

Marsellou, Emilia G.

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# Testing the Bhaduri–Marglin model for the demand regime of Greece

Emilia G. Marsellou

Centre of Planning and Economic Research (KEPE), Athens, Greece  
emarsellou@kepe.gr

*This paper investigates the demand regime of the Greek economy for the period 1960–2017, using a Bhaduri/Marglin (1990) type macroeconomic demand-led model and applying the empirical methodology developed in Stockhammer et al. (2009). The analysis finds that Greece is another European economy with a wage-led demand regime, both domestically and as an open economy. Moreover, the analysis shows that the results are robust across different sample sizes, for the period 1960–2017, and the two sub-periods 1960–1989 and 1990–2017.*

**Keywords:** functional income distribution, demand regimes, consumption, investment, foreign trade, Greece, wage moderation

**JEL codes:** E12, E20, E21, E25, F40

## 1 INTRODUCTION

The implementation of the Economic Adjustment Programs in Greece during the 2010–2017 period, as a result of the bailout agreement between Greece and the Troika (International Monetary Fund, European Central Bank, and European Commission), contained austerity measures aiming at fiscal consolidation, as well as structural reforms in labour and product markets aiming at the stimulation of private investments and exports. However, it took six years of severe austerity and heavy bailout programs for the Greek economy to stabilize and record, for the first time since the start of the programs, a positive growth rate of 1.1 per cent in 2017, while the unemployment rate was as high as 21.5 per cent. Leaving the fiscal issues aside, these policies, for the most part, rely on the theoretical belief which attributes the persistent high levels of unemployment in Europe and the anaemic growth rates to the labour market rigidities. Policies designed to remove these rigidities are contained in the wage moderation strategy (European Commission 2011, 2012). However, during times of economic turmoil, such as the one Greece was exhibiting, such policies take the more drastic form of internal devaluation.

Wage moderation policies are those that lead to a decline in the share of wages in total product, without direct intervention at the level or course of profits (Lavoie/Stockhammer 2012). In other words, the decline in the wage share occurs through institutional interventions in the operation of the labour market and the wage-setting mechanisms. Such interventions aim at the weakening of collective bargaining, the limitation of trade unions activities, the degradation of labour legislation, and the reduction/abolition of minimum wages. On the other hand, a pro-labour economic strategy includes policies that promote

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the role of trade unions, strengthen legislation that protects labour, and support the welfare state.<sup>1</sup>

A country's production structure determines, for the most part, the impact of a (re)-distribution of income on economic growth.<sup>2</sup> The literature suggests a typology of aggregate demand regimes based on whether a change in the functional income distribution stimulates or restrains the demand growth; if an increase in the wage share has a positive effect on aggregate demand, it is considered as wage-led; otherwise as profit-led. It is quite straightforward, then, that in a wage-led economy, policies which compress the wage share are recessionary.

The aim of this paper is to determine the Greek demand regime in an attempt to contribute to the dialogue on why the Economic Adjustment Programs had such a devastating and prolonged impact on the Greek economy from a demand-side point of view. The investigation period is set to 1960–2017, which is split into two sub-samples: the first for the period 1960–1989 and the second for the period 1990–2017. The split is chosen to be in the middle of the full sample in order to guarantee that both sub-samples have the same and the appropriate number of observations for the estimations. The year 1989 was confirmed exogenously as a breakpoint according to the Chow test for all the demand component equations. In addition, the 1990s was a decade of several transformations for the Greek economy as the country was preparing itself to meet the convergence criteria of the EU Treaty (1992)<sup>3</sup> in order to adopt the euro.

The paper is organized as follows: Section 2 briefly introduces the theoretical model, Section 3 reviews the literature, and Section 4 presents the econometric analysis and regression results. Section 5 discusses policy implications.

## 2 THEORETICAL BACKGROUND

A private sector open economy model is estimated based on the theoretical framework of Bhaduri/Marglin (1990) and the empirical work of Stockhammer et al. (2009).<sup>4</sup> Given this model has become a standard point of reference for the specific empirical literature, its exposition will remain brief. Aggregate demand ( $AD$ ) is the sum of consumption ( $C$ ), investment ( $I$ ) and net exports ( $NX$ ). All demand components depend on income ( $Y$ ), the wage share ( $\Omega$ ) and some control variables ( $z$ ).

$$AD = C(Y, \Omega, z) + I(Y, \Omega, z) + NX(Y, \Omega, Y^f, z) \quad (1)$$

1. The main goal of the pro-labour policies is to reduce income inequality and poverty among the citizens of a country. The adverse economic effects of rising income inequality on economic growth are explored in the mainstream literature, see for instance: Alesina/Rodrik (1993; 1994), Persson/Tabellini (1994) and Alesina/Perotti (1996).

2. Apart from the production structure, Lavoie/Stockhammer (2012) report a number of other factors which may contribute to whether an economy is wage-led or profit-led: the existing functional distribution of income, the propensity to consume class differential, business responsiveness to sales or profit margins, the sensitivity of exports to changes in production costs, exchange rates and external demand, and the relative size of the components of aggregate demand (consumption, investment, net exports).

3. <https://eur-lex.europa.eu/legal-content/EL/TXT/PDF/?uri=CELEX:11992M/TXT&from=EN>.

4. In a similar framework, Obst et al. (2020) developed a multi-country model that incorporates the government sector.

Consumption is expected to increase with the wage share. Investment depends on income, the profit share, and the interest rate. It is expected that investment rises with income and the profit share and falls with the interest rate. Net exports are negatively related to domestic demand ( $Y$ ) (through rising imports) and positively to external demand ( $Y^f$ ) (through rising exports). Income distribution enters the equation of net exports through the real unit labour cost (RULC) which is an alternative to the wage share variable. RULC is expected to have a negative impact on net exports, since a rise in the wage share increases RULC, generating a loss in competitiveness driving net exports down (Blecker 1999).<sup>5</sup>

Financial variables except for the interest rate are omitted. A potential weakness of the single-equation approach is that it treats income distribution as exogenous. This is because it is possible to have simultaneous causality between distribution and demand (Blecker 2016).<sup>6</sup>

To analyse the impact of income distribution on economic activity, it is assumed that, in the product market, aggregate demand equals aggregate supply: ( $Y^* = AD$ ). Differentiating  $Y$  with respect to  $\Omega$  we have:

$$\frac{dY^*/Y}{d\Omega} = \frac{b_2}{1 - b_1} \quad (2)$$

where  $b_1 = \left(\frac{\partial C}{\partial Y} + \frac{\partial I}{\partial Y} + \frac{\partial NX}{\partial Y}\right)$  and  $b_2 = \left(\frac{\partial C/Y}{\partial \Omega} + \frac{\partial I/Y}{\partial \Omega} + \frac{\partial NX/Y}{\partial \Omega}\right)$

The left-hand side of equation (2) gives the percentage change in GDP due to a change in the wage share since  $dY^*$  is divided by  $Y$ . The right-hand side of equation (2) consists of the standard multiplier  $\left(\frac{1}{1 - b_1}\right)$ , which is assumed positive for equilibrium in the products market, and parameter  $b_2$ , which is the sum of the partial derivatives of the demand components with respect to the wage share. This sum is called private excess demand and represents the change in aggregate demand which is produced by a change in income distribution given a certain level of income (Stockhammer et al. 2009). A positive  $b_2$  implies that a higher wage share affects positively aggregate demand and hence the Greek demand regime is wage-led; otherwise, it is profit-led.

The analysis is divided into two stages: first, the domestic sector (consumption and investment) is estimated, and second, the external sector (net exports) is included. This is because empirical results often change significantly when the external sector is taken into account (Blecker 2016). Usually, as consumption is more sensitive than investment to changes in the wage share, the domestic economy is wage-led (Onaran/Galanis (2012) for instance, provide evidence that domestic demand is wage-led for the G20 countries). However, if this regime is weak, when the impact on net exports is added, the overall demand regime might shift to profit-led.

5. Note that in macro-econometric models, ULC is usually used in order to capture the effect of prices on exports and imports (Stockhammer et al. 2009). However, to the extent that RULC is used, then prices should also enter the equation as a proxy for competitiveness.

6. Blecker et al. (2022) correct the simultaneity bias problem by estimating a similar econometric model consisting of seven equations for the US economy using the system GMM estimator. The authors find that by correcting the endogeneity problem the wage-led result becomes even stronger compared with the result of the OLS estimates.

### 3 LITERATURE REVIEW

Since the publication of the Bhaduri–Marglin model in 1990, an extensive empirical literature examines whether the demand regime of several countries is wage-led or profit-led. As this literature evolves, it differentiates with respect to the econometric method – single-equation approach or (Structural) Vector Autoregressive (S/VAR) models, the behavioural equations used, if the analysis incorporates one country/group of countries or a panel of countries, the size of the country, the stage of development, and the time span.

The first and more numerous group of the literature, in which this work belongs, applies the single-equation approach where consumption, investment, and the external sector functions are estimated separately, and at a later stage of the analysis their marginal effects with respect to the wage share are added. Bowles/Boyer (1995) are the first to empirically examine the Bhaduri–Marglin model. They estimated five OECD economies for the period 1961–1987 and find that, as closed economies, France, Germany, Japan, the U.K. and the U.S. are wage-led, while as open economies, France, Germany and Japan turn to profit-led. Naastepad (2006) finds that the growth rate of aggregate demand in the Dutch economy is hardly sensitive to changes in the growth of real wages and that investment and the growth of exports are hardly affected by profits. Naastepad/Storm (2007) investigate eight OECD countries (France, Germany, Italy, Netherlands, Spain, the U.K., Japan and the U.S.) during 1960–2000. All of them as closed economies are found to be wage-led, but when the external sector is introduced, Japan and the U.S. shift to profit-led. Stockhammer/Ederer (2008) and Ederer/Stockhammer (2007) investigate Austria and France, respectively, and find that as closed economies they are wage-led, while as open economies they turn into profit-led. Hein/Vogel (2008) use data on six OECD countries for the period 1960–2005. France, Germany, the U.K. and the U.S. were found wage-led, while Austria and the Netherlands were profit-led. Stockhammer et al. (2009) find the euro-area demand regime for the period 1962–2005 to be wage-led. Onaran et al. (2011) examine the impact of financialization and income distribution on U.S. aggregate demand covering the period 1962q<sub>2</sub>–2007q<sub>4</sub> and find a weak wage-led demand regime. Stockhammer/Stehrer (2011) use quarterly data to estimate for the domestic demand of 12 OECD countries during the 1970–2007 period. The domestic demand of the majority of the countries was found wage-led mainly because of the positive impact of the wage share on consumption rather than because its impact on investment was weak. Onaran/Galanis (2014) investigate the G20 countries covering the period 1960–2007. Their results suggest that the euro area as a whole is wage-led, including the three major European economies, Germany, France and Italy. The U.K., U.S., Japan, Turkey and Korea were also found to be wage-led. China is found to be profit-led due to the strong impact on net exports, South Africa is profit-led as well due to the low propensity to consume class differential; Mexico was also found profit-led due to the strong effects of profits on both investment and net exports; and Argentina and India due to the weak effect on consumption. Other country-specific papers using the single-equation approach include Jetin et al. (2012), Molero-Simarro (2011), Wang (2009) and Molero-Simarro (2015), who all find the Chinese demand regime to be profit-led. Also, Jetin/Kurt (2016) find a profit-led demand regime for the period 1970–2011 in Thailand.

As far as Greece is concerned, Marsellou (2013) only for Greece for the period 1960–2009 and Onaran/Obst (2016) for EU-15 for the period 1960–2013 follow the modelling and estimating strategy of Stockhammer et al. (2009), both finding a wage-led demand regime. More recently, Obst et al. (2020) developed a multi-country model to estimate the effects of income distribution and fiscal policy on the components of

aggregate demand and the budget balance in EU-15 countries. All individual countries and the EU-15 as a whole are found to be wage-led. This paper attempts to add to the aforementioned literature, and especially the one regarding the Greek economy, in three ways: (a) extending the period of the analysis to capture the lost decade of economic crisis; (b) exploring the robustness of the estimation results through the split of the sample into two sub-periods; and (c) utilizing, except for the first differences estimator, other modelling techniques when applicable.

The other literature group utilizes the vector autoregressive model approach (S/VAR models). Gordon (1995a) used an autoregressive model to assess the effect of income distribution on consumption and investment in the U.S, and contrary to Bowles/Boyer (1995), concludes that the growth of the U.S. economy is driven by profits. Gordon (1995b) extends the model for an open economy and finds that the U.S. economy is profit-led. Stockhammer/Onaran (2004) found no statistically significant effect of the profit share on capacity utilization and accumulation for the economies of the U.K., U.S. and France. Onaran/Stockhammer (2005) use an SVAR model to investigate the growth regime of Turkey (with estimation period 1965–1997) and South Korea (with estimation period 1970–2000) and find that although Korea and Turkey represent two different export-oriented growth experiences, their growth regime is stagnationist and their accumulation regime is not profit-led. Barbosa-Filho/Taylor (2006) test a VAR model inspired by Goodwin's (1967) model with two variables, effective demand and the wage share, for the U.S. economy and estimation period 1948–2002; they find a profit-led demand regime. Carvalho/Rezai (2015) estimate a two-variable threshold VAR model for the U.S. economy for the 1967–2010 period and find that the demand regime is profit-led. Nishi (2012) finds a profit-led accumulation regime for Japan during 1992–2010 and Burle/Carvalho (2021) find a profit-led demand regime for Brazil for the period 1997–2014.

The third literature group uses panel-data estimation techniques. Kiefer/Rada (2014) use a Goodwin-type model to examine the interaction between income distribution and economic activity for 13 OECD countries for the period 1971–2012 and find that economic activity is weakly profit-led. Hartwig (2014) estimates with the single behavioural equation approach a panel of 31 OECD countries for the period 1970–2011 and finds the demand regime weakly wage-led. Stockhammer/Wildauer (2016) for a panel of 18 OECD countries during the 1980–2013 period, find that the demand regime is wage-led.

## 4 ECONOMETRIC ANALYSIS

### 4.1 Methodological issues

The econometric analysis applies the single-equation approach, where consumption, investment and net exports equations are estimated separately. The analysis begins with the performance of conventional unit root tests – the Augmented Dickey–Fuller (ADF) and the Phillips–Perron (PP) unit root tests. Cointegrating relationships were established for most of the equations via the Johansen cointegration test<sup>7</sup> when all variables are  $I(1)$ ,

7. The lag structure for the Johansen cointegration test is based on the Akaike (AIC), Schwartz (SC) and Hannan–Quinn (HQ) information criteria. Then the Pantula Principle is applied to choose the optimal specification of the cointegration equation (Asteriou/Hall 2007).

and via the cointegration test suggested by Pesaran/Shin (1999) and Pesaran et al. (2001)<sup>8</sup> when there is a mix of I(0) and I(1) variables.

According to the results of the tests and other model-specific reasons discussed in more detail below, the consumption and export–imports equations are estimated using a first differences formulation and the investment equation using an Autoregressive Distributed Lag (ARDL) model. Finally, the elasticities of the demand components with respect to income distribution are transformed into partial marginal effects which are added to obtain the total impact on aggregate demand.

All variables, except for the interest rate, are taken from the annual macroeconomic database of the European Commission (AMECO). The long-run interest rate comes from the Central Bank of Greece (BoG).<sup>9</sup> All variables are in constant prices (2015 = 100) utilizing the respective deflator, and in logarithmic form (except for the real long-run interest rate). The definitions of the variables, the results of the unit root tests, and the cointegration tests are presented in the Appendix. The econometric analysis was performed using EVIEWS version 11.0.

## 4.2 Consumption

A typical Keynesian consumption function is estimated, where income is split into wages and profits, in order to estimate the impact of exogenous changes in income distribution. The consumption function is in a double-log form and is given by:

$$\log C_t = a + e_{C,W} \log W_t + e_{C,R} \log R_t + u_t \quad (3)$$

where  $\log C$ ,  $\log W$  and  $\log R$  are the logarithms<sup>10</sup> of consumption, wages and profits, respectively; and  $a$ ,  $e_{C,W}$ ,  $e_{C,R}$  and  $u_t$  are the constant term, the consumption elasticities with respect to wages and profits, and the error term, respectively. As self-employment in Greece is a significant portion of total employment, income distribution variables are adjusted attributing all self-employment income to labour income.

The unit root tests suggest that  $\log C$  and  $\log W$  are I(1) and  $\log R$  I(0), and hence the ARDL F-Bounds Test for cointegration suggested in Pesaran et al. (2001) is applied.

8. The ARDL bound testing method uses the F statistic to test the joint statistical significance of the coefficients of lagged variables under the null hypothesis of ‘no cointegration relation’. For the case of models with I(0) and I(1) variables, Pesaran et al. (2001) provide two asymptotic critical values where the lowest value assumes that all variables are I(0) and the highest value that all are I(1). The test leads to three possible results: (a) if the statistic F is lower than the lower critical value then we conclude that there is no cointegration relation; (b) if it is between the limits of the two critical values then the result is indefinite; and (c) if it exceeds the higher critical value then the null hypothesis is rejected and consequently there is a long-run equilibrium relationship. If the existence of a cointegration relationship is found, then the long-term and short-term relationships of the variables of the model can be estimated using the ARDL method (Asteriou et al. 2016).

9. Because the interest rates provided by the AMECO database are available since 1973, in order not to reduce the sample period, it is preferred to use the data provided by the Bank of Greece. Due to two breaks in the nominal interest rate time series, we had to link three time-series concerning non-financial enterprises, for loans of more than 1 million euro, and duration of more than five years. We then deflated the series using the price deflator GDP at market prices provided by AMECO.

10. The consumption function variables exhibit exponential growth and hence their variance increases over time. The logarithmic transformation is a linear transformation that reduces the variance of the variables without affecting their relationship.

According to the latter, there is evidence of cointegrating equations at the 10 per cent level of significance (Table A4) and usually this implies that we can proceed with an ARDL model. However, due to high correlation between the income distribution variables causing a serious problem of multicollinearity, the first difference formulation is used.

Table 1 reports the estimation results. For all three periods the model behaves well (adj.  $R^2$  ranges between 0.64 and 0.90), and there is no evidence of violation of the classical linear regression assumptions, namely heteroskedasticity (Breusch–Pagan–Godfrey and White test), serial correlation (Breusch–Godfrey serial correlation LM test), misspecification of functional form (RESET test), multicollinearity (Variance Inflation Factor (VIF) test) and non-normality of residuals (Jarque–Bera test).

As can be seen, the coefficients appear to have the predicted signs and they are statistically significant at 1 per cent and 5 per cent levels. The short-run consumption elasticities with respect to wages and profits range between 0.55 and 0.59, and between 0.06 and 0.17, respectively.

To calculate the direct effects of a change in the wage share on the consumption share in GDP, the elasticities are transformed according to the following relations

Table 1 Regression results for the consumption equation

	1960–2017		1960–1989		1989–2017	
	D(logC)		D(logC)		D(logC)	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
$c$	0.016	5.278***	0.025	4.975***	0.009	3.299***
D(logW)	0.564	9.594***	0.551	6.073***	0.587	10.059***
D(logR)	0.104	4.052***	0.062	2.081**	0.173	3.102***
D1985			–0.045	–2.331**		
D1991					0.029	1.911*
D2010_11	–0.046	–3.009***			–0.030	–2.552**
$R^2$	0.795		0.674		0.911	
Adj. $R^2$	0.784		0.635		0.896	
Jarque–Bera (p)	0.926		0.142		0.203	
BG-LM (2)	0.904		0.201		0.793	
BPG Het test/White test	0.278	0.171	0.747	0.549	0.062	0.091
RESET test (p)	0.195		0.672		0.057	
SSR	0.018		0.009		0.004	
Max VIF	1.3		1.0		1.1	
Chow test (1989)	0.002					
Obs.	57		29		28	

Notes:  $p$ -values MacKinnon–Haug–Michelis (1999). Statistical significance at: \*\*\* 1 per cent; \*\* 5 per cent; \* 10 per cent. All dummy variables were selected to correct the sample after inspection of outliers in the residuals. D2010\_11 captures the first two years of the first memorandum and 1985 was a national election year. It is worth noting that the second sub-period, which includes the economic crisis and the long-term recession that followed, is characterized by strong variability in all the time series used. D1991 corrects for an outlier observed in the residuals of the 1989–2017 regression most probably caused by a break in the (level and differenced) profits time series. Due to evidence of mild heteroskedasticity in the residuals of the 1989–2017 period, the HAC standard errors & covariance (Bartlett kernel, Newey–West) estimator is applied.



Table 2 Calculation of the partial direct effects

	$e_{C,W}$	$e_{C,R}$	av. $C$	av. $W$	av. $R$	$\frac{\partial C}{\partial W}$	$\frac{\partial C}{\partial R}$	$\frac{\partial C}{Y} / \partial \Omega$
1960–2017	0.563	0.104	90.016	74.517	52.905	0.681	0.177	0.504
1960–1989	0.551	0.062	56.650	56.061	36.220	0.557	0.097	0.460
1989–2017	0.587	0.173	124.577	93.773	70.195	0.780	0.307	0.473
2017	0.563	0.104	124.104	88.590	64.498	0.789	0.200	0.589

$e_{C,W} = \frac{\partial \log C}{\partial \log W} \cong \left( \frac{\frac{\partial C}{C}}{\frac{\partial W}{W}} \right) = \frac{\partial C}{\partial W} \cdot \frac{W}{C}$  and  $e_{C,R} = \frac{\partial \log C}{\partial \log R} \cong \left( \frac{\frac{\partial C}{C}}{\frac{\partial R}{R}} \right) = \frac{\partial C}{\partial R} \cdot \frac{R}{C}$  and substituted in equation (4), which is expected to be positive. The ratios  $\frac{C}{W}$  and  $\frac{C}{R}$  are evaluated at the sample mean. Table 2 summarizes this process for all estimated models.

$$\frac{\partial C/Y}{\partial \Omega} = e_{C,W} \frac{C}{W} - e_{C,R} \frac{C}{R} \quad (4)$$

The estimation results confirm the Kaldorian–Kaleckian differential savings hypothesis, namely the hypothesis of a lower propensity to consume out of profits than out of wages.<sup>11</sup> The impact of an increase of one percentage point in the wage share on the private consumption share in GDP is quite robust along the whole period and the sub-periods. In particular, the private consumption share increases by 0.50 percentage points (pp.) in the whole sample period, by 0.46 pp. in the 1960–1989 period, by 0.47 pp. in the 1989–2017 period and by 0.59 pp. at the end of the sample.

Obst et al. (2020) find similar results for Greece ( $e_{C,W}$  0.40 and  $e_{C,R}$  0.19). Likewise, Onaran/Obst (2016) find  $e_{C,W}$  0.63 and  $e_{C,R}$  0.11 and the direct effect on the consumption share equal to 0.56. In Marsellou (2013) the elasticity is similar, ranging between 0.38 and 0.62, depending on the specific formulation (ECM, DOLS, Johansen).

### 4.3 Investment

We estimate an accelerator-cash flow investment function. In particular, private investment is a double-log function that depends on demand and profits. The real interest rate is included to control for developments in financial markets. Because the variables of demand and profits are highly correlated, creating a serious problem of multicollinearity, we transform the demand variable into a growth rate following Bhaskar/Glyn (1995). So, the general form of the investment function is given by:

$$I_t = I(\text{growth}_t, R_t, \text{rint}_t) \quad (5)$$

where  $I_t$  stands for the gross fixed capital formation of the private sector,  $\text{growth}_t$  stands for the expectations for demand proxied by the growth rate of GDP,  $R_t$  stands for profits (gross operating surplus adjusted for the self-employed), and  $\text{rint}_t$  for the real long-term interest rate. According to economic theory, demand and profits are expected to have a positive effect on investment while the effect of the interest rate is expected to be negative.

ADF and PP unit root tests suggest that  $\log I_t$  and  $\text{rint}_t$  are I(1) while  $\text{growth}_t$  and  $\log R_t$  are I(0). As a result, to test for cointegration among the variables, we use the F-Bounds cointegration test suggested by Pesaran/Shin (1999) and Pesaran et al. (2001).

11. See Kalecki (1991 [1943]), Kaldor (1955/1956, 1957) and Robinson (1986 [1969]).

The test reveals the variables are cointegrated so there is a long-run equilibrium relation (Table A5).

The general form of the conditional unrestricted equilibrium ECM of the ARDL model is the following:

$$\begin{aligned} \Delta(\log I)_t = & \eta_0 + \eta_1 \log I_{t-1} + \eta_2 growth_{t-1} + \eta_3 \log R_{t-1} + \eta_4 rint_{t-1} \\ & + \sum_{j=1}^p \delta_j \Delta(\log I)_{t-j} + \sum_{j=0}^{q_1} \varphi_j \Delta(growth)_{t-j} + \sum_{j=0}^{q_2} \theta_j \Delta(\log R)_{t-j} \\ & + \sum_{j=0}^{q_3} \omega_j \Delta(rint)_{t-j} + \varepsilon_t \end{aligned}$$

Where we estimate the short- and long-run coefficients of the variables using the following form of the model:

$$\begin{aligned} \Delta(\log I)_t = & \eta_0 + \eta_1 (\log I_{t-1} + \gamma_1 growth_{t-1} + \gamma_2 \log R_{t-1} + \gamma_3 rint_{t-1}) \\ & + \sum_{j=1}^p \delta_j \Delta(\log I)_{t-j} + \sum_{j=0}^{q_1} \varphi_j \Delta(growth)_{t-j} \\ & + \sum_{j=0}^{q_2} \theta_j \Delta(\log R)_{t-j} + \sum_{j=0}^{q_3} \omega_j \Delta(rint)_{t-j} + u_t \end{aligned}$$

where  $\eta_1$  is the adjustment term in the long-run equilibrium relation and we obtain the coefficients of the long-run variables according to the following relations:

$$\gamma_1 = -\eta_2/\eta_1, \gamma_2 = -\eta_3/\eta_1 \text{ and } \gamma_3 = -\eta_4/\eta_1.$$

Table 3 Regression results for the investment equation

	1960–2017 ARDL(1,2,3,0)		1960–1989 ARDL(1,1,0,1)		1989–2017 ARDL(1,2,2,1)	
	D(log I)		D(log I) <sup>a</sup>		D(log I) <sup>b</sup>	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
c	0.131	7.385***	0.195	6.511***	-0.045	-5.964***
D(log R)	0.160	1.130	-0.327	-2.306**	-0.030	-0.141
D(log R) <sub>t-1</sub>	-0.748	-4.605***			-1.297	-5.976***
D(growth)	1.508	4.993***			1.595	3.802***
D(growth) <sub>t-1</sub>	-0.603	-2.214**			-0.664	-2.057*
D(growth) <sub>t-2</sub>	-0.465	-2.186**				
D(rint)			-0.007	-2.556**	0.011	2.755**
D1984	-0.146	-2.688***	-0.153	-2.492**		
D2007	0.241	4.508***			0.282	6.440***
CointEq <sub>t-1</sub>	-0.216	-9.277***	-0.262	-7.164***	-0.174	-11.229***
<i>Levels Equation</i>						
log R	0.559	5.174***	0.580	6.064***	0.709	1.989*
growth	16.680	4.912***	9.515	3.821**	24.735	7.444***
rint	0.002	0.274	0.009	0.660	0.026	2.229**

Table 3 Regression results for the investment equation (continued)

	1960–2017 ARDL(1,2,3,0)		1960–1989 ARDL(1,1,0,1)		1989–2017 ARDL(1,2,2,1)	
	D(logI)		D(logI) <sup>a</sup>		D(logI) <sup>b</sup>	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
R <sup>2</sup>	0.852		0.780		0.933	
Adj.R <sup>2</sup>	0.826		0.744		0.910	
Jarque–Berra (p)	0.930		0.771		0.917	
BG–LM(2)/ Ljung–Box Q(2)	0.669		0.026	0.119	0.091	0.295
BPG Het test/ ARCH test	0.200		0.400		0.033	0.293
RESET test (p)	0.949		0.780		0.146	
SSR	0.121		0.087		0.028	
Max VIF	7.0		6.7		4.9	
Chow test (1989)	0.001					
Obs.	57		29		28	

Notes: Statistical significance at: \*\*\* 1 per cent, \*\* 5 per cent, \* 10 per cent. The lag structure was selected according to the Akaike and Schwartz information criteria. Time dummy variables for 1984 and 2007 are correcting the outliers in the residuals of the regressions. <sup>a</sup>The HAC estimator is applied for the correction of mild serial correlation. <sup>b</sup>The HAC estimator is applied for the correction of mild serial correlation and heteroskedasticity.

The estimation results show that the explanatory power of the model is high (adj.R<sup>2</sup> between 0.74 and 0.91). In general, the residual and stability diagnostics do not show evidence of serious serial correlation (Breusch–Godfrey serial correlation LM and Ljung–Box tests) or heteroskedasticity (BPG Het test/ARCH) in the residuals. Some conflicting evidence of serial correlation in the first sub-period and heteroskedasticity in the second sub-period is treated with the application of the HAC estimator. The Jarque–Bera normality test shows that the residuals are normally distributed, the Variance Inflation Factors (VIF < 0.7) shows there is no serious issue of multicollinearity, and the RESET test shows there is no specification error.

The adjustment term of the long-run relation is negative, less than unity and statistically significant in all sample periods at 1.0 per cent level. The elasticities of investment with respect to profits and with respect to the expectations of demand are positive and statistically significant.<sup>12</sup> The real long-run interest rate is significant only in the second sub-period and quite oddly has a positive yet slight impact.<sup>13,14</sup>

12. The issue that the elasticity of investment with respect to demand is usually stronger than that with respect to profits has been discussed in several works (Stockhammer et al. 2009). It seems that in regressions that use aggregate time series demand has a stronger impact on investment than the profits, and the opposite happens when firm-level data are used (Chirinko 1993; Fazzari/Mott 1986).

13. The statistically insignificant interest rate, regarding the investment function, is not rare in the literature. See for instance Chirinko (1993), Clark et al. (1979), Shapiro et al. (1986), Fazzari (1994), Fazzari et al. (1988) and Hemming et al. (2002), among others.

14. It is worth noting that private investment also depends on public investment. Although according to the neoclassical theory public investment crowds out private investment, there is a large empirical literature even in the mainstream strand (starting from the seminal works of Aschauer 1989a, b and Easterly/Rebelo 1993) that supports the crowd-in effect of public investment on private investment. However, while initially public investment was included in the private investment

Table 4 Calculation of the direct partial effects

	$e_{IR}$	$I/R$ (av.)	$\frac{\partial I}{\partial \pi}$
1960–2017	0.559	0.429	0.240
1960–1989	0.580	0.553	0.320
1989–2017	0.709	0.362	0.257
2017	0.559	0.231	0.129

The elasticity of investment with respect to the profit share is converted into direct partial effect as follows:

$$\frac{\partial I}{\partial \pi} = e_{IR} \frac{I}{R} \tag{6}$$

In Obst et al. (2020) and Onaran/Obst (2016) the elasticity of investment with respect to the profit share, in first differences formulation, has no statistically significant effect on private investment. In Marsellou (2013) for the 1963–2003 period, however, the impact of the profit share is statistically significant and equals 0.15 and 0.31 in the ECM and the DOLS specifications, respectively. Table 4 reports the direct partial effect of the profit share on investment that shows a tendency to decline over time, ranging from 0.32 to 0.13.

4.4 Net exports

For the estimation of the effect of income distribution on net exports, a four-step indirect estimation strategy is applied instead of a direct estimation because the latter may fail to correctly specify the full effects on net exports (Stockhammer et al. 2009; Onaran et al. 2011; Onaran/Galanis 2012).

Domestic prices  $P$  and export prices  $P_x$  are determined by the following relations:

$$\log P = p_0 + p_{ulc} \log ULC + p_m \log P_m \tag{7}$$

$$\log P_x = p_{x0} + p_{xp} \log P + p_{xm} \log P_m \tag{8}$$

where  $ULC$  is the unit labour cost<sup>15</sup> and  $P_m$  the import prices. The exports equation is defined as:

$$\log X = x_0 + x_{yf} \log Y^f + x_{pxpm} \log P_x / P_m \tag{9}$$

where  $Y^f$  stands for the GDP of 15-EU countries which are Greece’s main trade partners, and it is used as a proxy for external demand. When the income of the trading partners increases, exports are expected to increase. The coefficient of export prices is expected to be negative as higher export prices imply a decline in the competitiveness of Greek products. The coefficient of the import prices is expected to be negative to the extent they represent

equation, it turned out statistically insignificant, both as a short- and a long-run variable in all sample periods. Perhaps, this is due to the very strong impact of demand and profits. In consequence, it seems reasonable to exclude it from the final regression.

15.  $ULC$  is defined as  $ULC = \left(\frac{W}{PY}\right) \left(\frac{ET}{EE}\right)$ , where  $\frac{W}{PY}$  the wage share and  $\frac{ET}{EE}$  the ratio of total employment ( $ET$ ), to dependent employment ( $EE$ ). The real  $ULC$  is defined as  $RULC = \left(\frac{W}{Y}\right) \left(\frac{ET}{EE}\right)$  with GDP at current market prices.

capital equipment or intermediate raw materials, making the cost of production of export products more expensive. On the other hand, if the prices of the imported goods for direct consumption increase, the Greek products will be relatively cheaper and hence its impact might be positive. The ratio of those two price variables reflects the terms of trade,  $P_x/P_m$ . Exports are negatively related to ULC through its impact on domestic prices and hence export prices. An increase in the wage share tends to increase the cost of exported products causing exports to fall.

The imports equation is defined as:

$$\log M = m_0 + m_Y \log Y + m_{ppm} \log(P/P_m) + m_{reer} \log REER \quad (10)$$

where  $M$  stands for imports and depends on domestic demand,  $Y$ , which is expected to have a positive sign and the ratio of domestic prices to import prices which is expected to have a positive effect on imports. The real exchange rate ( $REER$ ) is also included as a control variable.

All variables are in 2015 constant prices and in logarithms. Unit root tests indicate that all series are  $I(1)$ . The Johansen cointegration test provides evidence of cointegration among the variables of export equation (Table A6); however, experimentation with the error correction models did not give satisfactory results. Hence, the first differences formulation was chosen, the results of which were relatively better (Table 5). The predictability of the model as reflected by the adj. $R^2$  ranges between 0.28 and 0.56. The diagnostic and stability tests do not show any evidence of violation of the basic assumptions of the classical linear regression model. The elasticity of exports with respect to the terms of trade has

Table 5 Regression results for the exports equation

	1960–2017		1960–1989		1990–2017	
	D(log $X$ )		D(log $X$ )		D(log $X$ )	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
c	0.010	0.548	0.068	1.694	-0.011	-0.855
D(log $Y^f$ )	2.082	3.474***	0.382	0.366	3.429	5.009***
D(log $P_x/P_m$ )					-1.373	-1.840*
D(log $P_x/P_m$ ) $_{t-1}$	-0.716	-1.840*	-0.980	-2.130**		
D1982	-0.166	-1.905*	-0.188	-1.913*		
D1999					0.131	8.931***
D2002					-0.085	-6.458***
$R^2$	0.318		0.361		0.628	
Adj. $R^2$	0.280		0.281		0.563	
Jarque–Bera (p)	0.896		0.831		0.737	
BG–LM test (2)	0.877		0.978		0.228	
White test	0.152		0.546		0.591	
RESET test (p)	0.194		0.445		0.387	
SSR	0.348		0.185		0.076	
Max VIF	1.1		1.2		1.07	
Chow test (1989)	0.076					
Obs.	56		28		28	

Notes: Statistical significance at: \*\*\* 1 per cent; \*\* 5 per cent; \* 10 per cent. D1982 is selected according to the Bai–Perron multiple breakpoint test, and D1999 and D2002 correct the outliers in the residuals of the 1990–2017 sub-period regression.

the expected sign, and it is statistically significant in all sample periods.<sup>16</sup> The elasticity of external demand is positive and statistically significant (1.0 per cent level) in the full and the second sub-sample.

Table 6 presents the regression results for the imports function. The F-Bounds cointegration test shows no evidence of cointegration between the variables. Hence, we estimate the imports equation with a first difference formulation. The adj.R<sup>2</sup> of the estimated regressions ranges between 0.37 and 0.75. The diagnostic tests showed no evidence of serious violation of the basic assumptions of the classical linear regression model.

The elasticity of imports with respect to internal demand is positive, strong, and statistically significant in all sample periods. The elasticity with respect to the terms of trade is statistically significant at the full and the first sub-sample periods, while at the second sub-period it has a negative sign, and it is insignificant. At the same period the elasticity with respect to internal demand becomes almost two times higher than in the full and the first sub-sample, implying that at the second sub-period the impact of internal demand

Table 6 Regression results for the imports equation

	1960–2017		1960–1989		1990–2017	
	D(logM) <sup>1</sup>		D(logM) <sup>2</sup>		D(logM) <sup>3</sup>	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
c	0.021	1.980*	0.026	1.646	0.037	4.108***
D(logY)			0.774	3.957***	1.616	6.317***
D(logY) <sub>t-1</sub>	0.841	4.322***				
D(logP/P <sub>m</sub> )	0.384	1.919*	0.424	1.826*	-0.295	-1.614
D(logREER)	-0.452	-3.102***	-0.157	-1.636	-0.113	-0.802
D1973			0.191	15.106***		
D2000	0.142	2.009**				
D2002					-0.119	-12.513***
D2009					-0.177	-12.512***
R <sup>2</sup>	0.413		0.691		0.793	
Adj.R <sup>2</sup>	0.367		0.640		0.746	
BG-LM test(2)	0.165		0.063		0.845	
White test	0.842		0.117		0.469	
RESET test (p)	0.100		0.206		0.690	
SSR	0.237		0.054		0.043	
Max VIF	1.06		1.4		1.1	
Chow test (1989)	0.042					
Obs.	56		29		28	

Notes: Statistical significance at: \*\*\* 1 per cent; \*\* 5 per cent; \* 10 per cent. <sup>1</sup>D2000 is one of the possible breaks suggested by the Bai–Perron tests of 1 to *M* globally determined breaks and is confirmed by the Chow test. <sup>2</sup>D1973 is one of the possible breaks suggested by the Bai–Perron tests of 1 to *M* globally determined breaks; it is confirmed by the Chow test and evident by the inspection of the residuals of the regression. <sup>3</sup>Time dummies D2002 and D2009 correct the outliers in the residuals of the regression.

16. Note at this point that we follow Obst et al. (2020), Onaran/Obst (2016), and Stockhammer et al. (2011) in handling the statistically insignificant coefficients and consider them as zero when calculating the total marginal effects on total excess demand.

increases and that of prices falls. A possible hypothesis that cannot be tested here might be that this development is attributed to the gradual fall in the interest rates and the expansion of household credit. The elasticity with respect to the exchange rate is statistically significant only at the full sample period.

These findings are in line with the studies surveyed above that include Greece in their sample. Obst et al. (2020) and Onaran/Obst (2016) for the period 1961–2013 in first differences formulation find the elasticity of imports with respect to the terms of trade to be 0.15 but not statistically significant, and the elasticity with respect to internal demand equal to 1.27 at 1.0 per cent level.

As is expected, price variables are highly correlated, hence we estimated the export prices equation in first differences. The predictability of the model is high (adj. $R^2$  between 0.87 and 0.96). The diagnostic and stability tests show that the basic assumptions of the classical linear regression model are not violated, except for the evidence of mild serial correlation in the regression of the full sample that was corrected with a second order autoregressive term (AR(2)).

Both the elasticity of the export prices with respect to the domestic prices and imports prices is positive and statistically significant at the 1.0 per cent and 5.0 per cent levels, though the latter has a greater impact. These estimations remain stable over time.

The unit root tests indicate that the variables of the domestic price equation can be considered as I(1) and the Johansen cointegration test implies that a long-run relationship exists. However, due to the high correlation among the regressors and the serious problem of multicollinearity that causes, it is more reasonable to estimate the equation with a first differences formulation instead with an ECM or an ARDL. The diagnostic tests are

*Table 7 Regression results for the exports prices equation*

	1960–2017		1960–1989		1990–2017	
	D(logPx)		D(logPx)		D(logPx)	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
C	–0.000	–0.025	–0.009	–0.800	0.004	1.607
D(logP)	0.251	3.243***	0.255	2.172**	0.255	5.217***
D(logPm)	0.718	10.672***	0.798	8.995***	0.681	12.899***
D1973	0.032	1.273				
D1982	–0.025	–1.045				
D1975			–0.075	–8.218***		
AR(2)	–0.324	–2.722***				
R <sup>2</sup>	0.914		0.888		0.963	
Adj.R <sup>2</sup>	0.904		0.874		0.960	
Jarque–Bera (p)	0.071		0.968		0.238	
BG–LM test(2)/Qstat(2)	0.446		0.264		0.288	
BPG Het test	0.126		0.598		0.843	
RESET test (p)	0.091		0.360		0.768	
SSR	0.033		0.030		0.003	
Max VIF	6.2		1.6		2.2	
Chow test (1989)						
Obs	57					

*Notes:* Statistical significance at: \*\*\* 1 per cent; \*\* 5 per cent; \* 10 per cent. Dummy variables for 1973, 1982 and 1975 are selected according to the Bai–Perron multiple breakpoint test.

Table 8 Regression results for the domestic prices equation

	1960–2017		1960–1989		1990–2017	
	D(logP) <sup>a</sup>		D(logP)		D(logP) <sup>a</sup>	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
C	0.006	1.574	0.019	2.079**	–0.002	–0.611
D(logULC) <sub>t</sub>	0.338	4.795***	0.381	4.103***	0.229	2.285**
D(logPm) <sub>t</sub>	0.105	2.014**	0.262	3.878***		
D(logPm) <sub>t–1</sub>					0.136	0.926
D(logP) <sub>t–1</sub>	0.505	6.316***			0.646	6.241***
D1973	0.108	5.240***				
D2015					0.022	2.858***
Trend			0.002	2.423**		
R <sup>2</sup>	0.939		0.918		0.937	
Adj.R <sup>2</sup>	0.934		0.908		0.927	
Jarque–Bera (p)	0.783		0.185		0.035	
BG–LM test (2)	0.326		0.339		0.841	
BPG Het test/ARCH test	0.051	0.439	0.183	0.845	0.083	0.409
RESET (p)	0.907		0.681		0.082	
SSR	0.018		0.013		0.006	
Max VIF	5.2		4.3		5.2	
Chow test (1989)	0.002					
Obs.	57		29		28	

Notes: Statistical significance at: \*\*\* 1 per cent; \*\* 5 per cent; \* 10 per cent. Dummy variable for 1973 according to the Bai–Perron multiple breakpoint test. The time trend corrects for the trend evidenced in the variables in first differences. Dummy variable for 2015 corrects for an outlier.  
<sup>a</sup>The HAC estimator was applied due to evidence of mild heteroskedasticity.

satisfactory, while minor evidence of heteroskedasticity is corrected by applying the HAC estimator. The elasticity of the domestic prices with respect to the ULC and the imports prices have the expected signs and are statistically significant at the 1.0 per cent and 5.0 per cent levels. These findings are lower than the 0.42 found in Obst et al. (2020) and Onaran/Obst (2016).

The above estimated elasticities are used for the calculation of the marginal effect of a change in the wage share on the exports and imports shares in GDP. These calculations apply the derivative chain rule as follows:

$$\begin{aligned} \frac{\partial X}{\partial ws} &= \frac{\partial \log X}{\partial \log ws} * \frac{X}{ws} = \left( \frac{\partial \log X}{\partial \log P_x} * \frac{\partial \log P_x}{\partial \log P} * \frac{\partial \log P}{\partial \log ULC} * \frac{\partial \log ULC}{\partial \log RULC} * \frac{\partial \log RULC}{\partial \log ws} \right) \frac{X}{ws} \\ &= \left( e_{XP_x} * e_{P_x P} * \frac{e_{PULC}}{1 - e_{PULC}} * \frac{1}{RULC} * \frac{ET}{ED} \right) * \frac{X}{Y} \end{aligned} \quad (11)$$

and

$$\begin{aligned} \frac{\partial M}{\partial ws} &= \frac{\partial \log M}{\partial \log ws} * \frac{M}{ws} = \left( \frac{\partial \log M}{\partial \log P} * \frac{\partial \log P}{\partial \log ULC} * \frac{\partial \log ULC}{\partial \log RULC} * \frac{\partial \log RULC}{\partial \log ws} \right) \frac{M}{ws} \\ &= \left( e_{M,P} * \frac{e_{PULC}}{1 - e_{PULC}} * \frac{1}{RULC} * \frac{ET}{ED} \right) * \frac{M}{Y} \end{aligned} \quad (12)$$



Table 9 Calculation of the direct effects on net exports

Exports	$e_{XP_x}$	$e_{P_xP}$	$e_{PRULC}$	$1/RULC$	$X/Y$	$ET/ED$	
2017	-0.716	0.251	0.511	1.805	0.326	1.087	-0.059
1960–2017	-0.716	0.251	0.511	1.672	0.141	1.108	-0.024
1960–1989	-0.980	0.255	0.614	1.560	0.077	1.122	-0.021
1990–2017	-1.373	0.255	0.296	1.793	0.210	1.093	-0.043
Imports	$e_{M,P}$		$e_{PRULC}$	$1/RULC$	$M/Y$	$ET/ED$	
2017	0.384		0.511	1.805	0.334	1.087	0.129
1960–2017	0.384		0.511	1.672	0.182	1.108	0.066
1960–1989	0.424		0.614	1.560	0.101	1.122	0.046
1990–2017	–		0.296	1.793	0.269	1.093	0
$X-M$							
2017							-0.187
1960–2017							-0.090
1960–1989							-0.067
1990–2017							-0.043

$$\text{where } e_{PRULC} = \frac{\partial \log P}{\partial \log RULC} = \frac{\partial \log P}{\partial \log ULC} * \frac{\partial \log ULC}{\partial \log RULC} = \frac{e_{PULC}}{1 - e_{PULC}}.$$

Table 9 summarizes the effects of a 1.0 per cent increase in the wage share on the share of exports, imports and finally the net exports in GDP for all sample periods. Obviously, the final effect depends on the elasticities of exports and imports to relative prices, and the impact of labour costs on prices but also on the shares of exports and imports in GDP at each respective sample-period. The elasticity of imports with respect to domestic prices is not presented in the second sub-period because it is not statistically significant.

#### 4.5 Total effects

Table 10 presents the marginal effects of the wage share on each aggregate demand component calculated in the previous sections. The domestic sector of the economy is wage-led since the positive effect of wages on consumption is stronger than the negative effect on investment. The estimates suggest that a one percentage point increase in the wage share leads to an increase in the share of domestic demand in GDP by 0.26 pp. in the full sample-period 1960–2017, by 0.14 pp. during the 1960–1989 period, by 0.22 pp. during the 1990–2017 period, and by 0.46 in 2017. Adding the respective effect on the net exports share in GDP implies that the private excess demand ( $b_2$ ) is positive (0.17 pp. (1960–2017), 0.07 pp. (1960–1989), 0.17 pp. (1990–2017) and 0.27 in 2017) and hence the demand regime is wage-led in all sample periods.<sup>17</sup>

17. The total private effect of a one percentage-point increase of the wage share on equilibrium demand is obtained as stated in equation (2) by multiplying  $b_2$  with the standard multiplier  $\frac{1}{1-b_1}$ . However, because the analysis that preceded, in its most part, is conducted through short-run estimations, which are not compatible with the long-run equilibrium relations, the total effect estimations are not exercised.

Table 10 Private excess demand caused by a one percentage-point increase of the wage share.

	1960–2017	1960–1989	1990–2017	2017
Consumption	0.504	0.460	0.473	0.589
Investment	0.240	0.320	0.257	0.129
<i>Domestic Excess Demand</i>	0.264	0.139	0.216	0.460
Net Exports	–0.090	–0.067	–0.043	–0.187
<i>Private Excess Demand (<math>h_2</math>)</i>	0.174	0.072	0.173	0.273

5 SUMMARY AND CONCLUSIONS

This paper analysed the relationship between functional income distribution and economic growth in the small open economy of Greece. The analysis was based on a demand-driven distribution and growth model for an open economy inspired by Bhaduri/Marglin (1990), which allows for profit- or wage-led growth. The empirical analysis in this paper shows that a decline in the wage share in the Greek economy leads to a decline in growth, implying that the demand regime is wage-led. Moreover, the analysis shows that the results are robust across different sample sizes, for the period 1960–2017, and the two sub-periods 1960–1989 and 1990–2017.

The results of this analysis may shed some light on the possible causes of the rather ineffective wage moderation policies from a demand-side orientation. In particular, the internal devaluation policies implemented in the Greek economy during the three Economic Adjustment Programs have caused a dramatic fall in the Greek GDP and a prolonged recession that lasted nearly a decade. Supply-side economic theory, however, predicts that wage moderation policies, even in their extreme form of internal devaluation, should stimulate investments, increase competitiveness and exports, that eventually pass-through to growth and along with that to private and public debt sustainability. However, what occurred was a collapse in demand, a record high unemployment rate, and a dramatic deterioration in the public-debt-to-GDP ratio. In addition to these consequences, households’ private-debt-to-income ratios rocketed due to wage cuts and job losses creating unprecedented numbers of bankruptcies.

The results of this paper do not suggest the adoption/maintenance on a consumer-led growth model but rather point out that an attempt to change the existing productive model towards a competitive one through wage moderation policies will have adverse economic consequences and a great social cost, as the recent Greek experience has clearly shown. Policies that are compatible with the wage-led character of the Greek economy are those that strengthen the wage share; namely those that support the welfare state,<sup>18</sup> labour market institutions, increase unemployment benefits, and enact and support minimum wages.<sup>19</sup>

The shift towards a more productive and competitive growth model should be pursued through the governments’ prioritization of public investments in infrastructure that

18. Furceri/Zdzienicka (2012) estimate the short-run impact of government social spending on GDP for a panel of OECD countries for the period 1980–2005 and find that they have expansionary impact on economic activity, especially during recessions.

19. The rise of the income of the lower-income classes reduces personal income inequality (workers’ income instead of managerial income) and stimulates economic activity and growth (Palley 2017).

increase the overall productivity of the economy (Deleidi et al. 2020; IMF 2014; European Commission 2014) but also provide incentives to the private sector in specific areas and activate chain improvements in labour productivity at the firm, industry, and sectoral level. Hein/Tarassow (2010) have shown that a higher wage share does not only stimulate economic activity and growth in a wage-led environment in the short run but also has favourable effects on capital accumulation and productivity growth in the long-run. During the global financial crisis of 2007–2008, and more recently the COVID-19 pandemic, the so-called emergency Keynesianism has arisen. For a wage-led recovery to take place, however, more permanent pro-labour policies are required than a temporal Keynesian program.

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## APPENDIX

Table A1 Definitions of variables

Symbol	Description
<b><i>W</i></b>	$W = \text{UWCD}/\text{PVGD}$ Compensation of employees: total economy (UWCD). Adjusted for imputed compensation of self-employed Price deflator GDP at market prices (2015=100) (PVGD)
<b><i>R</i></b>	$R = \text{UQGD}/\text{PVGD}$ Gross operating surplus: total economy: Adjusted for imputed compensation of self-employed (UQGD) Price deflator GDP (2015=100) (PVGD)
<b><i>C</i></b>	Private final consumption expenditure at 2015 prices (OCPH)
<b><i>Y</i></b>	$Y = \text{UVGD}/\text{PVGD}$ Gross domestic product at current market prices (UVGD) Price deflator GDP (2005=100) (PVGD)
<b><i>P</i></b>	Price deflator GDP (2015=100) (PVGD)
<b><i>I</i></b>	$Y = \text{UIGP}/\text{PVGD}$ Gross fixed capital formation: private sector (UIGP) Price deflator GDP (2015=100) (PVGD)
<b><i>i</i></b>	Real interest rate (%), BoG 1960–2017, GDP deflator (PVGD)
<b><i>X</i></b>	Exports of goods and services at 2015 prices (OXGS)
<b><i>M</i></b>	Imports of goods and services at 2015 prices (OMGS)
<b><i>NX</i></b>	Net exports, real, $NX = X - M$
<b><i>REER</i></b>	Real effective exchange rates (XUNRQ)
<b><i>Y<sup>f</sup></i></b>	GDP at 2015 market prices (OVGD) Euro area (12-countries)
<b><i>RULC</i></b>	Real unit labour costs (QLCD)
<b><i>ULC</i></b>	Nominal unit labour costs (PLCD)
<b><i>E<sub>D</sub></i></b>	Employees (NWTG)
<b><i>E<sub>T</sub></i></b>	Employment (NETD)
<b><i>P<sub>g</sub></i></b>	Gross fixed capital deflator (PIGT)
<b><i>P<sub>x</sub></i></b>	Price deflator exports of goods and services (PXGS)
<b><i>P<sub>m</sub></i></b>	Price deflator imports of goods and services (PMGS)

*Note:* All deflators are at base year 2015.

*Sources:* AMECO, BoG.

Table A2 Augmented Dickey–Fuller unit root test

Variable	Levels			Differences		
	Deterministic	Lags	t-stat.	Deterministic	Lags	t-stat.
logC	c, t	3	-1.242	c, t	0	-3.617**
logW	c, t	1	-2.075	c	0	-3.899***
logR	c, t	0	-4.118**	c, t	2	-3.632**
logI	c, t	0	-1.767	c	2	-2.691*
logY	c, t	1	-1.712	c, t	1	-3.422*
growth	c, t	2	-5.152***	–	0	-12.106***
Int	c	1	-1.285		3	-2.469**
logX	c, t	2	-1.767	c	2	-3.834***
logM	c, t	2	-1.507	c	2	-3.454**
logY <sup>f</sup>	c, t	2	-2.205	c	1	-3.644***
logULC	c, t	2	-1.593		1	-1.125
RULC	c, t	0	-3.998**		3	-3.126***
log(P <sub>X</sub> /P <sub>M</sub> )	c	2	-1.063	c	2	-5.030***
logP <sub>x</sub>	c, t	0	-1.355		6	-1.153
logP <sub>m</sub>	c, t	3	-0.379		4	-1.034
logP <sup>a</sup>	c, t	1	-1.286	c, t	2	-5.163**
log(REER)	c	0	-2.743*	c	0	-6.899***

Notes: Critical values according to MacKinnon (1996). Lag structure according to Modified AIC. Statistically significant at: \*\*\* 1 per cent, \*\* 5 per cent, \* 10 per cent. <sup>a</sup>The breakpoint unit root test suggests that the logP series is I(1) with a structural break in 1972.

Table A3 Phillips–Perron unit root test

Variable	Levels			Differences		
	Deterministic	Bandwidth	t-stat.	Deterministic	Bandwidth	t-stat.
logC	c, t	3	-0.818	c, t	3	-3.727**
logW	c, t	4	-1.617	c	0	-3.899***
logR	c, t	1	-4.331***	c, t	3	-9.266***
logI	c, t	4	-2.033	c	3	-5.265***
logY	c, t	4	-0.050	c, t	3	-5.152***
growth	c, t	3	-5.264***	–	2	-12.848***
int	c	4	-1.136		1	-4.330***
logX	c, t	2	-1.418	c	2	-6.380***
logM	c, t	2	-1.431	c	2	-5.485***
logY <sup>f</sup>	c, t	2	-2.355	c	2	-3.990***
logULC	c, t	5	-0.507		1	-1.677*
RULC	c, t	0	-3.998**		3	-8.115***
log(P <sub>X</sub> /P <sub>M</sub> )	c, t	2	-1.473	c	1	-7.632***
logP <sub>x</sub>	c, t	0	-1.335		4	-5.087***
logP <sub>m</sub>	c, t	4	-0.170		5	-2.032**
logP	c, t	3	0.341		1	-1.030
log(REER)	c	2	-2.751*	c	1	-6.900***

Notes: Critical values according to MacKinnon (1996). HAC corrected variance (Bartlett kernel). Statistically significant at: \*\*\* 1 per cent, \*\* 5 per cent, \* 10 per cent.



*Table A4 F-Bounds test results, consumption equation*

ARDL Bounds Test		
Sample: 1961–2017, Observations: 57		
Null hypothesis: No long-run relationships exist		
Test statistic	Value	k
F-statistic	4.56	2
Critical value bounds		
significance	I(0) bound	I(1) bound
10%	3.17	4.14
5%	3.79	4.85
2.5%	4.41	5.52
1%	5.15	6.36

*Table A5 F-Bounds test results, investment equation*

	F-Statistic	Lower-critical value 1%	Upper-critical value 1%	Obs.
1960–2017	20.079	4.828	6.195	54
1960–1989	11.227	5.333	7.063	29
1989–2017	26.792	5.333	7.063	28

*Table A6 Johansen cointegration test, exports equation*

Sample (adjusted): 1962–2017				
Included observations: 56 after adjustments				
Trend assumption: No deterministic trend (restricted constant)				
Series: $\log X$ $\log Y^f$ $\log P_x/P_m$				
Lags interval (in first differences): 1 to 1				
Unrestricted Cointegration Rank Test (Trace)				
Hypothesized no. of CE(s)	Eigenvalue	Trace statistic	0.05 critical value	Prob.**
None*	0.364210	40.82902	35.19275	0.0111
At most 1	0.199290	15.46734	20.26184	0.2008
At most 2	0.052517	3.020973	9.164546	0.5768

Notes: Trace test indicates one cointegrating equation at the 0.05 level. \*Denotes rejection of the hypothesis at the 0.05 level. \*\*MacKinnon–Haug–Michelis (1999)  $p$ -values.

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)				
Hypothesized no. of CE(s)	Eigenvalue	Max-Eigen statistic	0.05 critical value	Prob.**
None *	0.364210	25.36168	22.29962	0.0181
At most 1	0.199290	12.44637	15.89210	0.1615
At most 2	0.052517	3.020973	9.164546	0.5768

Notes: Max-Eigenvalue test indicates one cointegrating equation at the 0.05 level. \*Denotes rejection of the hypothesis at the 0.05 level. \*\*MacKinnon–Haug–Michelis (1999)  $p$ -values.

*Table A7 F-Bounds test results, imports equation*

ARDL Bounds Test		
Sample: 1961–2017		
Included observations: 57		
Null hypothesis: No long-run relationships exist		
Test Statistic	Value	k
F-statistic	1.126311	3
Critical Value Bounds		
Significance	I(0) bound	I(1) bound
10%	2.72	3.77
5%	3.23	4.35
2.5%	3.69	4.89
1%	4.29	5.61

*Table A8 Johansen cointegration test, domestic prices equation*

Sample (adjusted): 1963–2019				
Included observations: 57 after adjustments				
Trend assumption: Linear deterministic trend (restricted)				
Series: $\log P$ $\log P_m$ $\log ULC$				
Lags interval (in first differences): 1–2				
Unrestricted Cointegration Rank Test (Trace)				
Hypothesized no. of CE(s)	Eigenvalue	Trace statistic	0.05 critical value	Prob.**
None*	0.375804	48.18814	42.91525	0.0136
At most 1	0.220350	21.32457	25.87211	0.1661
At most 2	0.117684	7.136678	12.51798	0.3305

Notes: Trace test indicates one cointegrating equation at the 0.05 level. \*Denotes rejection of the hypothesis at the 0.05 level. \*\*MacKinnon–Haug–Michelis (1999) *p*-values.

Unrestricted Cointegration Rank Test (Maximum-Eigenvalue)				
Hypothesized no. of CE(s)	Eigenvalue	Max-Eigen statistic	0.05 critical value	Prob.**
None*	0.375804	26.86357	25.82321	0.0364
At most 1	0.220350	14.18789	19.38704	0.2418
At most 2	0.117684	7.136678	12.51798	0.3305

Notes: Max-Eigenvalue test indicates one cointegrating equation at the 0.05 level. \*Denotes rejection of the hypothesis at the 0.05 level. \*\*MacKinnon–Haug–Michelis (1999) *p*-values.

*Table A9 Johansen cointegration test, export prices function*

Sample (adjusted): 1963–2019

Included observations: 57 after adjustments

Trend assumption: Linear deterministic trend (restricted)

Series:  $\log Px$ ,  $\log Pm$ ,  $\log P$ 

Lags interval (in first differences): 1–2

Unrestricted Cointegration Rank Test (Trace)

Hypothesized no. of CE(s)	Eigenvalue	Trace statistic	0.05 critical value	Prob.**
None*	0.437110	58.92531	42.91525	0.0006
At most 1*	0.289055	26.16905	25.87211	0.0459
At most 2	0.111256	6.722929	12.51798	0.3743

*Notes:* Trace test indicates two cointegrating equations at the 0.05 level. \*Denotes rejection of the hypothesis at the 0.05 level. \*\*MacKinnon–Haug–Michelis (1999)  $p$ -values.

Unrestricted Cointegration Rank Test (Maximum-Eigenvalue)

Hypothesized no. of CE(s)	Eigenvalue	Max-Eigen statistic	0.05 critical value	Prob.**
None*	0.437110	32.75626	25.82321	0.0052
At most 1*	0.289055	19.44612	19.38704	0.0490
At most 2	0.111256	6.722929	12.51798	0.3743

*Notes:* Max-Eigenvalue test indicates two cointegrating equations at the 0.05 level. \*Denotes rejection of the hypothesis at the 0.05 level. \*\*MacKinnon–Haug–Michelis (1999)  $p$ -values.