Review of Keynesian Economics, Vol. 8 No. 2, Summer 2020, pp. 220–239

Autonomous expenditures and induced investment: a panel test of the Sraffian supermultiplier model in European countries

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This paper tests the main postulates of the Sraffian supermultiplier model for the case of 16 European economies during the period 1995–2018. We adopt the methodology of Girardi and Pariboni (2016) and extend it to a panel framework. We apply panel unit root, cointegration, and causality tests that are robust to endogenous regressors, cross-sectional dependence and heterogeneity across countries. Our results are supportive of the Sraffian supermultiplier model. In a heterogeneous panel framework, autonomous demand and output follow a long-run equilibrium relationship and there exists panel long-run causality that goes unidirectionally from autonomous demand to output. We also empirically verify the investment accelerator (the mechanism that enables the dynamic stability of the model) by confirming the existence of same-sign panel causality running unidirectionally from the growth rate of autonomous demand to the investment share. Our results call for national economic policies aimed at promoting the components of autonomous demand that act as locomotives of growth in each country.

Keywords: *long-run economic growth, Sraffian supermultiplier model, panel cointegration, panel causality*

JEL codes: *B12, B23, E12*

1 INTRODUCTION

In this paper, we empirically test for the case of 16 European economies the Sraffian supermultiplier model (Serrano 1995; Cesaratto et al. 2003; Dejuán 2017; Palley 2019, among others). The supermultiplier model introduces a mechanism that leads the economy towards normal capacity utilization, while safeguarding the main Keynesian message regarding the principle of effective demand (Lavoie 2016, p. 172).[1] Consequently, the supermultiplier approach to growth and distribution can be viewed as providing a possible consensus model for reconciling the Neo-Kaleckian and Sraffian positions.

The modern history of economic growth theory begins with a famous article by Roy Harrod in 1939 (Roncaglia 2006). By combining the multiplier principle with the accelerator mechanism, Harrod (1939) tried to extend the principle of effective demand to the

1. Recently, Schoder (2014) has provided empirical evidence on the compatibility between the principle of effective demand and the stationarity of the rate of capacity utilization.

long run (Vernengo and Rochon 2001). Harrod developed a dynamic model in which the effects that investment has on demand, on the one hand, and on productive capacity (supply), on the other hand, could be reconciled so as to maintain supply and demand balance over time. However, Harrod's model suffers from dynamic instability, since any increase (decrease) in investment increases (decreases) productive capacity and demand in equal proportions. Therefore, the long-run equilibrium with a normal (desired by entrepreneurs) degree of capacity utilization is dynamically unstable.

The solution to the Harrodian instability problem we present here is the Sraffian supermultiplier model, proposed by Franklin Serrano in 1995. Serrano (1995) states that the problem of fundamental instability of the Harrod equation can be solved by considering the existence of non-capacity-generating autonomous demand variables. Autonomous expenditures, Z, are expenditures that do not generate productive capacity and whose evolution is independent of that of effective output. According to Serrano (1995), Cesaratto et al. (2003), Freitas and Serrano (2015), Girardi and Pariboni (2016), Lavoie (2016), Pérez Caldentey and Vernengo (2017), and Palley (2019), among others, autonomous demand can be considered as the components of final demand that are not mechanically linked to the evolution of output. Thus, Z is constituted by 'all those expenditures that are not financed by wage income generated by production decisions, nor affect (directly) the productive capacity of the economy' (Serrano 1995, p. 71).

In Harrod's model, instability occurs because productive capacity, driven by the decrease or increase of investment, decreases or increases always in the same proportion as demand does. By introducing Z into the picture, however, Serrano's model can achieve local asymptotical stability. Dynamic stability is reached because the reaction of investment causes a reduction (increase) in the rate of growth of capacity greater (smaller) than the reduction (increase) in the growth rate of output. In sum, in the face of unexpected changes, there is a constant tendency to re-establish the desired or normal level of productive capacity utilization (see Serrano and Freitas 2007).

Recently, some Neo-Kaleckian growth models (Allain 2015; Lavoie 2016; Fiebiger and Lavoie 2019, among others) have incorporated the existence of autonomous expenditures and the tendency towards the normal degree of capacity utilization in the long-term equilibrium path. By doing so, the neo-Kaleckian model of growth and the Sraffian supermultiplier approach share an economic rationale about how economic growth works. When considering autonomous expenditures, the causal relationship between the rates of growth of capital and output is exactly the opposite of that considered by the traditional Cambridge and Oxford models: the rate of capital accumulation is determined by the growth rate of output. In turn, the evolution of output in the long run is determined by the evolution of autonomous demand, the true engine of long-run economic growth.

On the other hand, Aspromourgos (2013) doubts that any component of final demand can be considered fully autonomous in the long run, given the intertemporal budget constraints of each sector (ibid., p. 35). Trezzini and Palumbo (2016), Skott (2019), and Nikiforos (2018), among others, also question whether autonomous demand can really be considered autonomous. For this reason, we analyse the relationship between autonomous demand and output in the long run through a bivariate model. We also empirically test the investment accelerator mechanism implied by the Sraffian supermultiplier model, by analysing the relationship between the investment share and the rate of growth of autonomous demand. In Section 2, we explain the methodology of the econometrics employed. In Section 3, we present our empirical results. Section 4 contains our conclusions.

2 DATA AND METHODOLOGY FRAMEWORK

This paper empirically evaluates the Sraffian supermultiplier model. We test some postulates of the model for 16 European countries. These countries are: Germany, Spain, France, the United Kingdom, Austria, Italy, Holland, Sweden, Denmark, Ireland, Greece, Portugal, the Czech Republic, Finland, Lithuania, and Luxembourg.

As a proxy variable for output, we use the gross domestic product, *GDP*. Following Serrano (1995), the variables that we consider as potential components of autonomous demand are public expenditure, *G*, autonomous consumption, *AC*, and exports, *E*. We build a quarterly series between 1995Q1 and 2017Q4. However, for Italy, Holland, Austria and the Czech Republic we only have data from 1996Q1 to 2017Q4. We use Eurostat quarterly data (ESA 2010). We work with the variables in real terms (measured in chained 2010 euros), seasonally and calendar adjusted, and transform them into logarithms to avoid non-linear trends and problems of heteroskedasticity.

We consider that the variable dwellings is a good proxy of autonomous consumption, since in most cases the acquisition of housing is not made by current income, but through credit or income accumulated in previous periods (see Fiebiger 2018). Moreover, this type of investment is not destined to increase productive capacity and is not given by the production decisions carried out by firms (Freitas and Serrano 2015, p. 261).

For public expenditure, we use final consumption expenditure of general government. We consider public investment as directly affecting the productive capacity of an economy. We are aware that there may be some degree of endogeneity in public expenditure, but the supermultiplier approach considers public expenditure as not mechanically linked to the evolution of output in the long run.

The variable total exports of goods and services is used to represent exports. As Girardi and Pariboni (2015; 2016) state, there are also multiple reasons to expect some endogeneity in exports, but it does not necessarily imply that they cannot be considered as elements of autonomous demand.

Finally, as a proxy variable for induced investment, *I*, which will be used in Section 3 to test the investment accelerator mechanism, we use total fixed capital accumulation less dwellings.

We first consider the relationship between output and autonomous demand in a multivariate setting, as Médici (2011) does for the Argentinian economy. A multivariate analysis would imply considering the following long-run relationship:

$$GDP_{it} = c_i + \beta_{1i}G_{it} + \beta_{2i}AC_{it} + \beta_{3i}E_{it} + \varepsilon_{it}, \tag{1}$$

where $i = 1, 2, \dots, N$ denotes each country of the panel, and $t = 1, 2, \dots, T$ denotes temporal observations. However, the multivariate approach is not the most adequate to test the main postulates of the Sraffian supermultiplier model. On the one hand, the panel long-run causality test is by construction a bivariate specification (Herzer and Nunnenkamp 2012; Narayan and Popp 2012). The multivariate approach is suitable to test weak exogeneity, but is not widely considered a suitable tool for testing long-run causality (Calderón et al. 2015).

On the other hand, and more importantly, we proceed with a bivariate setting because, according to the supermultiplier model, the rate of growth of the whole autonomous demand shapes the rate of growth of output in the long run. According to Serrano (1995), Cesaratto et al. (2003), Girardi and Pariboni (2016; 2019), Pariboni (2016), and Girardi et al. (2019), autonomous demand must be considered as the sum of government expenditure, total exports, and autonomous consumption. Therefore, we test the main insight of the supermultiplier approach: a convergence of the growth rate of output to the growth rate of the whole autonomous demand in the long run.

Thus, in what follows, we analyse the relationship between *GDP* and Z in the long run through a bivariate panel framework, with $Z = AC + G + E$. In doing so, we adopt the empirical methodology introduced by Girardi and Pariboni (2016). Thus, the novelty of this paper is empirical, since we apply Girardi and Pariboni (2016) methodology to European countries. It is also methodologically novel because we extend their time-series methodology to a heterogeneous panel framework.

2.1 Tests for cross-sectional independence and homogeneity

We expect the economies under study to be highly integrated (since they are part of the European Union (EU)) through financial, commercial, fiscal, and cultural ties, among others. That implies a country may be affected by shocks or cyclical movements that take place in another country of the panel. Therefore, our first step consists in applying tests of cross-sectional dependence. To this aim, we use the following equation:

$$Y_{it} = c_i + \beta_i X_{it} + \varepsilon_{it}, \tag{2}$$

where β_i is estimated by pooled effects ordinary least squares (OLS). Y_{it} is the dependent variable (output), and X_{it} is the independent variable (autonomous demand).

The Lagrange multiplier cross-sectional dependence test, developed by Breusch and Pagan (1980), is suitable to check the presence of cross-sectional dependence. The null hypothesis checks whether there is cross-sectional independence between the errors, $H_0 : Corr(\varepsilon_{it}, \varepsilon_{jt}) = 0 \; \forall t$ and $i \neq j$; while the alternative hypothesis is that there exists cross-sectional dependence between the errors, $H_1 : Corr(\varepsilon_{it}, \varepsilon_{jt}) \neq 0$ for at least one pair of $i \neq j$. To test the null hypothesis, Breusch and Pagan (1980) elaborated the following Lagrange multiplier test statistic:

$$B - P\,\mathrm{LM} = T \textstyle\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \varnothing_{ij}, \tag{3}$$

where $\varnothing_{ij} = Corr(\varepsilon_{it}, \varepsilon_{jt})$ is the coefficient of pairwise correlation of the residuals obtained by estimating equation (2) through pooled OLS for each country. This test is adequate when N is small and T is large.

On the other hand, Pesaran (2004) argues that the Breusch and Pagan (1980) test is not adequate when N is large. Pesaran (2004) proposes a test for panels with large T and large N:

$$\text{Pesaran CD} = \sqrt{\frac{2T}{N(N-1)}} \times \textstyle\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \varnothing_{ij} \sim N(0,1). \tag{4}$$

However, Pesaran et al. (2008) show that when the average pairwise correlations differ from zero, the Pesaran (2004) test statistic lacks power. This is corrected by the Pesaran et al. (2008) modified Lagrange multiplier test:

$$\text{Pesaran scaled LM} = \sqrt{\frac{2T}{N(N-1)}} \times \textstyle\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \varnothing_{ij} \times \frac{(T-k)\varnothing^2{}_{ij} - \mu_{Tij}}{\sqrt{\nu^2{}_{Tij}}} \sim N(0,1), \tag{5}$$

where k denotes the number of regressors, and μ_{Tij} and $v^2{}_{Tij}$ represent the mean and variance of $(T-k)\varnothing^2{}_{ij}$, respectively. When $T\to\infty$ and $N\to\infty$, the Pesaran et al. (2008) test follows a standard normal distribution under the null hypothesis.

A common source of inconsistent panel estimation results is the assumption of cross-sectional homogeneity (similar economic dynamics) when the panel is heterogeneous. We expect the economies that conform to our panel to be heterogeneous. As Boyer (2013, p. 537) states, 'the eurozone member states were more heterogeneous than many thought, in terms of international specialisation, labour market institutions, welfare organisation and financing, priorities in public spending, financial markets, etc.' Then, we proceed to analyse the presence of heterogeneity across the countries under study. For this purpose, we employ the Pesaran and Yamagata (2008) test, which is an extended version of the Swamy (1970) test. This method tests the null hypothesis of $\beta_i = \beta_j \; \forall i$ against the alternative hypothesis of $\beta_i \neq \beta_j$ for a non-zero fraction of pairwise slopes for $i \neq j$ in equation (2). The test is adequate for any N and T by assuming that the error terms follow a normal distribution. This test can be described as follows:

$$SH = \sum_{i=1}^{N} (\overline{\beta}_i - \overline{\beta}_{WFE})' \times \frac{(\overline{\beta}_i - \overline{\beta}_{WFE}) X'_i M_\tau X_{it}}{\overline{\sigma}_i{}^2}, \tag{6}$$

where $\overline{\beta}_i = X'_i M_\tau y_i (X'_i M_\tau X_i)^{-1}$, M_τ is an identity matrix of order T, and $\overline{\sigma}_i{}^2$ is defined as:

$$\overline{\sigma}_i{}^2 = \frac{(y_{it} - X_{it}\overline{\beta}_i)' M_\tau (y_{it} - X_{it}\overline{\beta}_i)}{(T-k-1)}, \tag{7}$$

where $\overline{\beta}_{WFE}$ is the weighted fixed effect pooled estimator of slope coefficients:

$$\overline{\beta}_{WFE} = \left(\sum_{i=1}^{N} \frac{X'_{it} M_\tau X_{it}}{\overline{\sigma}_i{}^2} \right)^{-1} \times \sum_{i=1}^{N} \frac{X'_{it} M_\tau y_{it}}{\overline{\sigma}_i{}^2}. \tag{8}$$

Therefore, the standard test statistic for slope homogeneity, $\hat{\Delta}$, is calculated as follows:

$$\hat{\Delta} = \sqrt{N} \times \frac{SH - k}{N\sqrt{2k}}. \tag{9}$$

According to Pesaran and Yamagata (2008), the mean- and variance-bias adjusted version of $\hat{\Delta}$, $\hat{\Delta}_{adj}$ can improve the small sample properties of the $\hat{\Delta}$ test:

$$\hat{\Delta}_{adj} = \frac{SH - E(z_{iT})}{N\sqrt{var(z_{iT})}} \times \sqrt{N}, \tag{10}$$

where, following Pesaran and Yamagata's (2008) notation, $E(z_{iT}) = k$, and $var(z_{iT}) = \frac{2k(T-k-1)}{T+1}$.

2.2 Unit root tests

The causality tests implemented in this paper are based on stationary time series. Therefore, we examine the stationarity of the variables though different panel unit root tests. The first one is the Im–Pesaran–Shin (IPS) test of Im et al. (2003). The model can be described as:

$$\Delta y_{i,t} = a_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} c_{ij} \Delta y_{i,t-j} + e_{it}, \tag{11}$$

where $y_{i,t}$ denotes the variable under study, Δ denotes first differences, and p_i is the lag length for the first difference in the augmented Dickey–Fuller (ADF) regression of Said and Dickey (1984), which can vary across countries. The IPS test can be described as:

$$t_{IPS} = \frac{\sqrt{N}\left(\tilde{t} - bar_{NT} - E(\tilde{t}_T)\right)}{var(\tilde{t}_T)} \sim N(0,1), \tag{12}$$

where t_i is the individual ADF statistic for every country of the panel, $\tilde{t}$ is the corresponding average, that is, $\tilde{t} - bar_{NT} = \frac{\sum_{i=1}^{N} \tilde{t}_{iT}}{N}$, $E(\tilde{t}_T)$ denotes the mean of $\tilde{t}_{iT}$, and $var(\tilde{t}_T)$ denotes the variance of $\tilde{t}_{iT}$. By assuming possible heterogeneity between the different countries of the panel, the IPS test overcomes the autocorrelation problems of the Levin et al. (2002) test.

Since the IPS test does not consider possible cross-sectional dependence in the panel, we additionally apply the cross-sectionally augmented Dickey–Fuller (CADF) test, by Pesaran (2007). We refer to this statistic as the cross-sectional IPS (CIPS):

$$t_{CIPS}(N,T) = \frac{\sum_{i=1}^{N} t_i(N,T)}{N}, \tag{13}$$

where $t_i(N,T)$ denotes the cross-sectional ADF statistic for each cross-section unit.

2.3 Panel cointegration test

We apply cointegration tests to verify whether *GDP* and *Z* follow an equilibrium relationship in the long run. Since endogeneity is a critical issue in econometrics, we apply the panel cointegration test of Pedroni (2004), which is robust to regressors endogeneity.

The test of Pedroni (1999; 2004), based on the Engle and Granger (1987) two-step (residual-based) procedure, is also suitable for our purposes because it controls for heterogeneity. It also includes common time dummies that deal with cross-sectional dependence. This test provides seven statistics. Four of them are based on within-dimension tests, which are simultaneously robust to common time factors and heterogeneity across the members of the panel. The rest of the tests are based on the between-dimension approach, derived by the averages of individual autoregressive coefficients of the unit root test on the residuals of every country i.

The seven statistics are residual-based tests. We obtain the residuals from the following regressions:

$$Y_{it} = c_i + \partial_i t + \sum_{m=1}^{M} \beta_{mi} X_{mit} + \varepsilon_{it}, \tag{14}$$

$$\Delta Y_{it} = \sum_{m=1}^{M} \beta_{mi} \Delta X_{mit} + \mu_{it}, \tag{15}$$

$$\hat{\varepsilon}_{it} = \hat{\rho}_i \hat{\varepsilon}_{it-1} + \hat{u}_{it}, \tag{16}$$

$$\hat{\varepsilon}_{it} = \hat{\rho}_i \hat{\varepsilon}_{it-1} + \sum_{k=1}^{K} \hat{\rho}_{ik} \Delta \hat{\varepsilon}_{it-k} + \hat{u}^*_{it}, \tag{17}$$

where M is the number of independent variables (in our study $M = 1$), and $k = 1, 2, \ldots, K$ denotes the lag length in the ADF regression. The parameter ∂_i is the potentially heterogeneous country-specific deterministic trend, while c_i allows for the country fixed effects across countries of the panel, thus leading with the unobserved heterogeneity between the EU countries. The slope coefficients, β_i, can be heterogeneous across countries.

The variables Y_{it} and X_{it} (output and autonomous demand respectively in our study) are assumed to be integrated of order one for each member of the panel, and under the null of no cointegration the residuals will also be I(1) processes (Pedroni 2004, p. 599). The null hypothesis of the test is that there is no cointegration between Y_{it} and X_{mit}, that is, the residuals, ε_{it}, follow an I(1) process. Several series and parameters can be calculated as follows (see Pedroni 1999, p. 661):

$$\hat{s}_i^{*^2} = \frac{1}{T} \sum_{t=1}^{T} \hat{u}_{it}^{*^2}, \quad \tilde{s}_{N,T}^{*^2} = \frac{1}{T} \sum_{n=1}^{N} \hat{s}_i^{*^2}, \tag{18}$$

$$\hat{L}_{11i}^{-2} = \frac{1}{T} \sum_{t=1}^{T} \hat{\mu}_{it}^2 + \frac{2}{T} \sum_{s=1}^{k_i} \left(1 - \frac{s}{k_i + 1}\right) \sum_{t=s+1}^{T} \hat{\mu}_{it} \hat{\mu}_{it-s}, \tag{19}$$

$$\hat{\lambda}_i = \frac{1}{T} \sum_{s=1}^{k_i} \left(1 - \frac{s}{k_i + 1}\right) \sum_{t=s+1}^{T} \hat{u}_{it} \hat{u}_{it-s}, \tag{20}$$

$$\hat{s}_i^2 = \frac{1}{T} \sum\nolimits_{t=1}^{T} \hat{u}_{it}^2, \quad \hat{\sigma}_i^2 = \hat{s}_i^2 + 2\hat{\lambda}_i, \quad \tilde{\sigma}_{N,T}^2 = \frac{1}{N} \sum\nolimits_{n=1}^{N} \hat{L}_{11i}^{-2} \hat{\sigma}_i^2. \tag{21}$$

Pedroni (1999; 2004) proposes various ways of elaborating statistics to test the null hypothesis of no cointegration. The seven statistics are elaborated as follows (see Pedroni 1999, p. 660):

$$\text{Panel } v = T^2 N^{(2/3)} \left(\sum\nolimits_{i=1}^{N} \sum\nolimits_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{\varepsilon}_{it-1}^2 \right)^{-1} \tag{22}$$

$$\text{Panel } \rho = T\sqrt{N} \left(\sum\nolimits_{i=1}^{N} \sum\nolimits_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{\varepsilon}_{it-1}^2 \right)^{-1} \sum\nolimits_{i=1}^{N} \sum\nolimits_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{\varepsilon}_{it-1} \Delta \hat{\varepsilon}_{it} - \hat{\lambda}_i \right) \tag{23}$$

$$\text{Panel } t = \left(\tilde{\sigma}_{N,T}^2 \sum\nolimits_{i=1}^{N} \sum\nolimits_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{\varepsilon}_{it-1}^2 \right)^{-\frac{1}{2}} \sum\nolimits_{i=1}^{N} \sum\nolimits_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{\varepsilon}_{it-1} \Delta \hat{\varepsilon}_{it} - \hat{\lambda}_i \right) \tag{24}$$

$$\text{Panel}\,ADF = \left(\tilde{s}^{*2}_{N,T}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{\varepsilon}_{it-1}^{2}\right)^{-\frac{1}{2}}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{\varepsilon}_{it-1}^{*}\Delta\hat{\varepsilon}_{it}^{*} \tag{25}$$

$$\text{Group}\,\rho = T\frac{1}{\sqrt{N}}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{\varepsilon}_{it-1}^{2}\right)^{-1}\sum_{t=1}^{T}\left(\hat{\varepsilon}_{it-1}\Delta\hat{\varepsilon}_{it}-\hat{\lambda}_{i}\right) \tag{26}$$

$$\text{Group}\,t = \frac{1}{\sqrt{N}}\sum_{i=1}^{N}\left(\hat{\sigma}_{i}^{2}\sum_{t=1}^{T}\hat{\varepsilon}_{it-1}^{2}\right)^{-\frac{1}{2}}\sum_{t=1}^{T}\left(\hat{\varepsilon}_{it-1}\Delta\hat{\varepsilon}_{it}-\hat{\lambda}_{i}\right) \tag{27}$$

$$\text{Group}\,ADF = \frac{1}{\sqrt{N}}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{s}_{i}^{*^{2}}\hat{\varepsilon}_{it-1}^{*^{2}}\right)^{-\frac{1}{2}}\sum_{t=1}^{T}\hat{\varepsilon}_{it-1}\Delta\hat{\varepsilon}_{it}. \tag{28}$$

Then the tests are adjusted so they follow a normal distribution under the null hypothesis of no cointegration (Pedroni 1999, p. 662).

Our bivariate approach to cointegration is analytically valid since if there exists a cointegrating relationship between two non-stationary variables, the same cointegrating relationship also exists in an extended variable space (Johansen 2000). This justifies considering small sub-systems, such as the one in our paper, while the inclusion of additional variables would unnecessarily increase the number of cointegrating equations that would have to be identified and estimated (Herzer and Nunnenkamp 2012, p. 253).

2.4 Long-run elasticities

After testing for cointegration, we estimate the long-run parameters of the cointegration relationship between autonomous demand and output. We implement the panel fully modified ordinary least squares (FMOLS) method by Pedroni (2000). We can interpret the estimated coefficients as long-term elasticities, since we use the variables in natural logarithms.

Unlike the ordinary least squares (OLS) method, the panel FMOLS estimator is effective in dealing with regressors endogeneity and errors serial correlation. The FMOLS corrects these problems through non-parametric regressions. Moreover, this method is also robust to heterogeneity through allowing time trends, different short-run dynamics, and fixed effects. Cross-sectional dependence is also considered by taking into account time-specific effects. The panel FMOLS estimator is defined as:

$$\hat{\beta}_{FMOLS}-\beta = \left(\sum_{i=1}^{N}\hat{L}_{22i}^{-2}\sum_{t=1}^{T}\left(X_{it}-\overline{X}_{it}\right)^{2}\right)^{-1}\sum_{i=1}^{N}\hat{L}_{11i}^{-1}\hat{L}_{22i}^{-1}\left(\sum_{t=1}^{T}(x_{it}-\overline{x}_{it})\varepsilon_{it}^{*}-T\hat{\gamma}_{i}\right), \tag{29}$$

where β denotes the cointegrating vector of Y_{it} in equation (14), X_{it} is the independent variable in equation (14), and $\overline{X}_{it}$ denotes the specific individual mean of X_{it}; and

$$\varepsilon_{it}^{*} = \varepsilon_{it}-\frac{\hat{L}_{21i}}{\hat{L}_{22i}}\Delta X_{it}, \quad \hat{\gamma}_{i} = \hat{\Gamma}_{21i}+\hat{\Omega}_{21i}^{0}-\frac{\hat{L}_{21i}}{\hat{L}_{22i}}\left(\hat{\Gamma}_{22i}+\hat{\Omega}_{22i}^{0}\right),$$

where ε_{it} is the residual of the cointegrating equation, $\widehat{\Omega}_i^0$ denotes the contemporaneous covariance, $\widehat{\Gamma}_i$ denotes a weighted sum of autocovariances, and $\hat{L}_i$ denotes a lower triangular decomposition of $\widehat{\Omega}_i$. The t-statistic is:

$$t_{FMOLS} = (\hat{\beta}_{FMOLS,i} - \beta) \times \left[\sum_{i=1}^{N} \hat{L}_{22i}^{-2} \sum_{t=1}^{T} (Y_{it} - \overline{Y})^2 \right]^{-\frac{1}{2}} \sim N(0,1). \tag{30}$$

2.5 Heterogeneous panel long-run causality test

As we stated at the beginning of this section, we expected that conventional Granger-causality tests suggest the existence of bidirectional causality between *GDP* and Z. As Girardi and Pariboni (2015; 2016; 2019) explain, there are multiple reasons for that. However, our hypothesis is that this feedback causality might be only of a short-run nature, so that the impacts that changes in *GDP* have on changes in Z eventually vanish and do not have a lasting effect on the long-run evolution of Z. As Nikiforos (2018, p. 659) states, 'the Supermultiplier model is a model of the long run and has to be evaluated as such.' Therefore, in what follows, we test whether the evolution of autonomous demand unidirectionally affects long-run economic growth.

As specified by Engle and Granger (1987), if the variables have a unit root and are cointegrated, one can test for a causal relationship between them through a vector error-correction model (VECM). However, since we are working with a group of countries that are highly heterogeneous, we apply the heterogeneous panel causality test by Canning and Pedroni (2008). This approach permits a considerable degree of heterogeneity in the coefficient of the error correction term, in contrast to the traditional VECM-based approach. Denoting X_{it} and Y_{it} as two stationary variables, this test considers a dynamic error correction model for every individual country of the panel, as follows:

$$Y_{it} = \gamma_{1i} + \lambda_{1i}\varepsilon_{i,t-1} + \sum_{j=1}^{p_i^1} \varnothing_{11ij} Y_{i,t-j} + \sum_{j=1}^{p_i^1} \varnothing_{12ij} X_{i,t-j} + u_{1it}, \tag{31}$$

$$X_{it} = \gamma_{2i} + \lambda_{2i}\varepsilon_{i,t-1} + \sum_{j=1}^{p_i^2} \varnothing_{21ij} Y_{i,t-j} + \sum_{j=1}^{p_i^2} \varnothing_{22ij} X_{i,t-j} + u_{2it}, \tag{32}$$

where $\varepsilon_{i,t-1}$ denotes the lagged residual of the cointegrating equation between X and Y of each country i, estimated using FMOLS, and p_i is the specific lag-length for each country i.

Long-run causality is analysed by testing the significance of the adjustment speed, λ_{si}. We test $H_0 : \lambda_{1i} = 0$ in equation (31), and $H_0 : \lambda_{2i} = 0$ in equation (32). Canning and Pedroni (2008) developed two tests to check the null hypothesis of no panel long-run causality: the group mean (GM) test and the Lambda–Pearson (LP) test. Both tests are developed from the analysis of the error correction models specified for each country.

The GM test is constructed by the average of the t-tests of all countries, and follows a standard normal distribution:

$$\bar{t}_{\lambda s} = \frac{\sum_{i=1}^{N} t_{\lambda si}}{N} \sim N(0,1), \tag{33}$$

where $t_{\lambda si}$ is the t-test on λ for each country i, with $s = 1, 2$ in a bivariate setting.

The LP test, on the other hand, focuses on the p-values of the individual t-tests, and follows a chi-square distribution with $2 \times N$ degrees of freedom:

$$P_{\lambda s} = -2 \sum_{j=1}^{N} lnp_{\lambda si}, \tag{34}$$

where $lnp_{\lambda si}$ is the logarithm of the p-value of each t-test statistic corresponding to each country i. The LP test is an interesting complement to the GM test when there is heterogeneity in the parameters λ_{si}. When the GM test cannot reject the null of no causality and the LP test can reject it, Canning and Pedroni (2008) interpret that the average value of λ_{si} is zero, but it is not pervasively zero, that is, it is not zero for all the countries of the panel. It means that there is no panel causality, but that there may exist causality in some countries of the panel.

2.6 The investment adjustment mechanism

Finally, we check the empirical validity of the investment accelerator principle implied by the supermultiplier model to tackle dynamic instability. According to the model presented in the introduction, investment reacts, through the accelerating effect, to changes in the autonomous trend, thus permitting productive capacity to adapt to demand. We should check whether increases (decreases) in the growth rate of *GDP*, induced by increases (decreases) in in the growth rate of autonomous demand, g^Z, exert a sufficiently strong accelerating effect on investment. This accelerating effect must be such that investment increases (decreases) faster than *GDP* for a while, that is, the investment share, h, increases (decreases). For a detailed analysis of the fundamental equations of the model and of the functioning of this mechanism, see Serrano (1995), Freitas and Serrano (2015), and Girardi and Pariboni (2016; 2019).

Since we expect h and g^Z to be I(0) processes, we cannot test the causal relationship between these two variables in an error-correction framework. Therefore, we estimate a regression in the following form:

$$Y_{it} := \mu_{1i} + \varphi_{1t} + \sum_{j=1}^{p} \gamma_{1j} Y_{i,t-j} + \sum_{j=1}^{p} \delta_{1j} X_{i,t-j} + \varepsilon_{1i,t}, \tag{35}$$

$$X_{it} := \mu_{2i} + \varphi_{2t} + \sum_{j=1}^{p} \gamma_{2j} Y_{i,t-j} + \sum_{j=1}^{p} \delta_{2j} X_{i,t-j} + \varepsilon_{2i,t}, \tag{36}$$

where Y_{it} is the investment share (the growth rate of autonomous demand) and X_{it} is the growth rate of autonomous demand (the investment share) of country i in time t, μ_i represents unobserved individual time-invariant heterogeneity, which is considered constant over time, and φ is a (quarterly) time-specific effect that allows for unobserved shocks that are common across the countries.

For robustness, we control for the real interest rate r; thus, we analyse a trivariate model. As proxy variable for the real interest rate, we use quarterly data of the European Economic and Monetary Union (EMU) to converge criterion bond yields for long-term interest rates (central government bond yields on the secondary market, gross of tax, with around ten years' residual maturity). We employ data from Eurostat. Since r is an I(1) process, we use Δr in the model.

Since equations (35) and (36) contain fixed effects and lagged dependent variables, we expect the OLS estimator to be biased and inconsistent due to the so-called Nickell bias (Nickell 1981). The Arellano–Bond system generalized method of moments (GMM) estimator, by Arellano and Bover (1995) and Blundell and Bond (1998), is robust to the Nickell bias. This method uses orthogonal deviations to eliminate the fixed effects, thus preserving the orthogonality between transformed variables and lagged regressors.

However, since our time dimension is relatively large, the system GMM generates a large number of instruments that (in our case) increases linearly with T. Judson and Owen (1999) show that the fixed-effects model is more appropriate to alternative methods when $T \geq 30$, as in our case. Therefore, we run a regression that includes country fixed effects and time fixed effects, that is, the so-called two-way fixed-effects model. Evaluating whether $X_{i,t-j}$ ($Y_{i,t-j}$) predicts Y_{it} (X_{it}) implies testing $\delta_{1j} = 0$ ($\gamma_{2j} = 0$) $\forall j$ on equations (35) and (36).

3 EMPIRICAL RESULTS

In this section, we present the empirical results. Table 1 reports the results of the cross-sectional dependence (CD) tests and the Pesaran and Yamagata (2008) slope homogeneity test. The CD tests confirm the existence of cross-sectional dependence at the 1 percent significance level. The Pesaran and Yamagata (2008) test rejects the null hypothesis of slope homogeneity; therefore, we verify the presence of country-specific heterogeneity at the 1 percent significance level.

Due to the existence of cross-sectional dependence and heterogeneity in the panel, we apply panel unit root tests that consider them. The null hypothesis of the IPS and the CIPS tests is that the series follows a unit root process. The test results indicate that the variables GDP, Z, and r are non-stationary in level and stationary in first differences, while h is stationary in level, at the 1 percent significance level (Tables 2a and 2b). These results allow us to apply Pedroni's approach to panel cointegration to verify whether there is a panel long-run equilibrium relationship between autonomous demand and output.

Next, we check for the existence of a long-run equilibrium relationship (that is, cointegration) between output and autonomous demand though a panel cointegration

Table 1 Tests for cross-sectional dependence and for homogeneity

Test	Statistic	*P*-value
Cross-sectional independence		
Breusch–Pagan LM	4004.13***	0.00
Pesaran scaled LM	250.71***	0.00
Pesaran CD	14.43***	0.00
Pesaran and Yamagata slope homogeneity		
$\widehat{\Delta}$	4327.41***	0.00
$\widehat{\Delta}_{adj}$	10.23***	0.00

Notes: This table presents the results of the Breusch–Pagan LM, the Pesaran scaled LM, and the Pesaran CD tests of cross-sectional independence, and the Pesaran and Yamagata slope homogeneity tests of equations (3), (4), (5), (9), and (10), respectively. The asterisks *** denote rejection of the null hypothesis of cross-sectional independence, as well as the rejection of the null hypothesis of slope homogeneity at the 1% significance level.

Table 2a Panel unit root tests: exogenous variables – individual effects

Tests	Variables							
	GDP	Δ*GDP*	*h*	Δ*h*	*Z*	Δ*Z*	*r*	Δ*r*
t_{IPS}	−1.12	−18.17***	−7.39***	−34.27***	−0.2	−28.89***	7.06	−15.30***
t_{CIPS}	−1.05	−16.18***	−4.92***	−17.15***	−1.82	−4.78***	4.11	−11.70***

Notes: This table presents the results of the IPS and CIPS panel unit root tests proposed by Im et al. (2003) and Pesaran (2007) of equations (12) and (13), respectively, considering individual effects. The asterisks *** denote rejection of the null hypothesis of a unit root process at the 1% significance level. The long-run variance of the residual is computed by applying the Newey and West (1987) heteroskedasticity and autocorrelation-consistent method with Barlett kernel. The lag order for the Barlett kernel is selected by the Schwartz information criterion (SIC), by Schwarz (1978).

Table 2b Panel unit root tests: exogenous variables – individual effects and individual linear trends

Tests	Variables							
	GDP	Δ*GDP*	*h*	Δ*h*	*Z*	Δ*Z*	*r*	Δ*r*
t_{IPS}	−1.12	−18.17***	−7.39***	−34.27***	−0.20	−28.89***	5.46	−15.59***
t_{CIPS}	−0.24	−14.16***	−5.44***	−15.29***	0.12	−10.94***	2.62	−11.26***

Notes: This table presents the results of the IPS and CIPS panel unit root tests proposed by Im et al. (2003) and Pesaran (2007) of equations (12) and (13), respectively, considering individual effects and individual linear trends. The asterisks *** denote rejection of the null hypothesis of a unit root process at the 1% significance level. The long-run variance of the residual is computed by applying the Newey and West (1987) heteroskedasticity and autocorrelation-consistent method with Barlett kernel. The lag order for the Barlett kernel is selected by the SIC.

analysis approach. Table 3 shows the results of the Pedroni (2004) test. The seven Pedroni statistics reject the null hypothesis of no cointegration at the 5 percent significance level. Therefore, we find cointegration between output and autonomous demand.

Interestingly, we find no evidence of cointegration between output and each component of autonomous demand (which are also I(1) processes) separately. Therefore, it becomes necessary to aggregate the components of autonomous demand in a unique variable (Z) to find a long-run equilibrium relationship between output and autonomous demand. Thus, as Pariboni (2016) highlights, it is the average growth rate of autonomous expenditures that in the long run leads the growth rate of output (in the case that we find unidirectional long-term causality running from autonomous demand to output).

Once the panel cointegrating relationship between output and autonomous demand has been acknowledged, we estimate the panel long-term parameters of this relationship. Table 4 shows the estimated elasticity by the FMOLS method of equation (29). The long-term coefficient is statistically significant, positive, and very close to 1. It suggests that in the long run autonomous demand and output move in step.

After verifying the existence of cointegration between autonomous demand and output, we test for a causal relationship between them. The hypotheses of the supermultiplier model suggest that changes in autonomous demand determine changes in output in the long run. The long-run causal relationship between the variables is analysed through the Canning and Pedroni (2008) heterogeneous panel causality test.[2]

2. We thank Peter Pedroni for providing us their RATS codes.

Table 3 Pedroni (2004) heterogeneous panel cointegration test; dependent variable: GDP

	Independent variable							
	Z		*G*		*E*		*AC*	
	Weighted statistic	*P*-value	Weighted statistic	*P*-value	Weighted statistic	*P*-value	Weighted statistic	*P*-value
Panel *v*-statistic	1.83**	0.03	2.79***	0.00	−1.25	0.89	6.58***	0.00
Panel ρ-statistic	−2.81***	0.00	−0.03	0.48	0.53	0.7	0.44	0.67
Panel *t*-statistic	−2.99***	0.00	−0.57	0.28	−0.18	0.42	−0.12	0.45
Panel ADF-statistic	−2.57***	0.00	−0.68	0.24	−0.05	0.47	0.29	0.61
	Statistic	*P*-value	Statistic	*P*-value	Statistic	*P*-value	Statistic	*P*-value
Group ρ-statistic	−4.1***	0.00	−0.03	0.48	−1.23	0.11	1.93	0.97
Group *t*-statistic	−3.51***	0.00	−0.42	0.33	−1.22	0.11	1.25	0.89
Group ADF-statistic	−3.51***	0.00	−1.14	0.13	−1.13	0.12	1.52	0.93

Notes: This table presents the results of Pedroni's (2004) panel cointegration test between output and autonomous demand described in equations (22)–(28). It also presents the cointegration test between output and each component of autonomous demand separately. Deterministic trend and intercept are assumed. The lag length selection is based on SIC with a maximum lag of 11 (Newey–West automatic bandwidth selection and Bartlett kernel). The asterisks *** and ** indicate rejection of the null hypothesis of no cointegration at the 1% and 5% significance level, respectively.

Table 4 Panel long-run elasticity estimation results; dependent variable: GDP

Regressor	FMOLS	
	Coefficient	*t*-statistic
Z	0.98	134.97***

Notes: This table shows the results of Pedroni's (2000) panel long-run elasticity estimator of autonomous demand on output. We considered a deterministic constant and trend in the cointegrating equation. Automatic leads and lags specification are based on SIC criterion (Bartlett kernel, Newey–West fixed bandwidth). The asterisks *** indicate statistical significance at the 1% significance level.

Table 5 shows the results of the Canning and Pedroni (2008) tests. The LP test suggests that λ_1 is statistically significant at the 1 percent significance level, while the GM test indicates that λ_{1i} is statistically significant at the 10 percent significance level. It means that the causality running from autonomous demand to output can be considered homogeneous at the 10 percent significance level. On the other hand, in the analysis of the causality running from output to autonomous demand, the LP shows that λ_{2i} is statistically significant at the 1 percent significance level, while the GM test indicates that λ_{2i} is not statistically significant at any level. Therefore, the average value of λ_{2i} is zero, but it is not pervasively zero, that is, it is not zero for all the countries that compose the panel. Then, there exists long-run causality running from output to autonomous demand for some countries of the panel, but not for all of them.

According to Canning and Pedroni (2008), our results indicate the existence of panel causality running unidirectionally from autonomous demand to output at the 10 percent significance level. Liddle and Lung (2015) and Mahalingam and Orman (2018), among

Table 5 Canning and Pedroni (2008) heterogeneous panel long-run causality test

	1996Q1–2017Q4				1996Q1–2007Q4				2008Q1–2017Q4			
	GM		LP		GM		LP		GM		LP	
	Test	*P*-value	Test	*P*-value	Test	*P*-value	Test	*P*-value	Test	*P*-value	Test	*P*-value
$Z \rightarrow GDP$	−1.39*	0.08	78.48***	0.00	−0.80	0.21	56.64***	0.00	−1.20*	0.09	85.32***	0.00
$GDP \rightarrow Z$	0.30	0.62	56.73***	0.00	1.34	0.91	93.01***	0.00	0.48	0.68	57.64***	0.00

Notes: This table shows the results of the long-run heterogeneous panel causality test of Canning and Pedroni (2008). We report the GM and LP statistics of equations (33) and (34), respectively, showing the estimated t-test for the null hypothesis that $\lambda_i = 0$ and the corresponding *p*-value. The symbol → denotes the direction of causality. The asterisks *** and * denote rejection of the null hypothesis of no causality at the 1% and 10% significance level, respectively.

others, only accept a long-run causality for the panel when both the LP test and the GM test reject the null hypothesis of no causality (Liddle and Lung 2015, p. 50). Following these authors, we conclude that there is long-run panel causality running exclusively from Z to *GDP*. However, we found it was also interesting to highlight the existence of long-run causality running from *GDP* to Z for some countries of the panel. These results validate that the long-run evolution of autonomous demand is indeed autonomous (with respect to the evolution of output) in a heterogeneous panel framework.

Since autonomous demand is a generated variable obtained from aggregation of underlying variables, we must check whether some element of autonomous demand is influenced by output. We cannot check long-run causality between output and each component of autonomous demand separately, because there is no cointegration between them. Therefore, following Girardi and Pariboni (2016), we remove each element of Z and analyse how the model performs without it. When we remove G from autonomous demand $(Z = E + AC)$, we find no cointegration between output and autonomous demand. The same happens when we remove $E(Z = G + AC)$. When we remove $AC(Z = G + E)$, we find panel cointegration between output and autonomous demand. However, the panel long-run causality between output and autonomous demand remains the same as when we use $Z = G + E + AC$, while we obtain an FMOLS estimator smaller than one. Therefore, the inclusion of AC in Z is appropriate for a bivariate analysis between autonomous demand and output. This result is in line with Girardi and Pariboni (2019), who show that the effect of autonomous demand on the investment share is not driven by one single component, but is a property of autonomous demand as a whole (Girardi and Pariboni 2019, p. 8).

As a robustness check, we perform our analysis on the sub-sample periods 1996–2008 and 2008–2018. We choose these sub-samples to isolate the effects of the Great Recession. For the sub-sample 2008–2018 the results coincide with those of the full period: there is homogeneous causality running from Z to *GDP* at the 10 percent significance level, while the causality running from *GDP* to Z is heterogeneous. However, for the first sub-sample, we find bidirectional causality between *GDP* and Z, although this causality is heterogeneous (therefore, no panel causality). Nevertheless, the orthogonalized impulse response functions show that the effects of Z on *GDP* are stronger and more permanent than the effects of *GDP* on Z. Thus, the existence of bidirectional heterogeneous causality in this sub-period may be explained by the short span of data.

Finally, we report the results of the analysis of the investment-accelerator mechanism. To test for panel Granger-causality between h and g^Z, we have analysed nine different two-way fixed effects models with a different number of lags of g^Z and h as regressors (Table 6). We use a maximum of three lags of g^Z and h as regressors because all subsequent lags are statistically insignificant. In every model, we include three lags of the variable Δr; however, the results do not change if we modify the number of lags of this variable.

For the models in which h_t is the dependent variable, Table 6 shows that for the nine models, the coefficients of g^Z_{t-j} are positive and significant, and the joint significance test suggests evidence of causality running from g^Z to h at the 1 percent significance level. On the other hand, when we analyse the models in which g^Z_t is the dependent variable, we see that for six of the nine models the Wald test indicates no causality running from h to g^Z. But in the three models where we reject the null of no causality, only the coefficient associated with h_{t-3} is statistically significant, and it has a low and negative value. Overall, these results indicate that there is enough empirical evidence of same-sign causality running unidirectionally from g^Z to h.

Table 6 Panel causality tests between h and g^Z

Dependent variable: h_{it}					Lags of g^Z_{it} and h_{it}				
Coefficients	(1/1)	(2/1)	(3/1)	(1/2)	(2/2)	(3/2)	(1/3)	(2/3)	(3/3)
$g^Z_{i,t-1}$	0.30***	0.29***	0.26***	0.305***	0.29***	0.26***	0.32***	0.30***	0.27***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$g^Z_{i,t-2}$	–	0.13	0.12	–	0.25***	0.24***	–	0.26***	0.25***
		(0.14)	(0.17)		(0.00)	(0.00)		(0.00)	(0.00)
$g^Z_{i,t-3}$	–	–	0.18**	–	–	0.19**	–	–	0.22***
			(0.04)			(0.03)			(0.00)
F-statistic (H_0: g^Z does not cause h)	11.49***	6.80***	5.92***	13.36***	10.99***	8.95***	14.91***	12.32***	10.60***
p-value of the F-statistic	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Dependent variable: g^Z_{it}					Lags of h_{it} and g^Z_{it}				
Coefficients	(1/1)	(2/1)	(3/1)	(1/2)	(2/2)	(3/2)	(1/3)	(2/3)	(3/3)
$h_{i,t-1}$	–0.00	0.00	0.01	–0.00	0.00	0.00	–0.00	–0.00	0.00
	(0.33)	(0.59)	(0.11)	(0.24)	(0.98)	(0.29)	(0.21)	(0.95)	(0.36)
$h_{i,t-2}$	–	–0.01*	–0.00	–	0.00	0.00	–	–0.00	0.00
		(0.08)	(0.84)		(0.23)	(0.86)		(0.23)	(0.92)
$h_{i,t-3}$	–	–	–0.02***	–	–	–0.02***	–	–	–0.02**
			(0.00)			(0.00)			(0.01)
F-statistic (H_0: h does not cause g^Z)	0.93	1.97	4.33***	1.13	1.38	3.57**	1.59	1.50	3.28**
p-value of the F-statistic	0.33	0.14	0.00	0.24	0.25	0.01	0.21	0.22	0.02

Notes: This table shows the results of the panel causality tests between h and g^Z described in equations (35) and (36). The results are based on the two-way fixed effects regression. All regressions include three lags of the variable Δr and different lags (from one to three) of the dependent and the independent variables. We report the coefficients and the correspondent p-values (in parentheses) associated with the independent variable in the nine different models. The asterisks ***, **, and * denote statistical significance at the 1%, 5%, and 10% level, respectively. Rejection of the null hypothesis of the F-statistic implies the existence of Granger-causality running from $h\,(g^Z)$ to $g^Z(h)$.

As a robustness check, we have also analysed separately the sub-samples 1995–2008 and 2008–2018. In both sub-periods we also find statistical evidence of unidirectional same-sign causality going from g^Z to h. As an additional robustness check, we estimate the model through the system GMM. System-GMM estimators are similar in magnitude and statistical significance to those of the two-way fixed effects model. We have not included these results for reasons of space, but they are available upon request. Our results are in line with the findings of Girardi and Pariboni (2019), who find evidence of a causal effect of the growth rate of autonomous demand on the investment share for a panel of 16 OECD countries for the 1960–2015 period.

These results empirically verify the specification of the investment-accelerator mechanism implied by the supermultiplier. They are in line with the postulates of the Sraffian supermultiplier model and its ability to provide a stabilizing mechanism that allows the convergence of the effective degree of capacity utilization to the normal one.

Since g^Z and h are not cointegrated variables, if we measured the effect of g^Z on h, the obtained estimator may be biased, due to the existence of unobserved time-varying factors that simultaneously affect g^Z and h. To address this concern, it is necessary to employ an instrumental variables strategy. By doing so, one can identify an exogenous variation of autonomous demand and its correspondent effect on h. However, finding a valid (exogenous) instrument matrix for autonomous demand is not an easy task. This is, in fact, the main objective of the work in progress of Girardi and Pariboni (2019).

4 CONCLUDING REMARKS

This paper investigates the empirical validity of the Sraffian supermultiplier approach to growth and distribution. At present, there seems to be the possibility of reaching a consensus among some Neo-Kaleckian and Sraffian economists. This consensus revolves around: (i) the consideration of non-capacity-generating autonomous expenditures as drivers of long-run economic growth; (ii) an investment function based on the (flexible) accelerator principle; and (iii) the tendency towards a normal utilization rate of capacity. Some Neo-Kaleckian economists have accepted the importance of non-capacity-generating autonomous expenditures and the tendency towards a normal capacity utilization rate; whereas some Sraffians have stated that the discrepancies between the effective rate of capacity utilization, u, and the normal one, u_n, are the factor enabling the functioning of the mechanism capable of bringing u back to u_n (the investment-accelerator mechanism).

The main result of this potential consensus model, whose main economic intuitions date back to Serrano (1995), can be summarized in the following statement: the long-run evolution of aggregate demand (which, in turn, determines the evolution of the rate of capital accumulation) is determined by the growth rate of autonomous expenditures, the genuine engine of economic growth in the long run. In this paper, we verify the empirical validity of this statement for certain European economies during the period 1995–2018.

Our results are in line with Médici (2011) for the Argentinian economy and with Girardi and Pariboni (2016) for the US economy. However, these articles refer to a single economy and they use time-series techniques. We analyse 16 heterogeneous and highly interdependent European economies, and use a heterogeneous panel data methodology that is generally more robust than the time-series one, especially when the number of time observations is small. Besides, we believe it is necessary to complement causal connectivity approaches (like the one implemented in this paper) with effective connectivity approaches, such as maximum-likelihood or Bayesian estimation methods (see Schoder 2017 for the utilization of these methods to estimate Keynesian models). This will be the purpose of our future investigation.

Economic growth is a complex phenomenon. As growth models treat more variables as endogenous, we face the paradox that if everything is endogenous there is no independent variable with explanatory power. In the framework of demand-led models, the Sraffian supermultiplier approach isolates some demand components that can be reasonably assumed to be autonomous with respect to the evolution of income. By doing so, one can identify exogenous and spatial factors in which to situate the debate about growth. That has similarities with Keynes's (1936) *General Theory of Employment, Interest and Money* which established causal macroeconomic relationships by identifying demand variables that are sufficiently independent and persistent to determine the rest of the system.

ACKNOWLEDGEMENTS

We are grateful to three anonymous reviewers and to the three editors of *ROKE* for valuable comments and suggestions that helped to improve the paper. We are also in debt to Victor Troster, Riccardo Pariboni, Daniele Girardi, Óscar Dejuán, Ferran Portella-Carbó, Soren Johansen, Tomás Del Barrio, and Ramiro Álvarez for helpful comments and suggestions. Usual disclaimers apply.

FUNDING

This work was supported by the research grant ¿Mediterranean Capitalism?: successes and failures of industrial development in Spain, 1720-2020, PGC2018 093896-B-100, directed by Jordi Catalan and financed by the Spanish Ministry of Economy; and by the research grant FPI/1866/2016, financed by the European Social Fund and the Ministry of Research, Innovation and Tourism of the Balearic Islands, which is gratefully acknowledged.

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