than the other models, as documented by Vasilev (2016) and Vasilev (2017).

Public Sector

What follows are some possible theories that explain the consistently higher wages in the public sector. The discussion in section 3 pointed at some privileges when working in the public sector, such as having state-financed employee social contributions, higher after-tax income, due to government payment towards pension, over-representation of women because of the contrary happening in the private sector, etc. The public sector's objective can be to maximise employment, due to social considerations, or gain more votes through public employment. As there are not as many theories that could help explain our findings about the public sector, we took the private sector models and will now look at them from the public sector aspect. However, there is no best model to be selected in the case of the public sector.

In the government sector there is no profit motive, so perfect competition is not a good approximation to the problem of pricing labour in the public sector.

Unions and a single buyer of labour, i.e. the government

The government is a single employer of labour that determines public sector wages according to the government budget constraint. Furthermore, since some jobs exist exclusively in the public sector, the government is a single employer, or monopsonist that exclusively sets wages, as, for example, in the case of the railway (BDZ), the police, etc. As documented by Borjas (2013), unions in the public bargain for higher wages and better working conditions generally have more power than in the private sector, since they are not restricted by profit-maximisation. However, as union power is not as strong as it used to be, this model can be disregarded as less explanatory than others.

Search and Matching

Search theory suggests that people with the same abilities may often end up at differently paid jobs in the process of matching, due to, for example, favouritism, political considerations or information asymmetries. The model could explain the wage premium in the public sector. However, public sector employees can also be viewed as risk averse, since government positions are relatively secure compared to those in the private sector. Working in the government sector can also be an occupational choice if people find it more rewarding to work for the public good or for their country. Government employees are also less likely to change positions frequently or to quit their jobs, so the process of search and matching occurs less often, if not only once, as opposed to the frequency of job changes in the private sector. We conclude that search and matching frictions do not explain the labour market as well as in the private sector.

Efficiency Wages

The efficiency wage model is a possible explanation for the wage premium of the public sector. Furthermore, the government includes a rent allowance in wages to establish not only loyalty from its employees, but also a reputation as a good employer. Moreover, the public sector tends to attract people with higher education or those with advanced qualifications, which may account for their higher wages. Besides, employees tend to have more experience than their counterparts in the private sector and, in general, have higher long-term benefits compared to the quick money incentive in the private sector. The additional rent allowance that government employees receive is *de facto* an efficiency wage shared between employees and the government.

Minimum Wages

The official Bulgarian minimum wage will reach BGN460 in 2017; one reason for its continuous increase for the past several years is the government's attempt to fight the increasing percentage of the grey economy. Friedrich Schneider finds that the percentage of grey economy for 2015 is 30.6%, a figure that has increased in the last year. The Institute for Market Economics (IME) reports in 2015 that, for every 100 BGN the minimum wage grows, 1.4% decline of employment follows. However, as our findings indicate, the minimum wage is not an effective tool or a sufficient model, not only because it introduces unemployment, but also because it fails to reduce the percentage of the grey economy.

Political Economy Factors

As noted in Section 3, Rose (1985) points out that public sector workers are also voters and are often affiliated with a particular party in order to be patronised by it. Public wages and, in this case, minimum wages can be increased prior to an election to encourage people to vote for a particular party. This would also explain discrimination between workers with similar abilities and characteristics, as well as the higher public wage. There is also a theory, known as Parkinson's law (Parkinson, 1955), that bureaucracy self-breeds and creates an increasingly complicated bureaucratic system in order to hire more people (subordinates). This theory might explain the ever-growing and larger public employment, as well as the wage premium in the presence of political economy in Bulgaria. However, the cycle hierarchy described is mitigated by having finite finances for wages coming from the private sector in the form of taxation. This also means that public employment is a function of what happens in the private sector and is, thereby, logically following private wage changes.

Overall, these theories are in line with the observed IRFs, which show that private wage is the driver and that public wages respond to changes in private wages.

ASECU South-Eastern Europe Journal of Economics 1 (2019) 129-153 SEEJE

MICROECONOMIC DETERMINANTS OF PRIVATE RETIREMENT SAVINGS: THE CASE OF TUNISIA

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Abstract
In the last decade, the financial situation of pension funds in Tunisia has worsened because of demographic and economic factors. These changes will make solidarity between generations more difficult and will prompt the government to institute some reforms regulating basic pension. This scenario of probable changes involves voluntary contributions to retirement saving schemes offered by financial institutions in order to supplement future pensions of workers. The aim of this paper is to understand Tunisian working population's savings behaviour, referring, for this purpose, to microeconomic factors, while assessing the current status of life insurance and funded pension schemes in Tunisia. To this end, a survey[1] was carried out on a representative population of Tunisian individuals affiliated to the three main public pension schemes and approached through savings theories and the Life Cycle Model (Modigliani 1954), in particular.

JEL Classification: G22, D14, D15
Keywords: Life Insurance, Retirement Savings, Life Cycle Theory

We would like to thank Mr. Mongi Hmidi, Deputy Director of the Unit of Actuarial and Financial Studies at CRES, for his expert advice and his valuable support in carrying out this study. Thanks also go to Mr. Nizar Oueriemmi, statistician engineer at CRES, for providing us with the databases of workers and retirees of different statuses and to Mrs. Jouda Khemiri, Director of the General Insurance Committee and to Pr. Malek Saihi for their helpful comments and suggestions. We would also like to thank the entire MOBIDOC-PASRI team for ensuring the smooth running and funding of this research.

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1. This study was carried out in collaboration with CRES Research and Social Studies Centre, Tunisia (Ministry of Social Affairs), under the MOBIDOC convention funded by the European Union, under the PASRI programme, and administered by the ANPR.

I- Introduction

In Tunisia, the pay-as-you-go pension scheme is facing real financial difficulties because of critical demographic trends leading to an ageing population, and because of a difficult economic context, notably reflected in high unemployment rates and low economic growth. These factors and many others would make the number of retirees exceed the number of working people, rendering funded pensions a burden and inter-generational solidarity more difficult.

This increasing budgetary imbalance of public pension funds and the uncertainty surrounding the sustainability of pension plans has urged governments to put in place several tax and social incentives to promote funded pension schemes proposed in the form of individual or group life insurance plans offered by insurance companies. These incentives suggest that more assets need to be saved to ensure protection from uncertainty about the sustainability of basic schemes and from decreased retirement income (Modigliani 1954, Friedman 1957, Fisher 1973).

These funded pension products are managed by life insurance companies; they represent 60% of insurance activity worldwide and 11.5% in Tunisia. Indeed, despite the range of life insurance products offered on the Tunisian market, the turnover of the life insurance market consists mainly of premiums generated by the 'temporary life insurance' scheme generally required by lending financial institutions and of premiums on mortgages, equipment and leasing operations. In return, life insurance and funding (long-term insurance) remain scarce, despite the numerous tax benefits they offer.

Tunisian insurers offer two types of life insurance products to accumulate long-term savings, namely, pure savings contracts and mixed contracts (savings / pension).

- ***Savings contracts***: These contracts aim at accumulating capital or an annuity at maturity. The insured entrusts the insurer with managing the former's contributions and the latter will pay them benefits[2] at a predetermined date in the case of a living person or a capital sum in the event of the insured person's death. Insurers offer a variety of savings products, such as: life insurance policies, capitalisation contracts and account-linked contracts.
- ***Mixed contracts***: these are contracts that cover both pensions and savings, allowing for the payment of a capital sum in the case of a living person at the end of the contract or the beneficiaries in the case the insured dies before the contracted period ends.

2. Offered services may be a capital sum or a lifetime annuity (with the option of spouse coverage for those who wish to do so). The insured chooses either exit modality (capital or annuity). In Tunisia, the exit modality is almost always that of a capital sum.

These life insurance and pension savings products are then offered by life insurers under general conditions, guarantees and benefits that meet the needs of different categories of individuals and businesses. They can be purchased as individual or group policies.

This sector is experiencing real growth and offers promising prospects for the Tunisian market, mainly thanks to favourable tax measures when subscribing to a Life product. Moreover, the amount of premiums generated by this sector has been steadily growing for several years (109 MD in 2014 as compared to 90MD in 2013 and 80 MD in 2012). Nevertheless, distribution of these premiums by insurance sector shows that the share of life insurance and capitalisation contracts in the Tunisian insurance market remains low (17.36% in 2014 as compared to 16.6% in 2013 and 16.4% in 2012).

Several factors may explain why this branch remains at such a bottleneck. On the one hand, there are economic factors, such as a low household savings capacity, inadequate investment in other financial instruments and economic instability. On the other hand, there are psychological and ideological factors, such as lack of trust in insurers, persistence of religious beliefs and lack of information about life insurance and funded pension schemes (Spaenjers 2012, Jamuldin 2013).

Public authorities also play an important role in developing life insurance, as it is an important vehicle of savings collection in the medium and long term. Tax exemption should also put forward a set of prudence rules and regulations concerning insurance investments. The development of this branch remains one of the main challenges of the sector.

Moreover, the generous pension plans hinder the development of life insurance and pension products (Feldstein 1979). Indeed, in Tunisia, a very large proportion of the active population is covered by pension and health schemes that offer attractive coverage rates, which does not encourage individuals to seek additional coverage from life insurers.

Beyond the scope of financial stability, understanding the economic and social specificities of the active population will help shape the potential development of this sector, hence, our interest in studying the socio-demographic, economic and financial determinants that may influence an individual's investment and savings strategies (Caroll 1992, Attanasio & Weber 1995, Lavigne & Freitas 2002).

Extending research on the issue of public pension schemes in Tunisia (Ben Brahem (2004), Houssi (2009), Chekki (2013)), the aim of this paper is to empirically determine the factors disengaging Tunisian assets from private pension schemes in the form of individual or group life insurance products. Despite the large number of studies examining the pay-as-you-go pension scheme in Tunisia, no empirical study has directly studied the factors that determine private funded pension schemes in Tunisia, proposed in the form of life insurance products for retirement.

The study first reviews the different theoretical and empirical studies that used the life-cycle hypothesis and examined the determinants of private life insurance and private pension schemes. Next, we present the model to-be-estimated, the data used and the methodology to determine the impact of socio-demographic and financial factors of Tunisian workers[3] on disengaging from life insurance for retirement in the four samples studied (Entire Population and three sub-groups RSNA[4], RINA[5] , CNRPS[6]). We will then discuss the main results, while the last section concludes the paper.

II- Literature Review

II.1- Life cycle theory and its extensions

Household savings and investment behaviour represent a major financing instrument for the economy and are affected by permanent changes in financial markets, innovative savings products in addition to an ageing population. Studies have shown that saving is a difficult concept to comprehend and that it depends on different macroeconomic and microeconomic determinants and motives.

The main theory about savings behaviour is the life cycle theory (Modigliani & Brumberg 1954), which assumes that saving is a way to finance consumption in one's old age. Typically, individuals borrow money before becoming economically active, save during the activity period and consume their savings at retirement.

Parallel to this theory, Fisher (1930) and Friedman (1957) developed the permanent income theory, which retains the same retirement savings motive but adds that an individual also has the desire to leave some inheritance to their descendants, a motive known as the generational transmission motive.

Several economists extended the life cycle model, such as Kotlikoff & Summers (1981), who assume that households transfer resources to their descendants, contrary to the life cycle theory, which supports that households ultimately consume all their resources.

Moreover, purchasing a life insurance policy is considered to be a form of savings, justified by the life cycle model. By extending the life cycle model, Fisher (1973) shows that the desire to transfer inheritance has some effect on holding life insurance contracts. Yaari (1975) and Campbell (1980) suggest that, just as household savings are motivated by a concern for distributing income across different life periods, so as to sustain more regular consumption patterns, purchasing life insurance policy aims at providing additional income or capital at a certain point in time.

3. Public and private sectors.
4. non-agricultural employees' scheme in the private sector
5. non-agricultural, self-employed in the private sector
6. public sector employees' scheme

Knowing the economic, legal and financial environment is a crucial piece of information before deciding on purchasing savings contracts. Macroeconomic factors that influence pension schemes mainly include taxation (Feldstein 1995, Arrondel & Masson 2003, Attanasio 2004) and the nature of the public pension scheme (Feldstein 1979, Leiner & Leroy 1982, Kantor & Fishback 1996).

II.2- Socio-demographic determinants of the savings behaviour

Beyond the economic and fiscal environment, household savings will also depend on socio-economic, geographical and ideological characteristics. Empirical studies carried out in several countries have focused on microeconomic determinants that affect holding pension savings policies. The main conclusions drawn from these studies indicate that savings behaviour depends as much on economic development as on individual factors related to their resources, preferences and expectations. These factors include life cycle variables, such as age and income, which, all things being equal, have a significant effect on purchasing pension contracts (Browne & Kim, 1993; Arrondel, 1996; Munnell *et al.*, 2000; Dauriol, 2005; Brun- Schammé & Duée, 2009). These authors concluded that younger households hold fewer such contracts than middle-aged households. Nevertheless, older households have Life Insurance holding rates as low as those of young people. Therefore, holding life insurance policies depends on the individual's generation. Under the life cycle theory, Brun-Schammé & Duée (2008) found that high-income households frequently hold more long-term savings. The income effect reflects the household's savings capacity and proves that poor households find it more difficult to access financial markets and diversify their portfolios.

Similarly, some authors suggest that marital status and household structure affect savings behaviour and investment choices. This amounts to saying that a change in the number of household members during one's life cycle may influence consumption and, consequently, savings behaviour (Yaari, 1965; Fischer, 1973). In a similar line, Fournier V. & Vaillancourt F. (2011) deduce, from their study, that married, divorced or widowed individuals save significantly less than unmarried individuals.

Furthermore, several authors argue that creation of a real estate asset, mainly through the acquisition of a real estate property, is considered to be a factor that affects one's decision to purchase life insurance policies and contribute to pension schemes (Bosworth, Burtless & Sabelhaus, 1991; Burbidge & Davies, 1994; Brun-Shammée & Duée, 2008; Arrondel & Savignac, 2011). These authors state that homeowners save in pension products more than renters do. For many households, acquisition of a residence remains the central project of their wealth accumulation. Other types of savings and investments come second. There are several studies examining the relationship between one's vocation and their decision to purchase this type of products (Brun- Schammé & Duée, 2009; Marti, 2011; Fournier V. & Vaillancourt F, (2011).

The studies indicate that household obligations affect their ability to plan for financial investments and they opt for asset accumulation.

Moreover, Marti (2011) states that socio-demographic (age, social status, income, etc.) and economic differences between regions shape household wealth strategies and, consequently, savings behaviour across regions. Dauner, I (2002) notes that geographic location and the proximity of financial institutions may affect the probability of households holding savings products.

The contribution of this study is to further examine the determinants of holding funded pension schemes in the Tunisian context, referring to the literature mentioned above.

III- Methodology

In what follows, we present, successively, our data, the empirical model to-be-estimated and the estimation method.

III.1- The Data

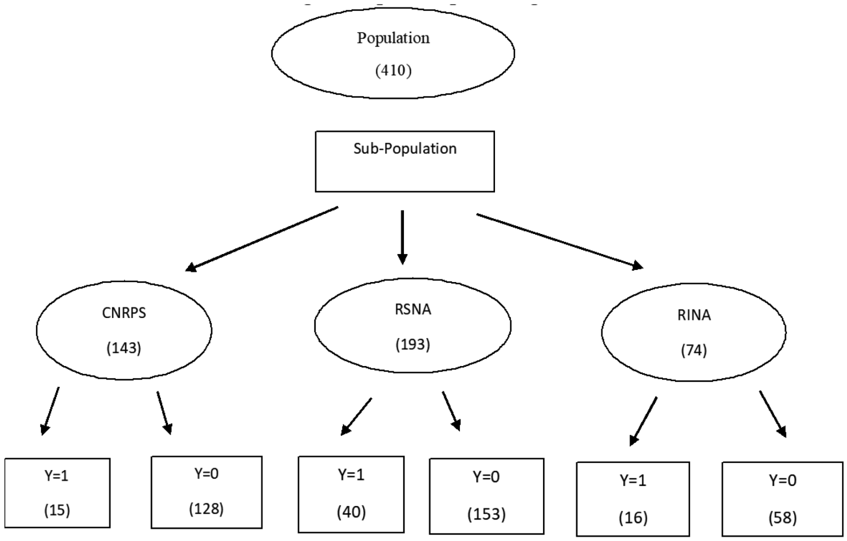
The main difficulties facing the analysis of household savings products retention relate to observations and quantification. Indeed, there are no accurate microeconomic data on pension schemes and life insurance purchasing behaviour. At this level, we would like to highlight that our study originally contributes to the research conducted on savings and life insurance in Tunisia (Dahmane D. 2000[7]; Khabcheche 2005[8]). Our study reports on a survey conducted in 2015 on a population of Tunisian workers affiliated to private and public sectors. The sample consists of male and female workers in different vocations, belonging to different age groups, with different income packages, and living in distinct geographic areas. The sample is distributed among the three main pay-as-you-go public pension schemes, namely, the Civil Service Regime (CNRPS), the Private Non-Agricultural Employees Plan (RSNA) and the Non-Agricultural Self-Employed Plan (RINA).

The representative sample is then divided into three sub-samples according to the pension scheme (private or public) and to activity type (employee or self-employed)[9]. They consist, respectively, of 142 assets belonging to the public plan (CNRPS), 197 assets to the Private Non-Agricultural Employees Plan (RSNA) and 76 to the Private Non-Agricultural Self-employed Plan (RINA). The survey[10] was conducted at the Research and Social Studies Centre.

7. Dahmane, D., 2000. Financial Liberalization and household saving behaviour: the case of Tunisia. Working Paper. IRD. University of Tunis 3.
8. Khabcheche (2005) « The Echo of Life Insurance in Tunisia », FTUSA Letter, December 2006- N°8.
9. CNSS or CNRPS
10. This study is carried out in collaboration with the Research and Social Studies Centre CRES Tunisia, under the MOBIDOC convention funded by the European Union, under the PASRI programme and administered by the ANPR.

The figure below shows the distribution of the entire sample and the sub-samples, specifying Life Insurance holding rates (Y = 1 or Y = 0).

Figure 1. Sample data processing



Source: Household Wealth survey 2015.

Our study tries to identify the determinants of holding Life Insurance contracts in the entire sample and according to the defined sub-samples. Details about the sample and sub-samples in terms of the different socio-demographic and financial variables mentioned above are presented in Table 1 below.

III.2- The model

Private retirement savings behaviour and Life insurance holding can be described by a discrete random variable $\tilde{Y}$, written as follows:

$$Y=\begin{cases}1 & \text{where 1 indicates holding a Life Insurance contract and 0 having} \\ 0 & \text{no pension scheme}\end{cases}$$

If we denote Y*i, j the probability that an individual holds a Life insurance contract can be formally written as follows: $Y^*_{i,j} = X_{i,j}\,\beta + \varepsilon_{i,j}$

i: Tunisian assets
j: worker affiliation (affiliated with one of the three pension plans (RINA, RSNA, CNRPS))
$X_{i,j}$: a vector of the characteristics of individual i of category j,
β: vector of the parameters measuring the influence of characteristics
$\varepsilon_{i,j}$: error term measuring the influence of the characteristics of non-observed assets.

By choosing to apply a logistic regression, we make the assumption that error terms are independent across individuals and follow a logistic distribution. We will try to estimate the probability that an individual i of category j holds or envisions purchasing savings contracts, taking into account their socio-demographic and financial characteristics, as follows:

$$\Pi i,j = P[yi,j = 1/Xi,j]$$
$$= P[Ti,j > 0] = P[\varepsilon i,j > -X'i,j\,ß]$$
$$= P[-\varepsilon i,j < X'i,j\,ß] = g(X'i,j\,ß)$$

where g is the distribution function of error terms εi, j defined by:

$$g(w) = P(-\varepsilon i,j < w)$$

In this study, we choose to apply a logistic regression and we assume that εi, j are independent across Tunisian assets and follow a logistic distribution.

The Logit function is defined by:

$$G(\pi) = logit(\pi) = ln\,(\frac{\pi}{1-\pi}) \qquad with\ g-1\,(x) = \frac{e^x}{1+e^x}$$

The ratio $\frac{\pi}{1-\pi}$ is the 'odds' that express a rating. Logistic regression is interpreted as linear modelling of log odds, while the coefficients of the variables express odds ratios. In our case, the probability that a Tunisian asset i of a sub-population j holds life insurance or pension savings contracts is written as follows:

$$\Pi\ i,j = \Lambda(X'i,j\,ß) = \frac{1}{1+e-X'i,j\,ß}$$

Furthermore, the estimation of a logistic regression model that allows for determining the log value of the holding rate of life insurance is as follows:

$$Ln[\frac{\pi_{i,j}}{1-\pi_{i,j}}] = \alpha + X'i,j\,ß$$

III.3- The variables of the model

The table below presents the variables Xi, j that will be introduced into the model and that represent each individual in the sample.

Table 1. Variables explaining the probability of holding a pension savings cont

Economic and socio-demographic characteristics	Variable	Description and values
Age	**Age**	Number of years
Generation	 **Gen1** **Gen2** **Gen3** **Gen4**	Age range Generation 1 [20-29[Generation 2 [30-49[Generation 3 [50-59[Generation 4 >60 years
Gender	**Gend**	Man / Woman
Annual Income	**IncBr**	Annual Income Bracket < 10000 TND, [10001TND;20000TND[, [20001TND;50000TND[, >= 50000TND
Geographical Area of residence	**East** **West**	North East, Greater Tunis, Central East (Sahel), South East Central West, South West, North West
Professional Category	**PubExec** **PrivExec** **SelfE**	Senior Manager, Middle-grade Manager, Senior executive, Executive, Researcher/Teacher, Self-employed/Investor. All these categories are grouped into three variables according to the sector of activity (public, private or self-employed)
Matrimonial status	**Sing/ Wid/ Div**	single, married , divorced (ee), widow (er)
Household Size	**HSize**	The number of individuals in the household dependent on the worker
Property ownership	**PropOw**	Owner or not of a principal residence

Source: Household Wealth survey 2015.

III.3- Specification and Estimation of the model

When examining the correlation matrix of the independent variables representing the entire sample and the sub-samples, we notice significant correlation coefficients between the following pairs: age and generation, household size and single status, and East variable and West variables. Bearing on these correlations, we specify two logistic regression equations to dissociate the joint effect of these pairs of variables on the sample studied.

$$LN\ [\![\pi i,j/(1-\pi i,j)]\!]=\alpha+\beta_1\ Age_{i,j}\ +\ \beta_2\ Age^2_{i,j}\ +\beta_3\ Gend_{i,j}+\beta_4\ Div_{i,j}+\ \beta_5\ Wid\ _{i,j}+\ \beta_6\ HSize\ _{i,j}\ +\ \beta_7 PubExec\ _{i,j}\ +\beta_8 PrivExec\ _{i,j}\ +\beta_9\ SelfE\ _{i,j}+\beta_{10}\ East\ _{i,j}\ +\beta_{11} IncBr\ _{i,j}\ +\beta_8\ PropOw_{i,j}\ (1)$$

$$LN\ [\![\pi i,j/(1-\pi i,j)]\!]=\alpha+\beta_1\ Gen2\ _{i,j}\ +\beta_2\ Gen3\ _{i,j}\ +\beta_3\ Gen4\ _{i,j}+\beta_4\ Gend_{i,j}+\beta_5\ Sing_{i,j}+\ \beta_6\ Div_{i,j}\ +\ \beta_7 Wid_{i,j}\ +\beta_8\ PropOw_{i,j}+\beta_{10}\ PubExec\ _{i,j}+\ \beta_{11} PrivExec\ _{i,j}\ +\beta_{12}\ SelfE_{i,j}+\ \beta_{13}\ West\ _{i,j}\ +\beta_{14}\ IncBr\ _{i,j}\ (2)$$

Equation (1) tests the joint effect of age, squared age, gender, vocational category, property ownership, the East variable, marital status, household size and the annual income range. As for equation (2), it tests the joint effects of generation, gender, marital status, property ownership, the West variable, vocational category and the annual income range.

The table below reports the fit quality of the pension savings / life regression model and regression significance.

Table 2. Goodness of Fit of Logistic Regressions

	Equation	Population	Sub-Population		
			CNRPS	RINA	RSNA
Number of observations	1	396	143	73	193
	2	396	143	73	193
Maximum Likelihood	1	330.067	88.738	65.861	167.127
	2	324.342	85.309	65.861	167.127
The Likelihood ratio test X^2	1	39.366	5.682	10.915	24.375
	2	45.092	9.110	10.915	24.375
Sig(X^2)	1	.000	.011	.001	.000
	2	.000	.017	.001	.000
R^2 Cox&Snell	1	0.095	0.041	0.139	0.122
	2	0.108	0.065	0.139	0.122
R^2 Nagelkerke	1	0.156	0.082	0.213	0.191
	2	0.177	0.129	0.213	0.191
$X^2_{Hosmer\text{-}Lemeshow}$	1	11.110	3.598	1.090	.000
	2	8.320	1.720	1.090	.000
Sig($X^2_{HosmerLemeshow}$)	1	.196	.165	.580	1.000
	2	.403	.632	.580	1.000

(1) the joint effect of age, squared age, gender, vocational category, property ownership, the East variable, marital status, household size and the annual income range.(2) the joint effects of generation, gender, marital status, property ownership, the West variable, vocational category and the annual income range.

In the two pre-established regression models, we seek to identify the variables that would effectively predict the probability of purchasing a life insurance / pension savings policy. The Hosmer-Lemeshow test indicates whether there is significant difference between values predicted and values observed.

The significance of the HL results, (Sig (X^2 Hosmer-Lemeshow)> 0.1)), indicates, in our case, that the values predicted and observed are the same. Coefficients determining Cox and Snell and Nagelkerke allow for assessing the fit quality of the final model. The higher the R2 coefficient, the better the model fits the data. In light of this finding, it can be said that the two models with the included variables can significantly predict the probability of holding a life insurance / pension savings contract than models with only constants.

Regression significance

The results above are supported by the deviance test (D law) and the likelihood ratio test (lr) to establish model significance.

- The deviance test

The results of the deviance test are summarised in Appendix 7. The table shows that introducing the variables mentioned above into the model allows for better prediction of the probability of holding a pension savings product. Indeed, coefficients 3.950 and 4.273 are significant at $p <0.01$. Thus, combining the independent variables significantly predicts the dependent variable Y.

- The Likelihood Ratio test

For the likelihood ratio test two hypotheses are considered:

H0: all coefficients βi of the independent variables are zero,

H1: there is at least one non-zero coefficient.

To this end, we use the likelihood ratio test. The relevant p-value uses Chi-square to indicate the impact of each independent variable on the model. The results of the likelihood ratio are summarised in the table below.

Table 3. Model Fit, the Deviance Test (ANOVA)

ANOVA	*Modèle*	*Population Globale*	*RSNA*	*CNRPS*	*RINA*
D	**1**	*3.950*	*3.018*	*1.525*	*1.313*
	2	*4.273*	*2.999*	*1.330*	*1.473*
Sig(D)	**1**	*.000*	*.002*	*.130*	*.248*
	2	*.000*	*.001*	*.210*	*.171*

(1) the joint effect of age, squared age, gender, vocational category, property ownership, the East variable, marital status, household size and the annual income range.(2) the joint effects of generation, gender, marital status, property ownership, the West variable, vocational category and the annual income range.

By assessing model parameters, we can determine the contribution of each variable of the model. Coefficient 'A' indicates the direction of the relationship between the independent variable and the dependent variable. In our case, property ownership, income range and vocational category 'private sector' vary in the same direction as the probability of holding a pension savings scheme, whereas the 'single' marital status variable varies in the opposite direction. The 't' value informs about the contribution of each variable of the model. T-values and their respective p show the impact of the independent variables on the dependent variable.

In Appendix 4, we found that the variable 'Annual Income IncBr' has a significant impact on the probability of holding a life insurance with $t = 3.140$ and $p < 0.01$. The second variable that has a significant impact on the probability of holding a life insurance is the 'PropOw Property Ownership' with $t = 2.380$ and $p < 0.01$.

The third variable that influences the probability of holding a life insurance is the vocational category 'Executive in Private sector PrivExec' with $t = 2.233$ and $p < 0.05$ as well as the 'SelfE self-employed vocational category' with $t = 2.484$ and $p < 0.1$. However, the fourth variable that has a negative impact on the probability of holding a life insurance is the 'Single status Sing' variable with $t = -2.172$ and $p < 0.05$.

In what follows, we will present, first, the descriptive statistics and, then, the results of the regression applied to the entire Population and to the three sub-population groups (RINA, CNRPS and RSNA).

Table 4. Goodness of Fit of Independent Variables (Xi)

Variable	*Equation*	*Coefficient*	*Population*	*Sub-population*		
				RSNA	*CNRPS*	*RINA*
Constant	1	*A*	-0.330	0.749	-1.116**	-0.223
		t	(-0.930)	(1.283)	(-2.054)	(-0.200)
	2	*A*	-0.072	-0.167		0.243
		t	(-0.662)	(-1.268)		(0.448)
Age	1	*A*	0.013	-0.050*	0.058**	0.030
		t	(-0.721)	(-1.611)	(2.112)	(0.504)
Age²	1	*A*	0.000	0.001	-0.010	0.000
		t	(0.507)	(1.549)	(-2.026)	(-0.648)
Gend	1	*A*	0.030	0.066	0.014	0.055
		t	(0.752)	(1.090)	(0.242)	(0.451)
	2	*A*	0.035	0.061	0.031	0.063
		t	(0.879)	(1.014)	(0.534)	(0.526)
IncBr	1	*A*	0.075*	0.057**	0.049	0.163*
		t	(3.140)	(1.529)	(1.352)	(2.729)
	2	*A*	0.074*	0.056	0.059	0.126**
		t	(3.116)	(1.467)	(1.619)	(2.096)
PropOw	1	*A*	0.099*	0.180*	0.048	-0.061
		t	(2.380)	(2.854)	(0.763)	(-0.540)
	2	*A*	0.085**	0.167*	0.036	-0.081
		t	(2.055)	(2.678)	(0.563)	(-0.715)
HSize	1	*A*	0.009	0.207	-0.007	0.017
		t	(0.609)	(0.819)	(-0.283)	(0.419)
Sing	2	*A*	-0.118**	-0.096	-0.37	-0.225***
		t	(-2.172)	(-1.277)	(-0.503)	(-1.667)
Div	1	*A*	0.145	0.268***	0.305	0.032
		t	(1.604)	(2.020)	(1.633)	(0.871)
	2	*A*	0.112	-0.253	0.246	-0.007
		t	(1.239)	(-0.643)	(1.329)	(-0.038)
Wid	1	*A*			-0.157	
		t			(-0.705)	
	2	*A*	-0.165	-0.218	-0.127	
		t	(-0.772)	(-0.557)	(-0.559)	
PubExec	1	*A*	0.072		-0.211	
		t	(0.947)		(-1.298)	
	2	*A*	0.088		-0.192	
		t	(1.160)		(-1.169)	
PrivExec	1	*A*	0.167**	0.246*		
		t	(2.233)	(2.823)		
	2	*A*	0.185**	0.264*		
		t	(2.484)	(3.015)		
SelfE	1	*A*	0.120			-0.513
		t	(1.445)			(-0.865)
	2	*A*	0.139***			-0.310
		t	(1.678)			(-0.586)
Est	1	*A*	0.001	-0.106	0.071	-0.010
		t	(0.019)	(-0.911)	(0.889)	(-0.037)
Ouest	2	*A*	-0.009	0.101	-0.072	-0.144
		t	(-0.135)	(0.875)	(-0.037)	(-0.520)
Gen2	2	*A*	-0.73	-0.158	0.060	0.023
		t	(-1.222)	(-1.844)	(0.649)	(0.115)
Gen3	2	*A*	0.019	-0.026	0.187**	-0.032
		t	(0.230)	(-0.203)	(1.537)	(-0.140)
Gen4	2	*A*	-0.130	-0.135	-0.037	-0.351
		t	(-1.212)	(-0.722)	(-0.503)	(-0.946)

(1) the joint effect of age, squared age, gender, vocational category, property ownership, the East variable, marital status, household size and the annual income range.(2) the joint effects of generation, gender, marital status, property ownership, the West variable, vocational category and the annual income range.

**** Significance (p<1%, p< 5% and p<10%) is indicated by the symbols *, **, ******

IV- Results and Interpretation

In order to compensate for loss of income when they retire, workers exhibit a life insurance savings behaviour, which can change during their working lives. Our results show that specific factors shape holding life insurance for retirement schemes. First, we will present the descriptive statistics of the entire population and the three sub-population groups. Secondly, we will interpret the regression results about the relationships that may exist between the different variables studied and the probability of individuals holding retirement savings contracts.

IV.1- Descriptive statistics

Holding retirement savings products in the population studied

Descriptive statistics show that of the sample studied, only 17.3% of individuals hold life insurance or pension savings products. This rate varies according to the sub-sample to which the individual belongs. Indeed, within the sub-sample 'CNRPS', the rate of holding a pension savings scheme is only 10.5%. As for 'RSNA' private sector affiliates, the rate is 20.7% and 21.6% for the 'RINA' self-employed affiliates regime.

Table 5 reports the different Life Insurance holding rates for each sub-population group and the entire population.

Table 5. Life-Insurance holding rates by category of workers

			Y		Total
			,00	1,00	
Sub-population	RSNA	Number	153	40	193
		% in the sub-population	79.3%	20.7%	100.0%
		% in the population	**37.3%**	**9.8%**	**47.1%**
	CNRPS	Number	128	15	143
		% in the sub-population	89.5%	10.5%	100.0%
		% in the population	**31.2%**	**3.7%**	**34.9%**
	RINA	Number	58	16	74
		% in the sub-population	78.4%	21.6%	100.0%
		% in the population	**14.1%**	**3.9%**	**18.0%**
Population		Number	339	71	410
		% in the population	82.7%	17.3%	100.0%

Source: Household Wealth survey 2015.

Rate differences between the private and the public sector can be explained by the fact that assets in the public system are not affected by a significant loss of income upon retirement, unlike assets of the private sector, the income in which declines after retirement. Coverage rate is significantly higher in the public system than in the private system. Insurers tend to offer their products to private sector assets rather than to the public sector, since these contracts are intended to supplement incomes of active individuals and are usually settled upon retirement. This also explains the difference in Life insurance contracts holding rates across affiliate schemes.

Table 6. Life-Insurance holding rates by type of subscription

Life Insurance / Retirement Savings holders			Type of Subscription		Total
			Individual	collective	
Sub- population	RSNA	Number	22	17	39
		%	**56.4%**	**43.6%**	**100.0%**
	CNRPS	Number	12	2	14
		%	**85.7%**	**14.3%**	**100.0%**
	RINA	Number	16	0	16
		%	**100.0%**	**0.0%**	**100.0%**
Population		Number	50	19	69
		%	72.5%	27.5%	100.0%

Source: Household Wealth survey 2015.

In the sample studied, only 27.5% of contract holders are members of group pension savings contracts, purchased by their employers, 89.5% of which work in the private sector.

Funded pension schemes are scarce, although they offer the same range of benefits as those of life insurance products, like disposing of the entire amount at retirement or collecting a reversible or non-reversible life premium. These contracts are mainly taken out by large private sector companies and are essentially meant for senior executives and managers.

The tax benefits offered to these contracts are, in principle, insufficient to convince employers about the usefulness of purchasing such products. Life insurers should deploy more efforts to increase the purchasing rate of these products in the market.

Life insurance policy-holders by socio-demographic and financial variables in the sample studied

Descriptive statistics of the sample studied indicate that the proportion of life insurance policy-holders differs according to pension scheme and also according to the socio-demographic characteristics of each individual in the sample.

The table below summarises the distribution of life insurance policy-holders and funded pension savings scheme-holders according to membership categories.

Table 7. Microeconomic Characteristics of « Life Insurance/ Retirement Savings » Holders

Variable		« Life Insurance/retirement Savings » holding							
		CNRPS		RINA		RSNA		TOTAL	
		Obs*	Rate%	Obs*	Taux%	Rate%	Taux%	Rate%	taux
Gender	M	10	67	13	81.25	27	67.5	50	70.41
	F	5	33	3	18.75	13	32.5	21	29.59
	Total	15	100	16	100	40	100	71	100
Age	20-29	--	--	--	--	7	17.5	7	9.85
	30-39	7	46.67	10	62.5	21	52.5	38	53.52
	40-49	2	13.33	2	12.5	3	7.5	7	9.85
	50-59	5	33.33	4	25	7	17.5	16	22.54
	60-65	1	6.67	--	--	2	5	3	4,22
	Over 65 years	--	--	--	--	--	--	--	---
	Total	15	100	16	100	40	100	71	100
Average age of the holder		42.8		40.87	100	38.1	100	39.71	
Property Ownership	Owner	12	80	11	68.75	31	77.5	54	76.05
	Not Owner	3	20	5	31.25	9	22.5	17	23.95
	Total	15	100	16		40	100	71	100
Geographical Area	Grand Tunis	9	60	9	56	30	75	48	67.6
	North East	3	20	1	6	0	0	4	5.63
	North West	0	0	0	0	0	0	0	0
	Centre East	2	13.33	2	13	6	15	10	14.08
	Centre West	1	6.67	0	0	0	0	1	1.4
	South East	0	0	4	25	2	5	6	8.45
	South West	0	0	0	0	2	5	2	2.81
	Total	15	100	16	100	40	100	71	100
Income Bracket	[10000TND-20000TND[	7	47	1	6.3	11	27.5	19	26.76
	[20000TND-50000TND[	7	47	7	43.7	21	52.5	35	49.3
	>=50000TND	1	7	8	50	8	20	17	23.94
	Total	15	100	16	100	40	100	71	100
Occupational Category	Senior Executive	12	36.36		--	21	63 .63	33	39.44
	Middle-grade Executive	3	20	--	--	19	47.5	22	30.98
	Self-Employed	--		16	100	--	--	16	12.68
	Total	15	100	16	100	40	100	71	100
Matrimonial Status	Married	13	86.67	15	94	29	72.5	57	80.28
	Single	1	6.67	0	0	7	17.5	8	11.27
	Divorced	1	6.67	1	6	4	10	6	8.45
	Total	15	100	16	100	40	100	71	100
Property Ownership	Owner	12	80	11	68.75	31	77.5	54	76.05
	Not owner	3	20	5	31.25	9	22.5	17	23.95
	Total	15	100	16	100	40	100	71	100
Household size	1	4	26.67	2	13	18	45	24	33.8
	2	0	0	2	13	5	12.5	7	9.86
	3	10	66.67	8	50	11	27.5	29	40.85
	>= 4	1	6.67	4	25	6	15	11	15.49
	Total	15	100	16	100	40	100	71	100

Source: Household Wealth survey 2015.

Descriptive results are presented with regression results, in order to identify the determinants of holding funded pension savings schemes that are presented in the next section.

IV.2- Regression Results

Interpretation of regression results will allow for determining the effect of socio-demographic and financial factors on the probability of holding a pension scheme funded by a Tunisian asset, whether positive or negative. In other words, there may be factors that increase or decrease the probability of holding a pension scheme funded by a Tunisian asset. As previously explained in the model, a logistic regression is interpreted as linear modelling of Log-Odds, while the coefficients of the independent variables express the Odds-Ratio.

O-R= $\frac{\pi_{i,j}}{1-\pi_{i,j}}$ with $\pi_{i,j}$ the probability that a worker i of population j

holds a life insurance or a funded pension scheme.
If O-R> 1: increase in Xj leads to increase in the probability that Y is achieved.
If O-R = 1: increase in Xj has no impact on the probability of holding a funded pension scheme.
If O-R <1: increase in Xj leads to decrease in the probability that Y is achieved.

The tables below report a summary of the results of the regressions, namely Odds ratios and Standard deviations. The coefficients will be interpreted, first, for the entire sample, then, for each sub-sample. This will allow us to determine, first, the common Life Insurance holding factors for the entire sample and, secondly, for each sub-sample.

Age and Generation

The analysis of the results of logistic regressions partially contradicts the life-cycle hypothesis, which states that age has a significant effect on holding life insurance for retirement. In the Tunisian context, the hypothesis about the relationship between age and purchasing life insurance contracts does not completely hold. Investment in life insurance products is a long-term endeavour and assets allocated to this type of product seek to supplement people's retirement rather than to achieve high profitability in the short term. Nevertheless, our results indicate that the probability of holding a life insurance or a funded pension product is not significantly influenced by age, irrespective of the public pension plan purchased (p. = 0.652> 0.1).

Table 8. "Life Insurance/Retirement savings" Holding: results of regressions

Variables	Model		Sub-population			
			Population	CNRPS	RINA	RSNA
AGE	1	Age	1.074	3.389**	1.215	1.004
	1	Age2	0.99	0.99	0.99	0.99
	2	Gen 2	0.503	--	0 .00	0.282***
	2	Gen 3	0.812	--	0	0.629
	2	Gen 4	0.272	--	0.062	0.261
Marital Status	2	Single	0.347**	0.448	0	0.415**
	1	Div	2.472	37.368**	1.376	4.558***
	2		2.18	5.622	0.969	3.14
Income Bracket		IncBr	1.768*	1.57	3.099**	1.463***
	2		1.842*	1.796	2.476**	1.578
	1	PubExec	4.084	0.053**	--	--
	2		4.695	0.099	--	--
Vocational Category	1	PrivExec	8.313**	--	--	--
	2		10.020**	--	--	--
	1	SelfE	5.534***	--	--	--
	2		6.274***	--	0.098	--
	1	PropOwn	2.051**	1.689	0.741	3.424*
Property Ownership	2		1.989**	1.514	0.826	3.307**
	1	East	1.232	6.134	--	0.541
Geographical Area	2	West	0.813	0.214	0	1.993
Household Size	1	TMen	0.47	0.42	0.15	0.69***
Gender	1	Gend	1.172	1.34	1.435	1.638
	2		1.379	1.385	1.67	1.67

Logistic regression coefficients "O-R", Significance (p<1%, p< 5% and p<10%) is indicated by the symbols *, **, ***

However, by introducing the generation factor into our model, we found a significant negative relationship between membership to the age group of 30 - 40 year olds and log odds (OR = 0.503 <1), significant at p = 0.095 <0.1. In other words, belonging to this generation decreases the likelihood that the asset will be allocated to life insurance in order to supplement one's retirement. This finding can be explained by the fact that at the beginning of their working lives, individuals lack vocational stability and face many financial constraints, such as debt repayment and large family expenses.

Purchasing life insurance or funded pension contracts often takes place during later years of one's working life, before retirement, which seems to be conducive to saving. Savings accumulated through these products respond to the need to supplement an individual's income after retirement.

In general, the generation variable highlights the effect of the socio-economic factor on savings behaviour. In other countries, uncertainty about the evolution of the financial situation of public pension funds and the anticipation of possible reforms encourages younger generations to save more resources for their retirement

(Brun-schammé and Duée 2009). Nevertheless, in the Tunisian context, we notice that belonging to a younger generation has the opposite effect on the probability of self-employed persons in the private sector holding life insurance. Moreover, results show that belonging to older, close-to-retirement generations does not have a significant impact on the probability of holding pension savings. These retirees are confident about their future retirement pension and believe they will not be affected by pension reforms that might be made to basic pension plans. In addition, when approaching retirement, these retirees consider that long-term savings products are not profitable investments in the short-term, nor are the tax benefits gained through such products.

Table 9. "Life Insurance/Retirement savings" Holding: results of regressions (Standard deviation)

Variable		Equation	Sub-population			
			Population	CNRPS	RINA	RSNA
AGE	Age	1	0.158	0.517	0.424	0.234
	Age2	1	0.002	0.006	0.005	0.003
Generation	Gen 2	2	0.531	--	--	0.674
	Gen 3	2	0.632	--	--	0.887
	Gen 4	2	0.896	--	--	1.224
Marital Status	Sing	2	0.461	0.159	9150.395	0.591
	Div	1	0.578	1.788	1.228	0.845
		2	0.58	1.408	1.25	0.849
Income Bracket	IncBr	1	0.19	0.385	0.457	0.274
		2		0.387	0.463	
Vocational Category	PrivExec	1	1.056	--	--	7688.64
		2	1.095	--	--	7618.815
	PubExec	1	1.084	1.739	--	--
		2	1.065	1.633	--	
	Self-E	1	1.083	--	40193.01	--
		2	1.089	--	48990.321	--
Property Ownership	PropOw	1	0.335	0.781	0.724	0.473
		2	0.337	0.79	0.719	0.472
Geographical Area	East	1	0.685	1.455	23184.905	1.022
	West	2	0.685	1.337	23205.422	1.027
Household Size	HSize	1	0.134	0.313	0.244	0.222
Gender	Gend	1	0.319	0.683	0.813	0.445
		2	0.324	0.667	0.807	0.45

Income

An analysis of regression results confirms the hypothesis that income has a significant effect on the probability of an individual holding life insurance for retirement. This observation confirms the life-cycle assumption that this factor has a prominent effect on savings-for-retirement behaviour. The logistic regression results indicate that holding a life insurance contract by a Tunisian asset positively relates to one's annual income. Indeed, results suggest that the probability of holding insurance products increases when an individual's income is in the upper range OR = 1.768> 1 and p = 0.003 <0.01 and individuals with low income are the least likely to purchase such contracts. These results confirm those found in previous studies.

This significantly positive relationship between income and the probability of purchasing a life insurance / funded pension contract holds for in the entire sample as well as for the RSNA and RINA sub-samples. High annual incomes increase the probability of life insurance ownership only among private sector employees with O-R = 1.463> 1 and p = 0.065 <0.1. Contracts may be purchased individually or collectively and this allows an employee to enjoy tax benefits. Similarly, results show that this positive relationship also holds for the RINA sub-sample with O-R = 3.099> 1 and p = 0.013 <0.05.

These results can be explained by the fact that private sector assets (employees and self-employed) with higher annual incomes are more likely to experience a large loss of income upon retirement, contrary to those holding public sector assets. The latter will receive a public pension quite similar to their working income, since the reference salary base taken into account while calculating the retirement pension is much more advantageous for the public sector than for the private sector. The probability of purchasing a funded pension plan increases with the growth of one's annual income. These highly-remunerated assets have significant savings capacity and express one's willingness to offset the sharp decline in income anticipated after their retirement, while benefiting from tax deductions (Caroll 1998, Arrondel 1996).

Marital Status and Household Size

Looking at the regression results, we found that marital status has a significant impact on holding life and pension products, which confirms the assumption made initially. In our study, a 'single' person is less likely to hold a pension savings product than a married person. Being single reduces the probability of holding life insurance / pension savings in the group with O-R = 0.347 <1 and p = 0.022 <0.05. This can be explained by the fact that a single person seems to be less interested in products that provide coverage for spouse and children than in other savings products that may be less expensive without offering these features.

However, results show that household structure and size do not have a decisive impact on holding pension savings for the entire sample. This confirms the results of some previous studies (Mahieu (2001), Mekkaoui de Freitas (2002)).

Although tax law regulating life insurance products allows for an increase in tax deductions with an increase in the number of dependent children, our results indicate absence of a positive relationship between household size and the probability of holding life insurance or retirement savings.

The tax advantage does not seem to be an incentive for purchasing such savings products. This may be explained, on the one hand, by the fact that most households are poorly informed about these benefits and tax deductions and, on the other hand, by the limited savings capacity of large families (Brun -Schammé & Duée 2008).

Nevertheless, we notice that belonging to the 'couple with 3 or more children' category under the RINA sub-sample increases the probability of holding a life insurance product. This can be explained by a concern to protect the future of one's family and to compensate for the risk of not having provided for oneself after retirement. The long-term saving decision expresses a desire to maintain a future income allowing one to cope with rising expenses, given their large household size (Lewis 1989, Fischer 1973).

Vocational Category

Vocational category in Tunisia reflects other socio-economic parameters, like income level, prospects of salary raise, and educational level.

According to our results, one's vocation has a significant impact on holding a life insurance plan only for private RSNA and RINA plans. In fact, 'senior managers' and 'senior executives' are more likely to hold supplementary pension products. The relationship is positive between vocational category and the probability of holding life insurance for RSNA-affiliated employees with $OR = 8.313 > 1$ and $p = 0.045 < 0.05$. This can be explained by the fact that members of this vocational category in Tunisia are generally known for high wages and salaries and prospects for favourable future income trends (Marti 2011).

Moreover, it should be noted that for the public sector, vocational category does not have a significant impact on the probability of holding pension products. Results indicate that the 'vocational category' variable does not significantly affect the probability of holding a life insurance policy. This means that one's vocation does not determine the purchase of a life insurance policy for the public sector regime. This result can be explained by the fact that affiliates to the Tunisian public sector regime do not risk significant loss of income when they retire. Their retirement pension paid by the Social Security Fund is not capped, as in the case of private sector affiliates, and the standard of living of public service retirees is similar to that of alternative assets.

Such pension coverage reduces the probability of holding funded pension supplements for this asset class.

Nevertheless, results show that the 'self-employed' category is more likely to contract pension savings with O-R = 6.247> 1 and p = 0.093 <0.1. This can be explained by the fact that members of this group wish to protect their future pensions, since the basic retirement amount of the self-employed is significantly lower than their actual activity income (Brun-Schammé 2009).

Property Ownership

Property ownership is a common holding factor to all Tunisian asset classes. Indeed, regression results indicate that ownership of a main property has a positive effect on the probability of holding a life insurance or a funded pension supplement with OR = 2.164> 1 and p = 0.021 <0.05. This significant positive relationship holds both for the entire sample and or the three sub-samples, i.e., RSNA, RINA and CNRPS. Moreover, it is owners of main properties who tend to purchase this type of contracts, which confirms the results of Arrondel & Savignac (2011), and Brun-Schammé & Duée (2008).

Results indicate that it is necessary to differentiate between owners of a main property and tenants who pay rent and who generally have lower savings capacity than homeowners. This makes them less interested in life insurance and pension products and also reflects the savings priorities of a Tunisian asset and shows that Tunisians allocate their resources to acquiring a house first, and that accumulating savings for retirement in the form of life insurance comes next. However, those who do not have access to owning a property are unlikely to purchase insurance or pension products.

Geographical Area

Several authors stipulate that the geographical area determines holding life and common insurance for both the private and public sectors. This assumption is not confirmed by our regression results, suggesting, all things being equal, that the geographical area of the household is not a significant factor determining holding life insurance for retirement.

Indeed, living in the Greater Tunis, Central East (Sahel) and South East (mainly Sfax) regions does not significantly impact the probability of holding a pension product. This is inconsistent with other studies in other countries, which consider the geographical area as a factor determining pension savings behaviour (Dauner 2002). In our study, geographical area does not have a significant impact, despite the difference in the demographic and economic characteristics of the sample in each region (income level, education, age, etc.). This can be explained by the proximity of financial institutions that offer such contracts in all regions of the country (Marti 2011, Dauner 2002). Indeed, life insurance agencies are located in all cities of Tunisia and offer the same life insurance and savings products to all citizens.

Gender

Previous studies state that asset type has a significant impact on the probability of holding life insurance / funded pension products, and that men are more likely to purchase such contracts than women. Nevertheless, according to the regression results, this stipulation is not confirmed. Gender does not have a significant effect on life insurance savings decisions.

All of the results above are summarised in the following table which indicates the negative or positive impact of each variable on retirement savings and holding a life insurance product in Tunisia.

Table 10. Effects of the variables on the probability of Life Isurance contracts holding

Variable		CNRPS	RINA	RSNA
AGE	Age	©	©	©
Generation	1981- 1990	-	-	-
	1971 - 1980	-	-	-
	1961 - 1970	©	©	©
	1951 - 1960	©	©	©
	1931 - 1950	©	©	©
Marital Status	Sing	-	-	-
	Div	©	©	©
	Wid	©	©	©
Income Bracket	IncBr	+	+	+
Vocational Category	PrivExec			+
	PubExec	–		
	SelfE		+	
Property Ownership	PropOw	+	+	+
Geographical Area	East	©	©	©
	West	©	©	©
Household Size	HSize	©	©	©
Gender	Gend	©	©	©

Source: Authors' calculations - 2015.

(©) The variable has no significant effect on the probability of holding a life insurance contract

(+) The variable has a positive significant effect on the probability of holding a life insurance contract

(-) The variable has a negative significant effect on the probability of holding a life insurance contract.

IV-Conclusion

The financial difficulties faced by Tunisian pension systems in addition to the harsh demographic and economic developments have been considered in this study. Our aim is to highlight the determinants of holding a funded pension or life insurance plan. We referred to many studies carried out in different countries to examine the behaviour of different population groups in regard to holding complementary retirement savings products. Our contribution is that we focus on a Tunisian sample and distinguish individuals according to the pension scheme they are affiliated to.

This allowed us to test our hypotheses, taking into account the specificity of each category.

The results of the logistic regressions pointed to the determinants of holding a pension plan that are common to all categories, notably life cycle variables and ownership property. The latter seem to have a significant impact on group pension savings behaviour.

Nevertheless, we found other factors specific to each asset class, such as vocational category, which is a significant factor for private assets but has no effect on public assets. This result can be explained by differences in the parameters considered and the rules for calculating pension rights for the private and public sectors.

As for marital status, results prove that this factor does influence the decision to hold a supplementary pension savings plan and that a 'single' Tunisian is less interested in these products than a 'married' one. Married people hold assets to try and protect the future of their spouses and children. Moreover, results indicate that socio-demographic and financial factors exert the same effect on the entire sample and on the sub-samples. Moreover, the level of impact of each factor varies from one category to another. Similarly, some factors can be decisive in a specific category of the sample but not in the entire sample.

Results highlight the socio-demographic and financial characteristics that influence a Tunisian in regard to holding a life insurance product for retirement. This might give a general idea about the potential of developing life insurance and funded pension products for Tunisia. Nevertheless, the present study examines data observed at a given point in time. It may be interesting to follow a developmental approach to savings behaviour over several years in the future, take into account the premiums paid and assess the amounts invested in such contracts.

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Book Review

Bridging the Prosperity Gap in the EU
The Social Challenge Ahead

edited by Ulf Bernitz, Moa Mårtensson, Lars Oxelhei and Thomas Persson

Edward Elgar Publishing, Cheltenham, UK, Northhampton, MA, USA, 2018, p.p. 264
ISBN: 978 1 78643 666 5 (cased)
ISBN: 978 1 78643 667 2 (eBook)

reviewed by Michele Raitano[1]

The book "Bridging the Prosperity Gap in the EU", edited by Ulf Bernitz, Moa Martensson, Lars Oxelheim and Thomas Persson, focuses on the crucial issue of how to reduce economic wellbeing discrepancies among EU countries. The authors note that, as a dramatic aftermath of the Great Recession started in 2008, the prosperity gap between rich and poor countries – and between rich and poor individuals within each country – has increased so much as to cast doubts on whether the concepts of economic integration and solidarity can be combined. Indeed, the relationship between economic policies and fiscal rules – carried out at the EU level –and social policies – carried out at the national level with soft, non-binding EU coordination – is highly asymmetrical, thus strongly limiting the scope for redistribution between states, on the one hand, and among individuals, on the other. The recent commitments towards a European Pillar of Social Rights and the anti-poverty targets stressed in the Europe 2020 strategy do not seem enough to effectively reverse the current trend of an increasing social deficit in the EU and help bridge the prosperity gap within the EU. In the wake of such a worrisome picture, the contributors to this interesting book inquire the possibility of addressing this huge social challenge, reducing differences across countries, and avoiding a possible EU collapse from several perspectives and disciplines.

To this end, the book includes 9 chapters – plus a detailed introductory chapter which points out the main research questions – in which the issue of the prosperity gap between countries is assessed from different perspectives; besides, some, mostly EU, policy suggestions are proposed so as to restore the balance of the roles played by market integration and social protection.

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The book is well-structured and full of original and, usually, under-emphasized insights, regarding, for instance, the capacity of social dialogue to strengthen EU legitimacy (Chapter 6), the role trust plays as a driver of gaps in well-being (Chapter 9), and the effectiveness of effective institutions to improve social cohesion through positive effects on social capital and trust (Chapter 10). Some chapters focus more explicitly on political challenges brought about by recent events, such as those related to the emergence of populist right-wing parties across almost all EU countries (Chapter 3) and to the effects of austerity programmes on voters' preferences (Chapter 5). Other chapters take a right-based approach to assess recent trajectories of social security systems (Chapter 8) – also related to workers' mobility across countries (Chapter 4) – labour market institutions (Chapter 7), and, more generally, social rights (Chapter 2).

The general conclusion of the book is that the EU may re-gain legitimacy only if it proves capable of reducing the prosperity gap, which has been deepening since the emergence of the Great Recession, by designing an actual new path for the EU to make the dream of a social Europe come true.

The research agenda pursued in the book is, therefore, clear; therefore, those who are interested in the building of a real social Europe fostering social cohesion cannot possibly disagree with nearly any of the statements made by the contributors.

However, despite the range of issues raised by the editors, certain crucial issues have been neglected, which, in my opinion, should also be considered in order to paint an exhaustive picture of the current social Europe stance, on the one hand, and of the main obstacles that need to be removed to effectively increase social cohesion, on the other.

In particular, the book lacks the deepening of two phenomena – one related to micro distributive characteristics and the other to macro structural features – the drivers and implications of which should be carefully investigated to foster a parallel increase in economic growth and social protection/cohesion.

First, the many contributors of the book focus almost exclusively on the prosperity gap between countries, without taking into consideration an equally important increasing prosperity gap, i.e., that associated with the increasing socio-economic inequality within almost every EU country. And the increase in "within-country" inequality, starting from relatively high levels, even before the crisis, in many countries might represent a driver of the diminishing confidence that many EU citizens feel for both national and EU political institutions.

From this perspective, the responsibilities of austerity programmes are clear, and their impact may be undervalued if average gaps between countries are considered alone. Furthermore, the role of the EU as a constraint to social cohesion should not only be assessed in relation to the limits imposed by austerity programmes to redistributive measures or through cuts in social spending. In order to truly assess how EU actions might help foster social cohesion, the question to be asked is how EU policies influence market equilibria. To foster ~~a~~ highly inclusive growth, indeed, redistribution does not suffice, and one may ask what the features of the equilibria engendered in the markets are – firstly in the labour market – before possible redistribution may be implemented. From this perspective, frequent suggestions, such as increasing labour market flexibility and decentralising bargaining, or ineffective actions, such as contrasting tax competition among countries, might have contributed towards both exacerbating both functional distribution between wages and profits and reducing labour share and market income inequality, as reflected in the upward trend of Gini Indices of market incomes in almost all major EU countries since the 1980s. To this end, what should be highlighted more are the responsibilities of EU policies in the processes that have risked an increase in market inequalities and a decrease in the redistributive role played by national governments.

Additionally, certain constraints to the social dimension arising from belonging to the European Monetary Union (EMU) should be more clearly emphasised. Indeed, even if the authors focus on the entire EU, instead of on countries belonging to the EMU, what should be stressed are the consequences of the asymmetrical and incomplete EMU implemented as well as the need to design more effective policy proposals.

As correctly pointed out by the authors, the EU budget is too limited to allow the EU to improve solidarity between and within countries and innovative tools – e.g., an EU-based unemployment benefit scheme should be soon implemented, which jointly transfers resources to countries that need it more and, within countries, to more disadvantaged individuals. However, apart from the possibility of introducing these redistributive tools, which clearly clashes with the current political sentiment in the EU, some basic lessons from the optimal currency area theory originally proposed by the Nobel Prize winner, Mundell, in the early 1960s should be kept in mind to highlight the steps necessary for reducing the prosperity gap in the EU.

Indeed, it is widely accepted that, if structural economic conditions of EU countries are not similar – which, on the contrary, have further diverged in recent years – there are only two options to reduce asymmetries between countries. The first one is a sort of 'low equilibrium' option and, unfortunately, this is what has, so far, been followed by the EU. In other words, increasing workers' mobility and labour market flexibilization to increase the cost-competitiveness of disadvantaged countries. However, this option is very painful for many citizens in these countries: even if it might bridge the average prosperity gap between countries – as measured, e.g., by the GDP growth rate – it is also likely to increase prosperity gaps within countries, worsening social cohesion. The second option is the 'high equilibrium' one, that should have, instead, been pursued by anyone who cares about the future of the EU; in other words, strengthening the EU budget and complementing the Monetary Union with an effective Fiscal Union, where a significant portion of national budgets is shared among countries. This 'high equilibrium' option would, indeed, help restore legitimacy in EU building, weakening the appeal of anti-Europe parties and fostering the idea of EU solidarity and social cohesion. However, paraphrasing the words of the editors of the book, presently, the dream of this type of EU is unlikely to come true in the next few years.

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The ***South Eastern Europe Journal of Economics*** gratefully acknowledges the continued support of the distinguished members of the Editorial Board and the following roster of economists who have contributed their services as referees during the period Fall 2016 - Spring 2019.

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