Residential Investment as a Luxemburg-Kalecki External Market in the United States: An Empirical Investigation

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Abstract: We study the residential investment-economic activity nexus in the United States during the period 1960–2020. We employ an asymmetric Granger causality test based on the frequency domain. The combination of the asymmetric and the spectral analysis is robust to structural changes in the relationship among the variables and allows for both positive and negative shocks over different horizons. We find evidence of symmetric and asymmetric frequency-domain Granger causality running unidirectionally from residential investment (RES) to output. This unidirectional causality relationship is both permanent and transitory: transitory shocks in RES have transitory effects on GDP, while permanent shocks in RES have permanent effects on GDP. Our results validate the hypothesis of Fiebiger (2018) and Fiebiger and Lavoie (2019), who state that housing investment in the US can be analogous to a Luxemburg-Kalecki external market. Thus, we highlight the role of residential investment as a driver of both the business cycle and the secular growth in the US.

Keywords: Effective demand; Growth; Residential investment; Autonomous demand; US economy

JEL Codes: B51; E11; E12; O41; O47; O51

1. Introduction

In the title of his famous article, Leamer (2007) stated that “Housing IS the Business Cycle.” He documented that the dynamics of residential investment (RES from now on) has played and is still playing a central role in explaining the business cycles in the United States during the post-WWII era. Leamer also suggested that, with very few exceptions, recessions can be predicted by analyzing the dynamics of RES. Thus, a reduction of the RES/GDP ratio would always precede a recession, while the growth of the RES/GDP ratio would always precede a recovery in the United States.

In recent years, and especially in the aftermath of the global economic crisis of 2008, several research projects have studied the influence of RES on macroeconomic performance. Most of these works are mainly empirical and are consistent with the neoclassical approach to the determination of prices and output. However, some recent works have studied the interaction between RES and growth from a demand-led, i.e., Keynesian, perspective. We argue that in this framework, those models that focus on the role played by the autonomous components of demand (more on this later) in driving growth—such as the Sraffian supermultiplier (SSM) developed by Serrano (1995), Bortis (1997), and Dejuan (2005), and some recent versions[1] of the neo-Kaleckian model—are a promising benchmark for studying the relationship between RES, economic activity, and growth. In this article, we argue, in particular, that with these analytical tools one can go a step beyond Leamer (2007, 2015) and establish that RES not only drives US cycles, but contributes to the long-run path of the economy.

If, following an endogenous credit money approach—as in Palley (1997), Fontana and Setterfield (2009), Cesaratto (2016), and Deleidi and Fontana (2020), among others—one considers that RES is mostly funded through external sources of finance, such as the extension of credit through the creation of new money, it follows that RES's course is independent from the pace of output and of capital accumulation. Thus, the banking system creates *ex nihilo* purchasing power that requires no abstinence on the part of other agents (Fiebiger and Lavoie, 2019: 256). This allows us to establish a link with the Luxemburg-Kalecki external market (Luxemburg 1913; Kalecki, 1968, 1971) intuition and the literature stemming from it, such as Fiebiger (2018) and Fiebiger and Lavoie (2019). In this sense, the exercise we propose in this article can be considered an attempt to provide an empirical validation of the theses of these contributions.

The main aim of this work is to show that RES dynamics can be considered a driver of both cycles and long-run growth in the United States, as predicted by theoretical models such as the SSM and some recent contributions in the neo-Kaleckian tradition. In

[1] See for example Allain (2015, 2019), Lavoie (2016) and Fiebiger and Lavoie (2019).

terms of a time-series empirical analysis, this assertion suggests the existence of both short- and long-run Granger causality running unidirectionally from RES to output. We investigate whether these causal relationships exist. In Section 2, we review the works that test the relationship between RES and economic activity. In Section 3, we briefly introduce a class of models that stress the role of autonomous demand as a driver of economic growth, which will provide the analytical interpretative tool for our empirical findings. In Section 4, we explain the econometrics methodology used in this research. We also explain its benefits: (i) it allows testing symmetric and asymmetric relationships between the variables; (ii) it allows testing permanent and transitory causality relationships; and (iii) it is robust to structural changes and parameter instability in the relationships among the variables. Finally, Section 5 presents our empirical results, while Section 6 concludes the paper.

2. Empirical Literature on the Housing-Economic Activity Nexus

Several works have analyzed whether the strong affirmations of Leamer (2007) ("Housing IS the Business Cycle") and Leamer (2015) ("Housing Really Is the Business Cycle") find empirical support. Some papers seem to empirically validate Leamer's assertions, while others seem to refute them. However, most of the literature does not focus directly on RES; instead, it uses proxies of housing activity, such as housing starts, housing permits, and housing prices. In general, these works consider housing supply indicators, and analyze how they affect macroeconomic variables. To avoid spurious causality between housing and economic activity, they control for supply factors (user costs, borrowing constraints, etc.) and other elements generally related to monetary shocks (interest rate).

Goodhart and Hofmann (2008) find evidence of a multidirectional link between house prices, monetary variables, and the macroeconomic activity for 17 industrialized countries. Ghent and Owyang (2010) assess the relationship between housing variables and employment over the business cycle in a set of 51 US cities, finding no evidence that housing permits or housing prices influence the business cycle. Musso et al. (2011) state that housing demand has a strong role in transmitting monetary policy shocks in the US. Similarly, Cesa-Bianchi (2013) suggests that house price shocks are quickly transmitted to the domestic real economy, leading to a short-run expansion of real GDP and consumer prices in the world economy. Jordà et al. (2016) show that contemporary business cycles are increasingly shaped by the dynamics of mortgage credit for 17 advanced economies since 1870, which supports the growing importance of housing and housing finance for the overall economy. Huang et al. (2018) point out that housing factors, such as housing starts and housing prices, autonomously drive the business cycle in the US. Similarly, Ren and Yuan (2014) focus on the leading role of housing indicators on macroeconomic activity for the US.

On the other hand, research on the macroeconomic implications of residential investment in terms of volume—the housing output produced for the market—is much scarcer (Kohlscheen et al., 2018: 2). Arestis and González-Martínez (2015) indicate that residential investment is influenced by real disposable income in the US and in other OECD countries. However, Kydland et al. (2016) state that RES leads output, whereas nonresidential investment lags output in the US; nevertheless, these authors emphasize that RES by itself is not capable of driving business cycles, since it is determined by changes in nominal interest rates. Thus, housing variables are simply channels through which monetary policy propagates. On the other hand, Kohlscheen et al. (2018) find that information on residential investment growth consistently improves the performance of the yield curve as a predictor of recessions across nine OECD countries. Similarly, Arestis and Zhang (2020) suggest that in China, the real estate sector is the main field of fixed investment and the key driver of the nation's GDP growth.

The second strand of works—those that investigate the housing-economic growth nexus in terms of volume—is more relevant to our research, since, as we will explain in Section 3, we consider RES a potentially autonomous component of final aggregate demand. There are some exceptions but most of these works are generally informed by the neoclassical approach to growth and distribution, and, thus, consider economic growth to be supply-led. Some of these works consider RES an exogenous component of demand; but, since growth is supply-led in their research framework, the dynamics of RES can only explain the cycle, not the secular growth rate. Thus, with the exception of Huang et al. (2018), they only study whether RES can explain business cycles.

We consider the housing-growth nexus under a different theoretical framework, upon which we base our empirical analysis. Inspired by Fiebiger (2018), Fiebiger and Lavoie (2019), and Petrini and Teixeira (2020)—and by the autonomous demand-led growth literature in general—we will investigate the role of residential investment not only in business cycles in the US, but also as a driver of long-run growth, in a context where the borrowing capacity of households has been relatively autonomous with respect to income in the long run. Therefore, we do not narrow our attention to the short run (as most of the vast literature we briefly summarized does) but consider the long-run nexus between the dynamics of RES and output growth in the US.

3. Autonomous Demand-Led Growth Models as a Tool for Studying the Nexus between RES and GDP Growth

In recent years, (broadly speaking) Keynesian growth theory has yielded a class of models that point to the role of autonomous demand in driving economic growth. Since we believe that these analytical constructs are a promising tool for studying the RES-economic activity nexus, we provide a brief overview.[2] The starting point can be found in the Sraffian supermultiplier (SSM), inaugurated by Serrano (1995). In his attempt to develop an alternative theory of production and long-term accumulation, Serrano builds upon two main staples: (i) the validity of the Keynesian-Kaleckian principle of effective demand in the long run; and (ii) the compatibility of this approach to the determination of quantities with the Classical Surplus approach to the theory of value and distribution, revived by Sraffa (1960) (Serrano, 1995: 67), which implies considering income distribution an exogenous variable and, therefore, rejecting any mechanical link between distribution and economic growth.

The main message of the SSM is that output growth tends to be driven by the growth of the non-capacity-generating autonomous components of demand. According to Serrano (1995) and the following literature, autonomous demand comprises and 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: 71). In the long run, effective demand determines normal productive capacity, while autonomous demand generates induced consumption (through the multiplier) and induced (capacity-creating) investment (Serrano, 1995: 67). Indeed, in this model, firms continuously and gradually attempt to reach the normal (cost-minimizing) rate of capacity utilization, so the system is in equilibrium when capacity is utilized normally. This mechanism operates through a flexible accelerator mechanism, in

[2] The interested reader can refer to Freitas and Serrano (2015) and Lavoie (2016) for in-depth treatments of the issue.

which investment is treated as induced by the dynamics of output. Very similar conclusions can be drawn from more recent amended versions of the neo-Kaleckian model, such as Allain (2015) and Lavoie (2016).[3]

The expenditures that should be considered as non-capacity-creating autonomous demand, according to the SSM approach, include: "the consumption of capitalists; the discretionary consumption of richer workers that have some accumulated wealth and access to credit; residential "investment" by households; firms discretionary expenditures that do not include the purchase of produced means of production such as consultancy services, research & development, publicity, executive jets, etc.; government expenditure (both consumption and investment); and total exports (both of consumption and of capital goods since the latter do not create capacity within the domestic economy)" (Serrano, 1995: 71).

Since autonomous demand represents financial dis-saving and is financed to a large extent through debt, the SSM approach ties in with the endogenous money approach and the credit-creating powers of banks (Cesaratto, 2016, 2017). Thus, as highlighted by Cesaratto (2015) and Fiebiger (2018), debt-financed expenditures, while internal to the dynamics of a private capitalist economy, are analogous to an *external market*, as defined by Rosa Luxemburg and Michał Kalecki (Fiebiger and Lavoie, 2019: 254).

If one considers that RES evolves largely independently from output and, at the same time, it is not destined to increase productive capacity (Freitas and Serrano, 2015: 261), then RES can be treated as a non-capacity-creating autonomous component of demand that drives long-run growth. Although the SSM approach explicitly considers residential investment as potential driver of growth, according to Petrini & Teixeira (2020), this feature has not yet been incorporated into a demand-led growth model. The aim of our research is to fill this gap. In this sense, our work complements other attempts such as Nah and Lavoie (2017), Deleidi and Mazzucato (2019), Deleidi and Mazzucato (2019) and Deleidi et al. (2020), where the authors single out and study the contribution of specific components of autonomous demand.[4]

[3] See Palley (2019) and Fazzari, et al. (2020), which introduce into this framework an explicit consideration of labor market dynamics. See also Hein (2018) for a model depicting an economy whose growth is driven by the evolution of autonomous government expenditures, but without convergence to normal utilization.
[4] The first article of the list deals with exports, the other three with government innovation expenditures.

With this in mind, we analyze the short- and long-run relationship between RES and economic growth in the US. In this way, we single out a specific component of autonomous demand that plays a prominent role in the US capitalism. To consistently analyze the RES- GDP nexus with the "autonomous demand-led growth" approach, we consider a multivariate setting and control for the variables that, according to the related literature, constitute other potential components of autonomous demand, such as public expenditure, G; exports, E; and research and development expenditures, RD (see Médici, 2011; Girardi and Pariboni, 2016; and Perez-Montiel & Manera, 2020a, b; and Deleidi and Mazzucato, 2019).

A final, important comment is in order. Pedagogically, the class of models we briefly referred to in this section provides an accessible and coherent story on how expenditures originating in the household, government, or foreign sectors can stabilize as well as drive growth processes. However, the canonical versions of these models do not consider explicit financial relations and constraints; thus, few would defend their strict empirical realism.[5] The consideration of debt-financed expenditure is undoubtedly complex, since it constitutes a non-wage and non-profit source of effective demand that carries numerous complexities from debt servicing (Setterfield and Kim, 2016). Hence, we do not try to provide a complete and thorough validation of autonomous demand-led growth models here. We do not deal with the very important issue of how residential investment is financed either, even though we recognize that the investigation of financial dynamics is the necessary complement of non-capacity creating autonomous demand-led growth models, as recently and persuasively claimed in Hein and Woodgate (2020). The aim of this paper is more modest: we investigate only whether RES can empirically be considered an autonomous driver of final demand and growth, both in the short and in the long run.

4. Data and Methodology

In the next sections, we empirically analyze the relationship between the dynamics of RES and economic growth within a SSM-inspired approach. To this end, we apply the nonlinear auto-regressive distributive lag model (NARDL) of Shin et al. (2014), together with the Granger causality in the frequency domain test of Breitung and Candelon (2006).

[5] See Skott, 2017a, b for a critique of the empirical validity of these models, together with Lavoie's (2017) reply.

As we explain in Section 4.4, we adopt this methodology in order to overcome several weaknesses of standard Granger causality tests, so that the results of our analysis, given the asymmetric framework we use, are not affected by parameter instability or undetected structural breaks. Moreover, the frequency-domain causality approach also allows us to discriminate between short- and long-run effects between our variables of interest.

4.1. Data

We employ quarterly series between 1960:Q1 and 2019:Q2. For the proxy variable for economic activity, we use the evolution of the gross domestic product, GDP. For the proxy variable for residential investment, we use fixed residential investment, RES. Additionally, we consider other potential components of autonomous demand, such as public expenditure, G; exports, E; and research and development expenditures, RD. For G, we use the variable real government consumption expenditures and gross investment. To represent E, we employ the variable real exports of goods and services. As a proxy for RD, we use the variable research and development expenditures. We work with the variables in real terms (measured in billions of chained 2012 dollars, seasonally adjusted annual rate) and transform them into logarithms for consistent and reliable empirical results.

4.2. Unit Root Test

To apply the asymmetric-spectral Granger causality test employed in this research, we first need to know the degree of integration of the variables. Thus, we first check for the stationarity of the variables GDP, RES, G, E, and RD. A non-stationary series that needs to be differenced d times to become stationary is integrated of order d, i.e., is a $\mathrm{I}(d)$ process. We apply the Breakpoint Unit Root test of Perron (1997), which is robust to an unknown structural change. The test considers various methods of selecting the break points and the asymptotic and finite sample distributions of the corresponding statistics (Perron, 1997: 355). The methodology of the Perron (1997) unit root test with breakpoint is widely known; therefore, for reasons of space, we do not describe it here.

4.3. The Asymmetric Relationship between Output and Residential Investment

In this section, we consider the relationship between GDP and RES in a nonlinear framework. Since the literature has widely recognized that macroeconomic variables and processes have nonlinear structures, the information obtained from linear models might not be enough to make reliable predictions. Keynes (1936: 314) himself stated that "the substitution of a downward for an upward tendency often takes place suddenly and violently, whereas there is, as a rule, no such sharp turning point when an upward is substituted for a downward tendency." Shin et al. (2014) warn that the assumption of linear adjustments may be too restrictive in many economically interesting situations, especially where transaction costs are important and where policy interventions are observed in-sample. These authors note that the fields of behavioral finance and economics (Kahneman and Tversky, 1979; Bliss and Shiller, 1995; and Cooper and Shiller, 2000, among others) have stressed that nonlinearity is endemic in the social sciences and that asymmetry is fundamental to the human condition.

In this section, we employ the nonlinear auto-regressive distributive lag model (NARDL) of Shin et al. (2014) to investigate the presence of a relationship between GDP and RES in the US. The nonlinear ARDL model of Shin et al. (2014) is preferred to other nonlinear methods, since this model allows us to simultaneously consider both short- and long-run asymmetries. Indeed, the Markov-switching vector error correction model (MS-VECM), developed by Hamilton (1989) and continued by Krolzig (1997) and Psaradakis et al. (2004), and the smooth transition regression error correction model, developed by Kapetanios et al. (2006), do not allow detection of asymmetry in the long run. On the other hand, the threshold framework, developed by Enders and Siklos (2001), accounts only for long-run asymmetry.

The NARDL, developed by Shin et al. (2014), has been discussed in recent works such as Alsamara et al. (2017), Ivanovski and Churchill (2019), Al-mulali et al. (2019), and Bahmani-Oskooee and Gelan (2019), among others; thus, we only display the basic details of the method. As stated, Shin et al. (2014) propose a nonlinear ARDL model that considers asymmetries, developing a flexible dynamic parametric framework to model relationships that exhibit combined long- and short-run asymmetries. This implies that the explicative variable, x, is decomposed into its positive and negative partial sums:

$$x_t = x_0 + x_t^+ + x_t^-, \tag{1}$$

where $x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} \max\left(\Delta x_j, 0\right)$ and $x_t^- = \sum_{j=1}^{t} \Delta x_j^- = \sum_{j=1}^{t} \min\left(\Delta x_j, 0\right)$ are the partial sum processes that accumulate positive and negative changes in x_t (Granger and Yoon, 2002; Shahzad et al., 2017). Thus, the model considers the following asymmetric long-run regression:

$$y_t = \beta^+ x_t^+ + \beta^- x_t^- + u_t, \tag{2}$$

where β^+ and β^- are the asymmetric long-run parameters associated with positive and negative changes in x_t. If $\beta^+ = \beta^-$, there is a symmetric effect of x on y. If $\beta^+ > 0$ and $\beta^- > 0$, this implies that an increase in x increases y, and a decrease in x decreases y. This means that there are two possible scenarios: either $\beta^+ > \beta^-$ or $\beta^- > \beta^+$. The former suggests that there is a positive effect on y caused by an increase in x, which exceeds the negative impact on y caused by a comparable decline in x. The latter situation indicates that a decrease in x has a larger negative effect on y than the positive effect from an increase in x. Shin et al. (2014) combined Equation (2) with the conventional ARDL(p,q) approach of Pesaran et al. (2001), obtaining the following nonlinear ARDL model:

$$y_t = \sum_{j=1}^{p} \theta_j y_{j-i} + \sum_{j=0}^{q} \left(\mu_j^{+\prime} x_{t-j}^+ + \mu_j^{-\prime} x_{t-j}^-\right) + \varepsilon_t \tag{3}$$

where x_t is a k×1 vector of multiple regressors, defined so that $x_t = x_0 + x_t^+ + x_t^-$, θ_j is the autoregressive parameter, $\mu_j^{+\prime}$ and $\mu_j^{-\prime}$ are the asymmetric distributed-lag parameters, and ε_t is an independent and identically distributed process with zero mean and constant variance, σ_ε^2. Following Pesaran et al. (2001), Equation (3) can be represented by the following asymmetric error correction model:

$$\begin{gathered}\Delta y_t = \delta y_{t-1} + \varphi_x^+ x_{t-1}^+ + \varphi_x^- x_{t-1}^- + \sum_{j=1}^{p-1} \gamma_j \Delta y_{j-i} + \sum_{j=0}^{q-1} \left(\alpha_j^+ \Delta x_{t-j}^+ + \alpha_j^- \Delta x_{t-j}^-\right) + \varepsilon_\mathrm{t}, \\ \Delta y_t = \delta \mathrm{ECT}_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta y_{j-i} + \sum_{j=0}^{q-1} \left(\alpha_j^+ \Delta x_{t-j}^+ + \alpha_j^- \Delta \mathrm{x}_{\mathrm{t-j}}^-\right) + \varepsilon_t,\end{gathered} \tag{4}$$

where $\delta = \sum_{j=1}^{p} \theta_j - 1$; $\gamma_j = \sum_{i=j+1}^{p} \theta_i$ for $j = 1, \dots, p-1$; $\varphi_x^+ = \sum_{j=0}^{q} \mu_j^+$; $\varphi_x^- = \sum_{j=0}^{q} \mu_j^-$; $\alpha_0^+ = \mu_0^+$; $\alpha_j^+ = -\sum_{i=j+1}^{q} \mu_j^+$ for $i = 1, \dots, q-1$; $\alpha_0^- = \mu_0^-$; $\alpha_j^- = -\sum_{i=j+1}^{q} \mu_j^-$ for $i = 1, \dots, q-1$; and $\mathrm{ECT}_t = y_t - \beta^+ x_t^+ - \beta^- x_t^-$ is the nonlinear error correction term where $\beta^+ = -\varphi_x^+/\varphi_y$ and $\beta^- = -\varphi_x^-/\varphi_y$ are the asymmetric long-run coefficients. The long-run asymmetry is represented by φ^+ and φ^-. The short-run asymmetry is represented by α_j^+ and α_j^-. We test the long-run asymmetry through a Wald test with the null hypothesis that $\varphi^+ = \varphi^-$. We test the short-run symmetry through a Wald test of the null hypothesis that $\alpha_j^+ = \alpha_j^-$ $\forall i$.

4.4. Frequency-Domain Granger Causality Test

As stated in Section 4.3, economic events, such as changes in the economic environment, changes in monetary and/or fiscal policy, etc., can create room for a nonlinear (rather than linear) relationship between RES and GDP. This nonlinearity may also arise due to frictions in the economy and/or asymmetric information flow, which leads to market inefficiency (Tiwari, 2012: 1573). In fact, it is important to note that one of the key assumptions of Granger causality tests, both in time and in frequency-domains, is that of parameter stability, i.e., the parameters in equations relating the variables do not change over the sample under consideration (Gupta et al., 2015: 803). Since we consider a relatively long span of time (1960:Q1-2019:Q2), it is likely that the relationship between GDP, RES, G, E, and RD suffers from structural breaks. However, unlike with symmetric causality testing, the results of the asymmetric causality tests are not affected by parameter instability/structural breaks, since they control for any bias that parameter instability would cause (Ranjbar et al., 2017: 27). Thus, we analyze the issue of Granger causality between RES and GDP in an asymmetric framework.

With insights from Hatemi-J (2012), Toumi and Toumi (2019), and Kumar et al. (2019), we investigate asymmetric causality relationships between RES and GDP by incorporating the partial sum decompositions of Equation (1) (x_t^+ and x_t^-) using the augmented vector autoregression (VAR) approach developed by Toda and Yamamoto (1995). However, instead of implementing the single test statistic in the time domain, which assumes that the causal relationship holds for all points in the frequency distribution, we use the frequency-domain Granger causality test of Breitung and Candelon (2006), which is built on the previous works of Granger (1969), Geweke

(1982), and Hosoya (1991). Thus, with insights from Ranjbar et al. (2017), Chang et al. (2017), and Saliminezhad and Bahramian (2020), we combine the time-domain-asymmetric Granger causality framework with the symmetric-frequency-domain Granger causality approach.

In contrast to the time-domain Granger causality analysis, which produces a single, one-shot statistic regarding predictability, thereby describing a relationship between the variables that is supposed to apply across all periodicities (in the short run, over the business cycle frequencies, and in the long run), frequency domain or spectral analysis decomposes the variability of a time series into different frequencies that contribute to the fluctuations of the variable. Long-run dynamics are represented by lower frequency components of the spectra, while short-run linkages are represented by higher frequency components. Thus, it is assumed that the sensitivity of a variable to a temporary (high-frequency) shock in another variable is not the same as that to a permanent (low-frequency) shock.

Since spectral causality is defined by the spectral density of the "effect" variable, frequency-wise measures are nonlinear functions of the parameters of the vector autoregressive regression (VAR) model. However, Breitung and Candelon (2006) developed a spectral Granger causality test that is easier to implement. The test can identify whether a particular component of the "cause" variable at frequency ω is useful for predicting the component of the "effect" variable at the same frequency one period ahead (Tastan, 2015: 1158).

The details of the methodology developed by Breitung and Candelon (2006) have recently been applied and discussed in regard to studying causal relationships between macroeconomic variables in works such as Krätschell and Schmidt (2017), Fromentin and Tadjeddine (2019), Bouri et al. (2019), and Manera et al. (2020). Thus, we only display the primary conclusions of the method, stating that Granger causality between RES and GDP can be tested at any frequency ω between 0 and π.

Following Breitung and Candelon's (2006) notation, let us consider $\mathbf{Y}_\mathrm{t} = (x_t, y_t)'$ a time series vector with stationary covariance. It can be described by a VAR(p) process such as

$$\boldsymbol{\Theta}(\mathrm{L})\mathbf{Y}_{\mathrm{t}} = \boldsymbol{\varepsilon}_t, \tag{5}$$

where $\boldsymbol{\Theta}(\mathrm{L}) = \mathbf{I}_2 - \boldsymbol{\Theta}_1\mathrm{L} - \boldsymbol{\Theta}_2\mathrm{L}^2 - \ \ldots - \boldsymbol{\Theta}_{\mathrm{p}}\mathrm{L}^{\mathrm{p}}$ is a 2×2 lag polynomial with the backshift operator $\mathrm{L}^{\mathrm{i}}\mathbf{Y}_{\mathrm{t}} = \mathbf{Y}_{\mathrm{t-i}}$; $\mathbf{I}_2$ represents a 2×2 identity matrix; $\boldsymbol{\Theta}_{\mathrm{i}}$ with i =1 ,2,...,p denotes a 2×2 coefficient matrix associated with lag i; and $\boldsymbol{\varepsilon}_t$ is a vector white-noise process with $\mathrm{E}(\varepsilon_t)$=0 and positive-definite covariance matrix $\boldsymbol{\Sigma} = \mathrm{E}(\boldsymbol{\varepsilon_t}\boldsymbol{\varepsilon_t}')$. After Cholesky factorization, $\mathbf{G}'\mathbf{G} = \boldsymbol{\Sigma}^{-1}$ (where $\mathbf{G}$ is a lower-triangular matrix), a moving-average representation of the system in (5) can be represented as

$$\mathbf{Y}_t = \boldsymbol{\Phi}(\mathrm{L})\boldsymbol{\varepsilon}_t = \begin{bmatrix}\Phi_{\mathbf{11}}(\mathrm{L}) & \Phi_{\mathbf{12}}(\mathrm{L})\\ \Phi_{\mathbf{21}}(\mathrm{L}) & \Phi_{\mathbf{22}}(\mathrm{L})\end{bmatrix}\begin{bmatrix}\varepsilon_{\mathbf{1}\boldsymbol{t}}\\ \varepsilon_{\mathbf{2}\boldsymbol{t}}\end{bmatrix} = \boldsymbol{\psi}(\mathrm{L})\boldsymbol{\eta_t} = \begin{bmatrix}\Psi_{\mathbf{11}}(\mathrm{L}) & \Psi_{\mathbf{12}}(\mathrm{L})\\ \Psi_{\mathbf{21}}(\mathrm{L}) & \Psi_{\mathbf{22}}(\mathrm{L})\end{bmatrix}\begin{bmatrix}\eta_{\mathbf{1}\boldsymbol{t}}\\ \eta_{\mathbf{2}\boldsymbol{t}}\end{bmatrix},$$

where $\boldsymbol{\eta_t} = \mathbf{G}\boldsymbol{\varepsilon}_t$; $\mathrm{E}(\boldsymbol{\eta_t}, \boldsymbol{\eta'_t}) = \mathrm{I}$; $\boldsymbol{\Phi}(\mathrm{L}) = \Theta(\mathrm{L})^{-1}$; and $\boldsymbol{\psi}(\mathrm{L}) = \boldsymbol{\Phi}(\mathrm{L})\mathrm{G}^{-1}$.

The causality running from x to y at frequency ω can be measured, following Geweke (1982), as

$$\mathrm{M}_{x\to y}(\omega) = \log\left[1 + \frac{|\Psi_{12}(e^{-i\omega})|^2}{|\Psi_{11}(e^{-i\omega})|^2}\right]. \tag{6}$$

The condition $\Psi_{12}(e^{-i\omega}) = 0$ implies no Granger causality from x to y at frequency . Breitung and Candelon (2006) show that this condition is satisfied if the condition

$$\left|\Theta_{12}(e^{-i\omega})\right| = \left|\sum_{k=1}^{p}\theta_{k,12}\cos(k\omega) - i\sum_{k=1}^{p}\theta_{k,12}\sin(k\omega)\right| = 0 \tag{7}$$

is also satisfied, where $\theta_{k,12}$ is the (1,2) element of $\boldsymbol{\Theta}_k$. In this framework, the necessary and sufficient conditions for no causality are

$$\sum_{k=1}^{p}\theta_{k,12}\cos(k\omega) = 0, \tag{8}$$

and

$$\sum_{k=1}^{p} \theta_{k,12}\sin(k\omega) = 0. \tag{9}$$

To simplify the notation, following Breitung and Candelon (2006), let's denote $\gamma_k = \theta_{k,11}$ and $\beta_k = \theta_{k,12}$, and write the following VAR equation:

$$y_t = \gamma_1 y_{t-1} + \ldots + \gamma_p y_{t-p} + \beta_1 x_{t-1} + \ldots + \beta_p x_{t-\mathrm{p}} + e_{1t}. \tag{10}$$

The null hypothesis of no causality $\mathrm{M}_{x\to y}(\omega) = 0$ is equivalent to the linear restriction $\mathrm{H}_0\colon \mathrm{R}(\omega)\beta = 0$, where β is $\left[\beta_1, \ldots, \beta_p\right]'$ and $\mathrm{R}(\omega)$ is $\begin{bmatrix}\cos(\omega) & \cos(2\omega) & \ldots\cos(\mathrm{p}\omega)\\ \sin(\omega) & \sin(2\omega) & \ldots\sin(\mathrm{p}\omega)\end{bmatrix}$.

According to Breitung and Candelon (2006), in a cointegrating system such as

$$\Delta\mathbf{Y}_t = \Pi\mathbf{Y}_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \,\Delta\mathbf{Y}_{t-\mathrm{i}} + e_t,$$

testing the null of no causality at frequency zero is interesting, since in this special case $\Theta(e^0) = \Theta_1 - \mathrm{I} + \Theta_2 + \ldots + \Theta_\mathrm{p} = \Pi$, which is called the impact matrix. Testing the null of no causality at frequency zero in a cointegrating system would imply testing $\mu_{12} = 0$ in the following regression:

$$\Delta x_t = \mu_{11}x_{t-1} + \mu_{12}y_{t-1} + \sum_{i=1}^{p-1}\gamma_{11i}\,\Delta x_{t-i} + \sum_{i=1}^{p-1}\gamma_{12i}\,\Delta y_{t-i} + e_{1t}, \tag{11}$$

where μ and γ denote the (i,j) element of Π and Γ_k, respectively. Testing $\mu_{12} = 0$ is equivalent to testing long-run causality (Pittis, 1999).

If in the bivariate cointegrating system x is integrated of order 0 and y is integrated of order 1, the Wald test for the hypothesis $\mu_{12} = 0$ does not have a standard limiting distribution (Sims et al., 1990). Similar problems are faced in higher-dimensional systems if a block of the matrix Π is singular (Breitung and Candelon, 2006: 368). Breitung and Candelon (2006) suggested that the Toda and Yamamoto (1995) method can be used to overcome this difficulty.

The Toda and Yamamoto (1995) approach to causality is considered an upgraded version of the Wald test. The method obtains efficient and consistent estimates even when the variables have different orders of integration. The Toda and Yamamoto (1995) model is constructed through a VAR in levels. In our research, the VAR specification would be as follows:

$$\begin{aligned}
\mathrm{GDP}_t = \mu_{1t} &+ \sum_{i=1}^{p} \alpha_{1i}\,\mathrm{GDP}_{t-\mathrm{i}} + \sum_{k=p+1}^{d_{\max}} \alpha_{1k}\,\mathrm{GDP}_{t-k} + \sum_{i=1}^{p} \gamma_{11i}\,\mathrm{RES}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{11k}\,\mathrm{RES}^{+}_{t-k} \\
&+ \sum_{i=1}^{p} \gamma_{12i}\,\mathrm{RES}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{12k}\,\mathrm{RES}^{-}_{t-k} + \sum_{i=1}^{p} \gamma_{13i}\,\mathrm{G}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{13k}\,\mathrm{G}^{+}_{t-k} \\
&+ \sum_{i=1}^{p} \gamma_{14i}\,\mathrm{G}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{14k}\,\mathrm{G}^{-}_{t-k} + \sum_{i=1}^{p} \gamma_{15i}\,\mathrm{E}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{15k}\,\mathrm{E}^{+}_{t-k} \qquad (12) \\
&+ \sum_{i=1}^{p} \gamma_{16i}\,\mathrm{E}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{16k}\,\mathrm{E}^{-}_{t-k} + \sum_{i=1}^{p} \gamma_{17i}\,\mathrm{RD}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{17k}\,\mathrm{RD}^{+}_{t-k} \\
&+ \sum_{i=1}^{p} \gamma_{18i}\,\mathrm{RD}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{18k}\,\mathrm{RD}^{-}_{t-k} + e_{1t},
\end{aligned}$$

$$\begin{aligned}
\mathrm{RES}^{+}_t = \mu_{2t} &+ \sum_{i=1}^{p} \alpha_{2i}\,\mathrm{GDP}_{t-i} + \sum_{k=p+1}^{d_{\max}} \alpha_{2k}\,\mathrm{GDP}_{t-k} + \sum_{i=1}^{p} \gamma_{21i}\,\mathrm{RES}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{21k}\,\mathrm{RES}^{+}_{t-k} \\
&+ \sum_{i=1}^{p} \gamma_{22i}\,\mathrm{RES}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{22k}\,\mathrm{RES}^{-}_{t-k} + \sum_{i=1}^{p} \gamma_{23i}\,\mathrm{G}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{23k}\,\mathrm{G}^{+}_{t-k} \\
&+ \sum_{i=1}^{p} \gamma_{24i}\,\mathrm{G}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{24k}\,\mathrm{G}^{-}_{t-k} + \sum_{i=1}^{p} \gamma_{25i}\,\mathrm{E}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{25k}\,\mathrm{E}^{+}_{t-k} \qquad (13) \\
&+ \sum_{i=1}^{p} \gamma_{26i}\,\mathrm{E}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{26k}\,\mathrm{E}^{-}_{t-k} + \sum_{i=1}^{p} \gamma_{27i}\,\mathrm{RD}^{+}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{27k}\,\mathrm{RD}^{+}_{t-k} \\
&+ \sum_{i=1}^{p} \gamma_{28i}\,\mathrm{RD}^{-}_{t-i} + \sum_{k=p+1}^{d_{\max}} \gamma_{28k}\,\mathrm{RD}^{-}_{t-k} + e_{2t},
\end{aligned}$$

$$\begin{aligned}
\mathrm{RES}_t^- = \mu_{3t} &+ \sum_{i=1}^{p} \alpha_{3i}\,\mathrm{GDP}_{t-i} + \sum_{k=p+1}^{d_{\max}} \alpha_{3k}\,\mathrm{GDP}_{t-k} + \sum_{i=1}^{p} \gamma_{31i}\,\mathrm{RES}_{t-i}^+ \\
&+ \sum_{k=p+1}^{d_{\max}} \gamma_{31k}\,\mathrm{RES}_{t-k}^+ \sum_{i=1}^{p} \gamma_{32i}\,\mathrm{RES}_{t-i}^- + \sum_{k=p+1}^{d_{\max}} \gamma_{32k}\,\mathrm{RES}_{t-k}^- + \sum_{i=1}^{p} \gamma_{33i}\,\mathrm{G}_{t-i}^+ \\
&+ \sum_{k=p+1}^{d_{\max}} \gamma_{33k}\,\mathrm{G}_{t-k}^+ + \sum_{i=1}^{p} \gamma_{34i}\,\mathrm{G}_{t-i}^- + \sum_{k=p+1}^{d_{\max}} \gamma_{34k}\,\mathrm{G}_{t-k}^- + \sum_{i=1}^{p} \gamma_{35i}\,\mathrm{E}_{t-i}^+ \qquad (14) \\
&+ \sum_{k=p+1}^{d_{\max}} \gamma_{35k}\,\mathrm{E}_{t-k}^+ + \sum_{i=1}^{p} \gamma_{36i}\,\mathrm{E}_{t-i}^- + \sum_{k=p+1}^{d_{\max}} \gamma_{36k}\,\mathrm{E}_{t-k}^- + \sum_{i=1}^{p} \gamma_{37i}\,\mathrm{RD}_{t-i}^+ \\
&+ \sum_{k=p+1}^{d_{\max}} \gamma_{37k}\,\mathrm{RD}_{t-k}^+ + \sum_{i=1}^{p} \gamma_{38i}\,\mathrm{RD}_{t-i}^- + \sum_{k=p+1}^{d_{\max}} \gamma_{38k}\,\mathrm{RD}_{t-k}^- + e_{3t}
\end{aligned}$$

where p is the lag length, and $d_{\max}$ is the maximum order of integration that the variables need to become stationary. The null hypothesis of no causality running from RES^+ to GDP at a frequency ω ($\mathrm{M}_{\mathrm{RES}^+\to\mathrm{GDP}}(\omega)$) involves only $\gamma_{11i}, i = 1,\ldots,p$, which implies that the coefficients of the additional lagged variables are not included in the computation of the Wald statistic. The H_0 can be tested by means of the F-test, which follows an $\mathrm{F}(p,\ \mathrm{T} - 2p)$ distribution for every ω between 0 and π.

5. Empirical Results

In this section, we present the empirical results of our research. We implement the nonlinear ARDL test of Shin et al. (2014) and the symmetric- and asymmetric-frequency-domain Granger causality test of Breitung and Candelon (2006), using the modified Wald test of Toda and Yamamoto (1995).

Table 1. Summary statistics

	GDP	RES	G	E	RD
Mean	9.07	5.94	7.67	6.41	5.44
Std. Dev.	0.51	0.48	0.32	0.99	0.58
Skewness	–0.22	–0.07	–0.28	–0.13	–0.26
Kurtosis	1.84	2.14	1.83	1.69	1.96

Notes: This table shows the descriptive statistics of the series GDP, RES, G, E, and RD. The data cover the period between 1960:Q1 and 2019:Q2. We dispose of a total of 238 quarterly observations for each variable.

The first step of our research is to investigate the degree of integration of the variables through the Breakpoint Unit Root test of Perron (1997). The test suggests that the variables are integrated of order one (Table 2).

Table 2. Breakpoint Unit Root tests. Exogenous variables: individual effects and individual linear trends.

Tests	Variables									
	GDP	ΔGDP	RES	ΔRES	G	ΔG	E	ΔE	RD	ΔRD
Test Statistic	–4.32	–11.80***	–4.84	–8.95***	–3.91	–5.39**	–3.04	–19.99***	–4.71	–5.13*
P-value	0.32	0.00	0.11	0.00	0.58	0.03	0.95	0.00	0.15	0.05

Notes: This table presents the Perron (1997) unit root test statistics. The asterisks *, **, and *** denote rejection of the null hypothesis of a unit root process at the 10%, 5%, and 1% significance level, respectively. The break specification is intercept and trend, and the type of break is an innovational outlier. The lag length is selected by the Bayesian information criterion (BIC) of Schwarz (1978).

Tables 3 and 4 report the results of the nonlinear ARDL test of Shin et al. (2014). The lag selection is carried out following the Akaike's information criterion (AIC) of Akaike (1974) and the Bayesian information criterion (BIC) of Schwarz (1978). The lag length of the ARDL specification is validated by the absence of serial correlation in the residuals through the Breusch-Godfrey Lagrange multiplier test, which tests the null hypothesis that the errors are serially independent (see Breusch, 1978 and Godfrey, 1998). Additionally, the diagnostic tests reveal that there is no misspecification in the short- and long-run estimates, the non-normality of the error term, the serial correlation, and heteroscedasticity.

Table 3. Vector error correction model derived from the NARDL(6, 7, 7, 3, 0, 0, 1, 0, 1). Dependent variable: ΔGDP_t

Regressor	Coefficient	Std. Error	t-Statistic	*P*-value
ΔGDP_{t-1}	–0.06	0.06	–1.06	0.29
ΔGDP_{t-2}	0.08	0.06	1.44	0.15
ΔGDP_{t-3}	0.09	0.05	1.72	0.09
ΔGDP_{t-4}	0.07	0.05	1.28	0.20
ΔGDP_{t-5}	–0.10	0.05	–1.98	0.05
ΔRES_t^+	0.04***	0.02	2.99	0.00
ΔRES_{t-1}^+	0.04***	0.02	2.68	0.01
ΔRES_{t-2}^+	0.02	0.02	1.01	0.31
ΔRES_{t-3}^+	0.01	0.02	0.40	0.69
ΔRES_{t-4}^+	0.02	0.02	1.02	0.31
ΔRES_{t-5}^+	0.02	0.02	1.39	0.17
ΔRES_{t-6}^+	0.04	0.01	2.54	0.01
ΔRES_t^-	0.11***	0.02	7.13	0.00
ΔRES_{t-1}^-	0.06***	0.02	3.43	0.00
ΔRES_{t-2}^-	–0.02	0.02	–1.07	0.29
ΔRES_{t-3}^-	–0.03	0.02	–2.12	0.04
ΔRES_{t-4}^-	0.04***	0.02	2.67	0.01
ΔRES_{t-5}^-	–0.04	0.02	–2.10	0.04
ΔRES_{t-6}^-	–0.03	0.02	–1.82	0.07
ΔG_t^+	0.32***	0.05	6.37	0.00
ΔG_{t-1}^+	0.11	0.05	2.04	0.04
ΔG_{t-2}^+	0.09	0.05	1.72	0.09
ΔG_t^-	0.09	0.09	1.01	0.32
ΔE_t^+	0.09***	0.01	6.32	0.00
ΔE_t^-	0.00	0.02	0.27	0.79
ΔRD_t^+	0.12***	0.04	2.72	0.01
ΔRD_t^-	0.12	0.09	1.36	0.18
ECT_{t-1}	–0.28***	0.03	–8.65	0.00

	F-Statistic	*P*-value
Asymmetric short-run coefficients	13.47***	0.00

Notes: This table display the results of the vector error correction model derived from the NARDL of Shin et al. (2014) (Equation (4)). The asterisks *** indicate statistical significance at the 1% significance level. The lag order specification is adopted by the AIC and BIC. We also present the results of the short-run asymmetry test between the effects of ΔRES^+ and ΔRES^-. The asterisks *** indicate rejection of the null hypothesis of symmetry at the 1% level.

On the other hand, in Table 4 we see that the long-run coefficients of RES^+ and RES^- carry their theoretically expected positive and significant coefficients, but we are not sure if the coefficients are significantly different from each other. To determine if this difference is statistically significant, we apply the standard F-test on the null hypothesis that $\beta^+ = \beta^-$, while the null hypothesis of short-run asymmetry is described by $\sum_{i=0}^{q} \alpha_j^+ = \sum_{i=0}^{r} \alpha_j^-$. In Tables 3 and 4 we see that the F-tests are significant, which implies rejecting the null hypothesis of symmetry both in the short and in the long run. Thus, RES has asymmetric short- and long-run effects on GDP. Asymmetries are important, since, in our case, they indicate that cyclical upturns and downturns in RES have asymmetrical effects on economic activity. For example, Table 4 shows that the estimated long-run coefficients of RES^{-} are higher than those of RES^{+}. It implies that, in the long run, a decrease in RES has a stronger impact on economic activity than a proportionally equal increase in RES.

Table 4. Long-run coefficients. Dependent variable: GDP_t

Variable	Coefficient	Std. Error	t-Statistic	*P*-value
RES^{+}	0.09***	0.01	6.42	0.00
RES^{-}	0.13***	0.01	15.88	0.00
G^{+}	0.02	0.08	0.22	0.83
G^{-}	0.43***	0.16	2.63	0.01
E^{+}	0.30***	0.03	12.20	0.00
E^{-}	0.16***	0.02	8.78	0.00
RD^{+}	0.32***	0.03	10.27	0.00
RD^{-}	–0.03	0.08	-0.37	0.71
Constant	8.05***	0.01	651.86	0.00
		F-statistic	*p*-value	
Asymmetric long-run coefficients		8.99***	0.00	

Notes: This table presents the coefficients of the long-run equilibrium relationship among the variables under study (Equation 2). The asterisks *** indicate statistical significance the 1% significance level. We also present the results of the long-run asymmetry test between the effects of RES^{+} and RES^{-}. The asterisks *** indicate rejection of the null hypothesis of symmetry at the 1% level.

The calculated long-run effects are only meaningful in the presence of cointegration. In Table 3 we observe that the coefficient of the lagged error correction term (ECT_{t-1}) is negative and statistically significant. Thus, following the Engle and Granger (1987) approach, we conclude that a significant and negative coefficient for ECT_{t-1} supports the existence of cointegration between GDP, RES, G, E, and RD. A significant and negative coefficient for ECT_{t-1} does not only indicate convergence, but also the speed of adjustment to equilibrium. Thus, there is a nonlinear cointegrating relationship between the variables at the 1% significance level.

The existence of cointegration indicates that our proposed model is analytically valid, since, if there exists a cointegrating relationship between a set of nonstationary variables, the same cointegrating relationship also exists in an extended variable space (Johansen, 2000). The presence of cointegration avoid finding spurious causality relationships and justifies the non-inclusion of additional variables, which would unnecessarily increase the number of cointegrating equations that would have to be identified and estimated (Herzer and Nunnenkamp, 2012: 253; Perez-Montiel and Manera, 2020a: 227).

Once we have confirmed that the variables follow a long-run equilibrium relationship, we implement the Breitung and Candelon (2006) frequency-domain Granger causality test using the modified Wald test of Toda and Yamamoto (1995). Since the unit root tests suggest that the variables are I(1) processes, we establish $d_{\max}$=1. Following Toda and Yamamoto (1995) and Breitung and Candelon (2006), we fit a VAR(p+$d_{\max}$) model, but implement the test using p lags. We also include a time trend in the VAR system. We test symmetric and asymmetric causality by also considering the partial sum decompositions specified in Equation 2 in the Toda Yamamoto-Breitung Candelon approach. Thus, we test for symmetric causality relationships between RES and GDP conditioning on G, E, and RD; for asymmetric causality relationships between RES^{+} and GDP conditional on G^{+}, E^{+}, RD^{+}, RES^{-}, G^{-}, E^{-} and RD^{-}; and, finally, for asymmetric causality relationships between RES^{-} and GDP conditional on G^{+}, E^{+}, RD^{+}, RES^{+}, G^{-}, E^{-} and RD^{-}.

We run the test using different lag orders, p (from 3 to 6, since 3 is the minimum number of lags necessary to get a sufficient dynamic structure in the model to perform the frequency decomposition).[6] We run the frequency-domain Granger causality tests using the Stata command *bcgcausality* by Tastan (2015).

Table 5 displays the results of the frequency-domain Granger causality tests. The presence of Granger causality is examined at frequencies 0.05 and 2.5. These frequencies are defined as long (0.05) and short (2.5) periodicities. At the same time, $\omega = 0.05$ is defined as permanent causality, whereas $\omega = 2.5$ is defined as temporary causality (see Bodart and Candelon, 2009). Table 5 shows that we cannot reject the null hypothesis of no symmetric causality running from GDP to RES, conditional on G, E, and RD, at any significance level for frequencies 0.05 and 2.5. On the other hand, we reject the null hypothesis of no causality running from RES to GDP at the 1% significance level for frequencies 0.05 and 2.5. These results suggest both transitory and permanent symmetric

[6] We are thankful to Jörg Breitung for this suggestion.

Granger causality running unidirectionally from the growth rate of RES to the growth rate of GDP.

In Table 5, we also present the results of the asymmetric-frequency-domain Granger causality tests. In line with the symmetric causality results, we observe that the positive and negative decompositions have a unidirectional causal effect on output growth.

Table 5. Frequency-domain causality between GDP and RES, conditional on G, E, and RD. Symmetric model: $\mathrm{GDP} = f(\mathrm{RES}, \mathrm{G}, \mathrm{E}, \mathrm{RD})$;
Asymmetric model: $\mathrm{GDP} = f(\mathrm{RES}^+, \mathrm{G}^+, \mathrm{E}^+, \mathrm{RD}^+, \mathrm{RES}^-, \mathrm{G}^-, \mathrm{E}^-, \mathrm{RD}^-)$

	Lags	Symmetric causality		Asymmetric causality $\mathrm{RES}^+(\mathrm{GDP}) \rightarrow \mathrm{GDP}(\mathrm{RES}^+)$		Asymmetric causality $\mathrm{RES}^-(\mathrm{GDP}) \rightarrow \mathrm{GDP}(\mathrm{RES}^-)$	
		$\omega = 0.05$	$\omega = 2.5$	$\omega = 0.05$	$\omega = 2.5$	$\omega = 0.05$	$\omega = 2.5$
RES → GDP	3	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***
	4	0.00***	0.00***	0.00***	0.02**	0.00***	0.00***
	5	0.00***	0.00***	0.00***	0.37	0.01**	0.00***
	6	0.00***	0.00***	0.00***	0.25	0.02**	0.00***
GDP → RES	3	0.75	0.77	0.56	0.41	0.13	0.16
	4	0.75	0.64	0.21	0.09	0.06	0.20
	5	0.34	0.22	0.69	0.09	0.47	0.23
	6	0.38	0.34	0.96	0.13	0.51	0.17

This table shows the results of the frequency-domain Granger causality tests. We report the *p*-values corresponding to the estimated Wald test statistic for the null hypothesis of no symmetric Granger causality running from RES(GDP) to GDP(RES), conditional on G, E, and RD (first column). We also report the results of asymmetric-frequency-domain Granger causality running from RES^+(GDP) to GDP(RES^+), conditional on G^+, $\mathrm{E}^+, \mathrm{RD}^+, \mathrm{RES}^-, \mathrm{G}^-, \mathrm{E}^-$, and RD^- (second column); and from RES^-(GDP) to GDP(RES^-), conditional on G^+, $\mathrm{E}^+, \mathrm{RD}^+, \mathrm{RES}^+, \mathrm{G}^-, \mathrm{E}^-$, and RD^- (third column). We employ different lag lengths, *p*, (from 3 to 6). We test Granger causality at frequency ω=0.05 and ω=2.5. The asterisks *** and ** denote rejection of the null hypothesis of no Granger causality at the 1% and 5% significance level.

The results presented in Table 5 indicate that there is enough statistical information to confirm the existence of short- and long-run Granger causality running unidirectionally from the growth of RES to the growth of GDP, suggesting that the dynamics of RES leads the evolution of GDP both in the short and in the long run. Overall, our results suggest that RES can be considered a driver of the secular growth rate of the US economy. At the same time, RES has determined the business cycle in the US since the sixties.

As a side result, the data in Tables 4 and 5 allow us to relate our findings with the autonomous demand-led growth literature. According to this literature, at the theoretical level, the whole of autonomous demand is the driving force of output growth in the long run. In this paper, we have shown that in the US during the time span considered, a prominent role has been played by a specific component of autonomous demand: residential investment. It does not seem too far-fetched to interpret this as a specific feature of American capitalism and its growth process, which are heavily based on credit to households.

6. Concluding Remarks

In this article, we have highlighted the role of residential investment as a driver of both the cycle and secular growth in the US. The main methodological novelty of our contribution is the combination of the asymmetric time-domain causality approach with the spectral causality framework, which provides an empirical strategy that allows testing for asymmetric frequency-domain causality relationships between the variables. In this way, we have found evidence of symmetric and asymmetric causality running unidirectionally from the growth of residential investment to GDP growth, both in the short and in the long run.

According to Cesaratto (2015) and Fiebiger (2018), residential investment, while internal to the dynamics of a private capitalist economy, can be analogous to a Luxemburg-Kalecki external market. Our research provides an empirical validation for this assertion for the US economy, since we find evidence that the dynamics of residential investment drive the evolution of output, while the output does not exert any significant influence on the evolution of the residential investment. This allows us to relate our contribution to the growing body of Keynesian literature on autonomous demand-led growth. We single out a specific component of autonomous demand and describe its prominent role in the US variety of capitalism. Thus, we conclude that residential investment, despite constituting a small overall share of GDP, *is not only the cycle but is also the trend* of the US economy.

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