MACROECONOMIC MEASUREMENT OF EXPECTATIONS VERSUS REALITY

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Abstract

This study analyzes the combined effects of "sensitivity of economic expectations" (responsiveness of expectations to the effective dynamics of the economy) and the "degree of self-fulfillment of expectations" (proportion of these expectations that are translated into reality). Their interactions (named "expectational impulse") are revealed by the evolution of the concerned process (speed and direction). Unlike many previous studies that have focused on the specificity of this phenomenon in different business segments, our study examines its configuration at a macroeconomic level. We start from the premise that despite the heterogeneity of expectation forming mechanisms within the microenvironment, expectations manifest as a congruent economic zeitgeist at a macro level. Empirically, our study uses quarterly estimations of the economic sentiment index (ESIq), gross domestic product index in real terms (IGDPq), expectation sensitivity (mtzq), and self-fulfillment degree of expectations (sfdq) for all the European Union countries using the pre-Brexit format for the period 1995Q1–2020Q4. These data were processed using AR lag-distributed techniques. The key finding of the study is that when applied to successive post-sample simulations, all the experimented models generated steady-state type of estimates, defined as attractors. However, this view must be understood at a broader level as it is not only a matter of an equilibrium or normatively desirable levels of mtzq and sfdq, but more about their presumptive state provided the current dominant market behaviors do not change. In other words, the estimated attractors only represent signals concerning the economic expectations under a static business behavioral framework.

Keywords: economic expectations sensitivity, self-fullfilment degree of expectations, expectational impulse, AR lag-distributed model, post-sample simulation, attractor

JEL Classification: A14, C12, C15, C63

1. Introduction

The present research started from the premise that expectations are formed on the basis of incidence of two types of causal factors: i) those derived from the characteristics of the socio-economic, technological, and institutional framework that is dominant in the respective historical period, thus exerting a profound and stable enough influence; and ii) those consisting of permanent changes that occur in the behaviors of producers and buyers,

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thereby leaving a pronounced volatile impact on the entire process. The first category induces what we name the inertial component of expectations. Focusing on this component, this study attempts to identify whether it statistically manifests as a kind of attractor. The hypothesis is empirically analyzed at an aggregate level, with reference to the experience of the European Union countries, in a pre-Brexit format, between 1995 and 2020.

Early studies on the sensitivity of economic expectations were focused on prices, interest rates, and the movement of inventories (Metzler, 1941; Lachmann, 1943; Hicks, 1946; Tropeano, 1997; Lenfant, 2020). Despite the incontestable heterogeneity of the formation of expectations at a micro-economic level, we directed this analysis towards the most synthetical indicators of economic activity. In our opinion, two circumstances are responsible for the transformation of diversity of expectations at the micro-level (random-walk variants, a large series of auto-corrective algorithms and rational paradigm) into a relatively congruent economic zeitgeist. The sociological interpretation of this concept is discussed by multiple studies including Mayo and Nohria (2005), Rodrik (2015), Davies (2018), Krause (2019), and Hameleers and Vliegenthart (2020). On the one hand, the deep interdependencies developed among different economic activities compel the market operators to communicate with partners about their projected deals through which their initial business intentions adapt to the requirements of contractual transactions. On the other hand, the modern economy is characterized by intensive macroeconomic guidance exerted by governments, central banks, and some international organizations that inform economic agents about conditions in which the global or regional commercial and financial inter-flows operate, so that the national economy and their main sector can evolve within predictable coordinates. As a result of these two processes (Dobrescu, 2020), the heterogeneity of expectations of economic agents as separate entities is considerably attenuated by aggregation. Due to this homogenizing effect, the expectational mechanism comes closer to the rational paradigm at the scale of the entire economic system.

In the case of the European Union countries, the sensitivity of aggregate expectations and degree of expectations being self-fulfilled are approximated by using two leading measurements: i) the economic sentiment index (ESI), and ii) the gross domestic product as an index (IGDP), which are both considered at comparable prices. The first quantifies the economic expectations, while the second represents the reality as compared to the first outcome.

The ESI (Eurostat, 2019), a composite indicator, is estimated as the weighted average of managers' and consumers' answers about the actual state and short-run perspectives of the economy's key sectors (European Commission, 2016, 2020). The ESI estimation questionnaires distributed in 2020 to over 220,000 units revealed that 55–60% of the respondents provided the requested information effectively. The structural weights involved in the aggregation of partial indicators were defined based on the relative impact of the respective fields on global economic growth. Therefore, the ESI surveys directly or indirectly cover several micro and macroeconomic indicators, including production, export, import, investment, building activity, employment, unemployment, financial situations, selling prices, consumer prices, purchase intentions, saving intentions, companies' profitability, and general (sectoral, national and EU) economic conjuncture (European Commission, 2020). Consequently, the national ESI time series was considered as consistently opposable to the corresponding GDP index ones.

A reliable measurement of the expectational phenomenon that mirrors the economic reality imposes a particular constraint regarding the temporal frequency of the compared data. Annual indexes are inappropriate because many fluctuations of the expectations-reality binomial could be blurred within this period as a result of temporal aggregative operations. From this perspective, monthly data would be preferable; however, they can only currently be calculated for the ESI, and not for the IGDP, which is officially recorded at quarterly intervals. Thus, this presents two available technical possibilities: to generate monthly series of the GDP index from the existing information based on its monthly components, or to generate quarterly series of the ESI from its monthly levels. Following Dobrescu (2020, Box 3), we opted for the second alternative. Therefore, this study's statistical analysis is based on quarterly data (ESIq and IGDPq).

The paper proceeds as follows: Section 2 covers the techniques used to identify aggregate expectation sensitivities and degrees of self-fulfilment of expectations in the European Union. Simulated attractors are commented in Section 3. Section 4 highlights the conclusions derived from our study.

2. Methodology

Given the economic expectations (denoted by *x*) and the reference economic reality (denoted by *y*), the expectation sensitivity and the degree of self-fulfilment of expectations are formalized. Dynamically, the sensitivity of expectations can be studied either by referring to i) the first order differences, or ii) the integral involved indicators. According to Metzler (1941, p.119), the coefficient of expectation (η) is "defined as the ratio of the expected change of sales between periods t and t- i to the observed change of sales between periods t- i and t- 2." Thus, based on our notation, Metzler's definition is represented as:

$$\eta=(x-x(-1))/(y(-1)-y(-2))=dx/dy(-1) \qquad (1)$$

Such a coefficient has the advantage of immediately revealing any changes that occur in the expectational process. We adopt a slightly modified version of η, denoted by mtz. Using the indicator levels instead of their first order differences results in the following formula:

$$mtz=x/y(-1) \qquad (2)$$

Normally, mtz and η are congenitally linked. The equation x=mtz*y(-1) algebraically yields:

$$\eta=(mtz*y(-1)-x(-1))/(y(-1)-y(-2)) \qquad (1a)$$

$$mtz=\eta*(1-y(-2)/y(-1))+x(-1)/y(-1) \qquad (2a)$$

Upon comparing these two formalizations of expectation sensitivity, some technical challenges raised by η cannot be ignored. The numerator (y(-1)-y(-2)) can assume very small values, which may result in the time series of η registering abnormally high standard deviations. However, the more uncomfortable situations occur when η equals zero as this would require the adoption of some conventional substitutes or the amputation of the incriminated data. For example, within our database, such "divide zero" cases exceed four percent of the total number of estimable η. To avoid these possible disadvantages, the expectation sensitivity at the macroeconomic level is analyzed by mtz that is a robust indicator, with the added advantage to usually generate a positive time series of more limited standard deviations.

Admitting that real processes are a materialization of the antecedents inducing expectations, the degree of self-fulfilment of expectations (sfd) is approximated by the ratio:

$$sfd=y/x \quad (3)$$

If the expectational mechanism comes close to the rational paradigm (REH), the ratio (3) should tend towards unity.

Combining (2) and (3), the variation of economic processes envisaged during two successive time sequences can be quantified by the product of mtz and sfd:

$$y=mtz*sfd=mtz*y(-1)*sfd \quad (4)$$

$$y/y(-1)=mtz*sfd \quad (4a)$$

Our proposal is to define the expectational impulse as:

$$eim=mtz*sfd-1 \quad (5)$$

This means that eim=0 corresponds to the equality where y=y(-1), which implies the stability of the reached state; eim>0 reflects its expansion, while eim<0 represents a converse dynamic.

The previously discussed measures consider an essential feature of the expectational process, long known by the economics literature: "it remains true that our expectations for the future are a response to our experience of the past…" (Lachmann, 1943, p. 19). Thus, while the mechanism of forming expectations does not ignore forward information concerning the economic process of interest (Stavrakeva and Tang, 2020; Goldstein and Gorodnichenko, 2022), they do not annul but rather, interact with "our experience of the past."

In turn, these expectations (as a core of managerial decisions) have a crucial impact on the effective economic activity. As a consequence, the time series of expectation sensitivities and the degrees of their self-fulfilment are intrinsically auto-regressive processes. Besides, these series are lag-distributed because past experience is appreciated differentially by the current perceptions of market operators. The trajectory of such a series can be irregular or conform to certain patterns. Sometimes, they are drastic (in ascending or descending trends), while at other times, they tend to have asymptotical configurations (fixed points, periodical oscillations). The main goal of our empirical research is to identify whether the series of aggregate expectation sensitivities (mtz) and degrees of self-fulfilment of expectations (sfd) in the European Union countries contain second type sequences (steady-states).

The lag-length of the corresponding AR series is not a simple technical question. The informational criteria are likely to be involved, and our analysis will also apply some econometric tools, such as likelihood ratio (Lütkepohl, 2005; Hatemi-J and Hacker, 2009) and final prediction error (Ljung, 1999; Akaike, 1969 and 1973; Schwarz, 1978; Hannan and Quinn, 1979). The problem is that such tests particularly focus on the accuracy of regressions (squared residuals, R-squared), but do not indicate the likelihood that the examined series may generate steady-states.

2.1 The LStAR Model

The more appropriate mode in our context may be the longest stable auto-regressive model (LStAR), which is a model that represents an auto-regressive process with the highest admissible number of lags, while also satisfying the stability condition (no root of the characteristic polynomial lies outside the unity circle). The first feature (the highest number of lags) signifies the best attainable valorisation of the available sample data, while the second (satisfying the stability condition) recognizes a possible convergence of the successive post-sample iterations. The LStAR methodology (or LSVAR in other previous studies) can be segmented into the following five steps.

2.1.1 AR Length Determination

The starting point is an ARDL specification for the sample interval (representing s observations):

$$y_t = c_0 + \Sigma c_i * y_{t-i} + u_t \quad (6)$$

where: y is the modelled variable; t=1, 2…s; i represents the preliminary tried lags (i=1, 2…(s-1)); c are the econometric coefficients estimated for the tried lags; u is the error term (residuals). The resultant estimation output from each variant is applied to the stability condition test. Should the stability condition be observed, the number of tried lags is increased, and vice-versa. This exploratory operation is continued until the longest possible stable AR process is revealed (denoted by p).

2.1.2 The LStAR Model

The operational LStAR model is represented as:

$$y_t = c_0 + \Sigma c_j * y_{t-j} + u_t \quad \text{where } j=1, 2\ldots p \quad (7)$$

The residuals are then submitted for the usual econometric tests, depending on which model is accepted for the post-sample iterations or is correspondingly re-specified.

2.1.3 Post-sample Simulations (Interval τ)

$$\tilde{y} = c_0 + \Sigma c_i * y_{t+\tau-i} + \varepsilon_{t+\tau} \quad (8)$$

where: $\tilde{y}$ represents the post-sample estimation; τ=1, 2…h (h is the last term of the post-sample interval). Extrapolating the error term was not considered necessary because our application does not aim to anticipate the deviations of the future effective evolution of series y from the model. On the contrary, it only tries to approximate its most probable state if the current functional context is not modified. In other words, the post-sample simulations are not forecasts, but simply numerical iterations.

2.1.4 Monitoring the Volatility of Simulations

The dynamics of $\tilde{y}_\tau$ are evaluated by the successive relative first-order differences (r_τ) (in module):

$$r\tilde{y}_\tau = d(\tilde{y}_\tau)/\tilde{y}_{\tau-1} \quad (9)$$

The convergence of the series $\tilde{y}_\tau$ is admitted as significant if the below inequality (10) holds a given number of times (ϖ):

$$d(\tilde{y}_\tau)/\tilde{y}_{\tau-1} \leq \Lambda \quad (10)$$

where: Λ is the critical threshold below which the modification of the successive simulations could be considered negligible. Both ϖ and Λ are exogenous, as an expression of the modeler's experience.

2.1.5 Inertial Component of the mtz and sfd Series

If the inequality (10) remains valid during the interval ($\tau\tau$), the mean leading to the corresponding estimations $\tilde{y}_\tau$ is considered an acceptable approximation of the inertial component within the analyzed series (denoted by ỹ@). The exogenous factors adopted in our research are $\tau\tau$ =100 and Λ =0.00001.

As a result of modelling weaknesses and/or time series peculiarities, it is possible that in some cases the iterations do not generate socio-economically plausible results or do not even converge (despite a prior positive stability test). In such situations, the estimation work will be reconsidered by diminishing the number of specified lags, adopting other modelling techniques, or by simply accepting that the given series is unable to generate a convergent auto-regression.

How to interpret the asymptote of the successive post-sample LStAR iterations which are finalized into ȳ@? While mathematically it is a "steady-state", it is not clear if socio-economically this is also the case. We could define these as the "steady state of aggregated sensitivity of expectations" and the "steady state of aggregated degree of self-fulfilment of expectations," respectively. However, in the social sciences, the "steady state" syntagm has been consecrated during recent decades as a notion of larger and normative significance, derived from the durable development theory (Daly, 1974, 1980, 1991; Czech and Daly, 2004; Anderson, 2012; Kenton, 2022; *Rua, 2021; Czech, 2021).* By contrast, the previously described concept of ȳ@ substantially differs from this interpretation.

Initially, we used the definition "inertial component of economic expectation sensitivity," but a shorter expression would seem more appropriate. Such a notion could be termed an "attractor" when considering a problem. The literature dedicated to this issue outlines a distinctive feature of an attractor as its weak dependence on the starting state of the causing system. According to the theory of dynamic system, "an attractor is a set of numerical properties towards which a system evolves from a variety of initial conditions." (Beven and Davies, 2015, p. 5215); this property is discussed also in Milnor (2006), Hoeksma *et al*. (2007), Broer and Takens (2008), Ruelle (1981), Abraham and Shaw (1992), Goldstein (2011).

The quality of the attraction pool is assignable to those data points which, although quantitatively different, are structurally related to the series inducing the given attractor. The probability of finding such samples within the effective statistics is very low. Artificial data generating algorithms should be considered cautiously, particularly for series such as mtz and sfd, which are indubitably marked by "time's arrow." Therefore, a simple random mixture of the sample data would be highly questionable. According to McIntosh (2016), "Jackknife is inappropriate for correlated data or time series data (p. 6)." That is why, the y@ values (deduced according to the above presented methodology) are defined as attractors in a narrow sense, as expressions of a presumptive state of the economy if the current dominant market behaviors do not change.

2.2 Database and Modelling Design

The database is concentrated in the Supplementary Data SA, which contains the quarterly estimates for 1995Q1–2020Q4 of all the European Union countries (ante-Brexit format) concerning: i) economic sentiment index (ESIq); ii) gross domestic product index in real terms (IGDPq); iii) expectation sensitivity (mtzq); and iv) self-fulfilment degree of expectations (sfdq). Regarding ESIq, there are data for the entire sample (104 observations) of 23 countries, with shorter series being recorded in the case of Slovenia (103), Hungary

(100), Cyprus (78), Malta (72), Croatia (50). Only five countries have full samples of IGDPq, while the majority of the other (20) register 103 observations; shorter series appear for Netherlands (99), Czechia (99), and Malta (83). The longest series of mtzq comprises 102 terms (22 countries), with the following countries being relatively less covered from this viewpoint: Hungary (100), Netherland (98), Czechia (98), Cyprus (78), Malta (72), and Croatia (50). Concerning the sfdq, presently five countries (Germany, Finland, France, Sweden, United Kingdom) have complete series (104 observations) and other 17 countries have 103 observations; similar to mtzq, some shorter sfdq series are registered in the case of Hungary (100), Netherland (99), Czechia (99), Cyprus (78), Malta (72), Croatia (50).

For each indicator, the matrix of data contains 2912 cells (104x28); 95.98% are complete for ESIq; 98.25% for IGDPq; 94.27% for mtzq, and 95.09% for sfdq. Thus, the overall statistical support of our empirical research can be admitted as relevant.

The descriptive statistics reveal important inter-country differences concerning the temporal distributions of the observations.

The ESIq series (Appendix 1a) exhibit high oscillations of data, which are clustered into three groups. Seven countries are characterized by a coefficient of variation (CV) between 0.09 and 0.1. Half of the sample is placed within the median group, where CV reaches 0.1–0.11. The rest od countries register 0.11-0.146.

An opposite picture is offered by the IGDPq series with significantly lower coefficients of variation: for six countries the CV does not exceed 0.015; ten of them reach 0.015–0.02, and the remaining 13 surpass this limit (0.02–0.035).

As a result of the division of ESIq by IGDPq(-1), mtzq is characterized by a coefficient of variation that is nearer to that recorded by ESIq. The volatility of the expectation sensitivity within the European Union is predominantly linked with the dynamics of the expectational processes themselves. Thus, the higher coefficient of variation between sfdq and mtzq should also be noticed.

All statistical series were submitted to unit root tests in level. A large battery of procedures was involved: Augmented Dickey-Fuller (ADF; Fuller, 1996; Davidson, 2004; MacKinnon, 2010; Enders and Liu, 2014); Phillips-Perron (PP; Phillips and Perron, 1988; Leybourne and Newbold, 1999; Davidson, 2004); Elliott-Rothenberg-Stock (ERS DF-GLS; Elliott *et al*., 1996; Lopez, 2003; Davidson, 2004; Baum *et al*., 2017); Kwiatkowski-Phillips-Schmidt-Shin (KPSS; Kwiatkowski *et al*., 1992; Dongin and Schmidt, 1996; Syczewska, 2010); Elliott-Rothenberg-Stock (ERS; Elliott *et al.*, 1996; MacKinnon, 2010; Otero and Baum, 2017); Ng-Perron (NgP; Ng and Perron, 1995, 2001); Breakpoint (BP; Vogelsang and Perron, 1989; Perron and Qu, 2006; Vogelsang and Perron, 1998).

In the first instance, they were admitted as stationary for all the series in which Augmented Dickey-Fuller (ADF) or Phillips-Perron (PP) reject the null hypothesis “series has a unit root” (probability <0.05). When both ADF and PP did not observe this convention, the structural stability of the respective series was evaluated taking into consideration the results provided by the rest of the above-mentioned tests, including Ng-Perron in four variants: MZa, MZt, MSB, MPT.

Regarding ESIq, Appendix 2a synthetizes the ADF and PP tests indicating that a null hypothesis probability lies below 0.05, together or separately. The stationary nature is attested for 13 series by both ADF and PP tests, while 6 series are attested by one of them. Besides, this property is revealed by other URT tests in the following 8 cases. Only one series (HR_ESIq) did not respond positively to the ten identified stationarity procedures; this series is significantly shorter than the others (50 against the standard of 104 observations).

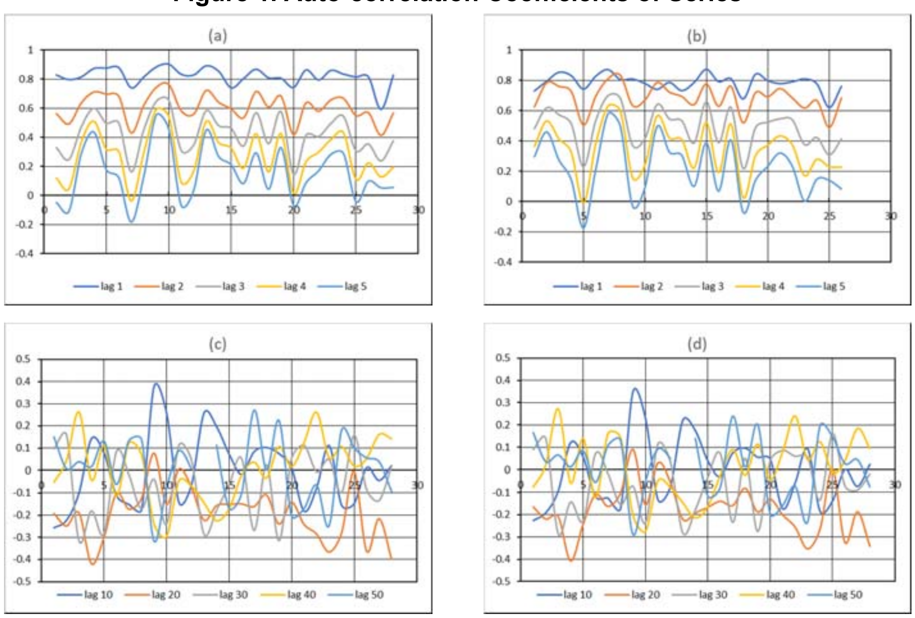
In the case of IGDPq, the URT null hypothesis did not surpass the threshold of 0.05 in both the ADF and PP algorithms in 27 cases; only the ES_IGDPq series observes this standard for the PP test, but not for the ADF test (Appendix 2b).

Naturally, the stationarity of the mtzq data is marked by the corresponding characteristics of the basic ESIq and IGDOq series (Appendix 2c). Both the ADF and PP tests confirmed the stationarity of 19 series, while at least one of them confirmed four more. Three others were attested by other procedures: CY_mtzq (by KPSS, NgP MSB variant); EL_mtzq (by ERS DF-GLS, ERS, NgP all four variants), and HU_mtzq (by ERS DF-GLS, KPSS, ERS, NgP all four variants; in this case even ADF showed a null hypothesis probability of 0.0567 and PP of 0.0502), and LT_mtzq (by ERS DF-GLS, KPSS, ERS, Ng-P all four variants). Stationarity was rejected by all ten processed tests only in the case of HR_mtzq.

Concerning sfdq (Appendix 2d), the stationarity of 20 series was validated by both ADF and PP tests, while four series were validated by at least one of these procedures. Three series did not pass the ADF or PP tests, but were accepted by other techniques — CY_sfdq (KPSS), EL_sfdq (ERS DF-GLS, Ng-P four variants) and ES_sfdq (ERS DF-GLS, KPSS, Ng-P variant MZa). Only HR_sfdq was rejected by all the applied unit root tests.

The correlograms of mtzq and sfdq (auto-correlation — AC; and partial auto-correlation — PAC coefficients) computed for the entire available samples are presented in Supplementary Data SAC1. Figure 1 describes the evolution of the auto-correlation coefficient.

Figure 1. Auto-correlation Coefficients of Series



Note: a. Auto-correlation coefficient of mtzq series, lags 1–5; b. Auto-correlation coefficient of sfdq series, lags 1–5; c. Auto-correlation coefficient of mtzq series, lags 10–50; d. Auto-correlation coefficient of sfdq series, lags 10–50.

Figure 1 shows that the auto-correlation coefficients of the mtzq and sfdq series, although failing with an increasing number of lags, remain non-negligible. Therefore, the lasting presence of auto-correlation should not be a surprise, since as a rule the new expectations always start from the anterior state of the envisaged deal, implicitly from the corresponding precedent expectations. Due to this characteristic, mtzq and sfdq series are suitable for the AR modelling techniques.

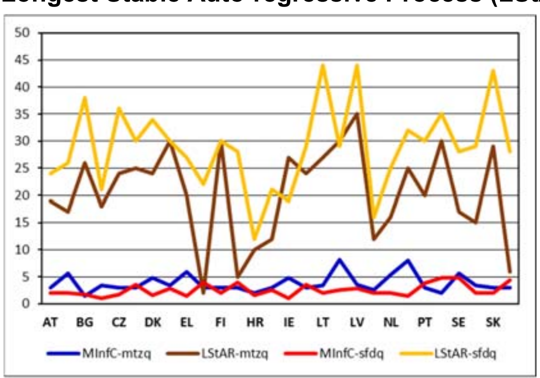
The lag-length of the auto-regressive process is a key issue, with two targets being essential for its determination. One comes from the pronounced persistence of serial correlations observed in nearly all the analyzed series. This would impose extended lag-lengths that are able to reflect the statistical properties of a greater part of the sample. On the other hand, the chosen specification allows to clarify if the examined time series are convergent or not.

The informational criteria are not a satisfactory answer to the mentioned targets as they are centered on the minimization of regression squared residuals alone. As a proof, the mtzq and sfdq series were submitted to five such tests: sequential modified likelihood ratio – LR (Lütkepohl, 2005; Hatemi-J and Hacker, 2009); final prediction error – FPE (Ljung, 1999); Akaike information criterion – AIC (Akaike, 1969 and 1973); Schwarz information criterion – SC (Schwarz, 1978); and Hannan-Quinn information criterion – HQ (Hannan and Quinn, 1979). Useful comments about these tests can also be found in Liew (2004) and Zahid and Irum (2007). Tables 1a and 1b from the Appendix [2] contain the number of lags indicated by the tests previously noted as optimal for the mtzq and sfdq series.

Therefore, except for the likelihood ratio (LR), the informational criteria recommend short lag-lengths. It is hard to say if such lengths can reveal whether the mtzq and sfdq series are convergent or not. They might accidentally converge as the informational criteria were not designed for such a scenario.

A more appropriate way to estimate the AR lag-length, in accordance with the previously noted targets, is the LStAR methodology. Figure 2 presents the results of using this methodology, compared with the corresponding means of lag-lengths indicated by the informational criteria (MInfC).

Figure 2. Mean Lag-length Indicated by the Informational Criteria (MInfC) and by the Longest Stable Auto-regressive Process (LStAR)



[2] *Available online at https://ipe.ro/rjef.htm.*

The LStAR lag-lengths are longer in all the cases — sometimes considerably — than the averages stipulated by the informational criteria. Thus, these estimations satisfy the stability condition.

In order to maximally valorise the statistical properties of the available data, the LStAR lag-lengths have been selected to model the mtzq and sfdq series as lag distributed auto-regressive processes. Adapting formula (7) to the mtzq and sfdq series, the operational LStAR models are defined as:

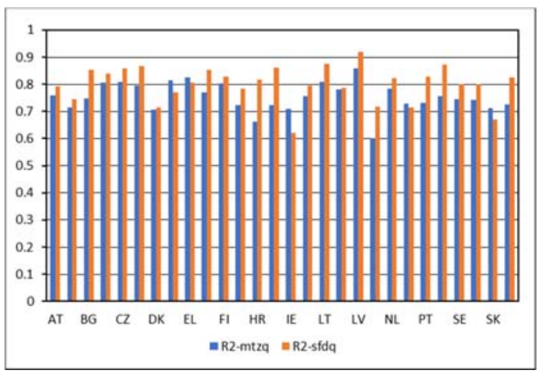
$$mtzq_{kt} = c_{k0m} + \Sigma c_{kjm} * mtzq_{k(t\text{-}j)} + u_{ktm} \quad (11)$$

$$sfdq_{kt} = c_{k0s} + \Sigma c_{kjs} * sfdq_{k(t\text{-}j)} + u_{kts} \quad (11a)$$

where: k is the country code; j=1, 2… p, and m and s indicate the modelled series.

Estimators for the (11) and (11a) specifications are detailed for all the EU countries in Supplementary Data SM. The accuracy with which the sample data are approximated by these models is presented in the following figure.

Figure 3. Goodness of Fit of the mtzq and sfdq AR-models

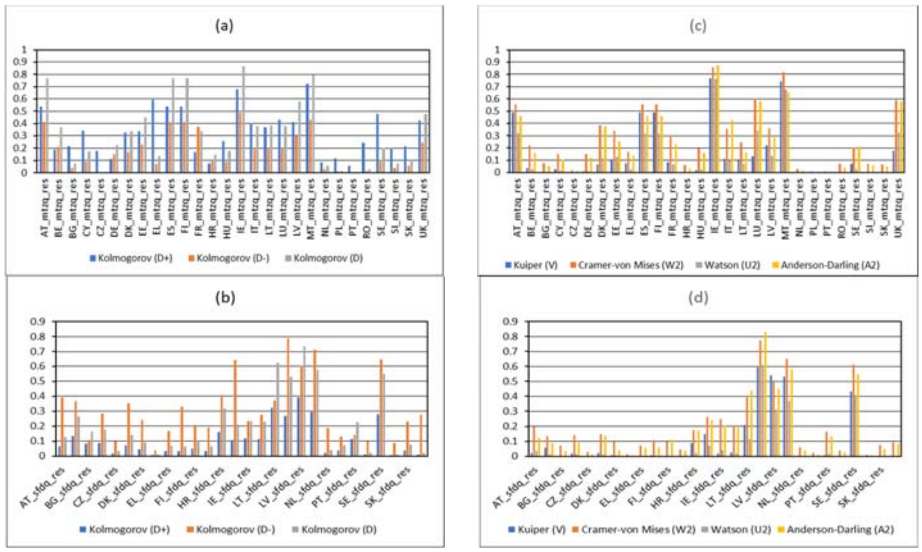


It is worth noticing that in both series of models, the coefficients of determination are ranged at acceptable levels. The econometric strength of these results was checked by residual tests, paying special attention to their normality, homoscedasticity, and serial autocorrelation. The residuals were applied to the empirical distribution tests (null hypothesis: normal) and to the long-run normality test.

The empirical distribution tests were used in the variants D+, D-, and D of the Kolmogorov procedure (Conover, 1972), and in other four variants: Kuiper V (Kuiper, 1960), Cramer-von Mises W^2 (Cramér, 1928; Anderson, 1962), Watson U^2 (Persson, 1979), and Anderson-Darling A^2 (Anderson and Darling, 1952). The conceptual framework and applicative problems of these tests were largely debated in the statistical literature by Anderson and Darling (1954), Lilliefors (1967), Stephens (1974), Dallal and Wilkinson (1986), Marsaglia *et al*. (2003), Hanusz and Tarasińska (2015), and Mishra *et al*. (2019).

The results obtained for the mtzq_res series using the Kolmogorov procedures are presented in Figure 4a, while the other methods are shown in Figure 4c.

Figure 4. Empirical Distribution Test – Probability of the Null Hypothesis (Normal)



Therefore, for 22 countries, the normal distribution was attested (probability of the null hypothesis higher than 0.05) by 7-4 tests. The normality was also recognized by three other tests for Bulgaria, two for Romania, and only one for Czechia, Netherland, Poland, and Portugal. It is worth remembering that the Kolmogorov (D+) test accepted the null hypothesis for all the countries.

The same tests applied to the sfdq_res series are plotted in Figures 4b and 4d. The normality of a large part of sfdq_res (18) was also validated by 4-7 tests, with the remaining nine such series registering fewer confirmations: three for Denmark, Slovakia, and the United Kingdom; two for France, Netherland, and Poland; and one for Czechia, Romania, and Slovenia. The residuals of Estonia were the only model to fail all the above-mentioned procedures.

Overall, the empirical distribution tests do not reject the estimated models (11) and (11a) from the "normality viewpoint". The long-run normality test (Bai and Ng, 2005) enforces this statement. Table 2 from the Appendix contains the results of this test.

The lowest probability of the long-run normality test is recorded by Greece (slightly more than 8%) for the sfdq_res series; the distribution of the remaining 55 tests is as follows: 8 between 10% and 20%, 15 between 20% and 40%, and 32 tests between 40% and 60%. The normality of the residuals of models (11) and (11a) is thus clearly corroborated.

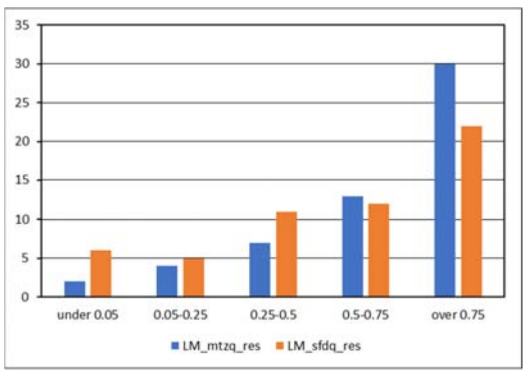
The possible heteroscedasticity of the residuals was checked using the Breusch-Pagan-Godfrey tests (Godfrey, 1978; Breusch and Pagan, 1979; Harvey, 1976 and Glejser, 1969). The test was computed separately for the F-statistic, Obs*R-squared, and Scaled explained SS, thus obtaining nine probabilities of homoscedasticity for each series of residuals. It should be noticed that out of the 252 null hypothesis probabilities for each residuals series, those that are 0.05 and higher represent approximately 83% of the mtzq_res and 90% of the sfdq_res series. The distribution of residuals confirming homoscedasticity probabilities is displayed in Table 3 from the Appendix 1.

There are very few cases where homoscedasticity is identified only by two tests; these include Estonia and the United Kingdom from the mtzq_res series, and Ireland from the sfdq_series; the number confirming the null hypothesis tests exceeds 5 for all the other countries.

The VAR residual serial Correlation LM tests were applied in the version developed by Breusch (1978), Godfrey (1978 and 1988), Davidson and MacKinnon (1993), and Doornik (1996). For each country-model, the residuals leading to a lag-length of (p+1) were checked. When this option technically failed, a higher number of lags was adopted for which the test worked. This was applicable to one country from the mtzq_res series (Latvia) and seven from the sfdq_res series (Bulgaria, Czechia, Denmark, Lithuania, Latvia, Romania, and Slovenia). As discriminating criteria, the LRE* statistic (*Edgeworth expansion corrected likelihood ratio statistic) and Rao F-statistic were adopted, which provided very close results. Probabilities were computed for the "no serial correlation at lags 1 to h" and are presented in Table 4 in the Appendix 1.

The H_0 probabilities compacted into a five-threshold classification (0–0.05; 0.05–0.25; 0.25–0.5; 0.5–0.75, and over 0.75) and described in the previous table, are presented in Figure 5.

Figure 5. VAR Residual Serial Correlation LM by Country Models. Distribution of the Null Hypothesis Probability of LRE*-stat. and Rao T-stat



Almost 93% of the processed tests rejected the residual serial correlation with a probability over 5%. However, there are six cases (ES_mtzq_res, BG_sfdq_res, CZ_sfdq_res, DK_sfdq_res, ES_sfdq_res, and LV_sfdq_res) with probabilities lower than this critical level. Additional correlograms were estimated for these cases (see Supplementary Data SCA2). For all these series, the Ljung-Box Q-statistics inform the presence of residual serial correlation at all computed lags with significant p-values, except for ES_mtzq_res, which appears more visible at a higher number of lags.

Therefore, the residual analysis revealed a satisfactory econometric validation of the mtzq and sfdq country-models, building trust in the following post-sample simulations

3. Post-sample Simulations

The main purpose of the post-sample simulations is to establish whether the series for the sensitivity of economic expectations and their degree of self-fulfilment, at least sub-textually, converge towards certain steady states within the European Union. They consist of solving each country's mtzq and sfdq models repeatedly and observing whether the obtained successive estimations are convergent or not.

The computational exercise is structured into two phases. First, the post-sample iterations are continued until the mtzq and sfdq results become compatible with restriction (10); the values observing it are denoted by smtzq and ssfdq. Secondly, the first hundred estimations of smtzq and ssfdq are used to approximate the attractors. Depending on the adopted lag-length interval (p) and on the statistical peculiarities of the involved samples, the number of simulations differs across countries (Table 1).

Table 1. Total Number of Simulations Operated on the mtzq and sfdq Models by Country

Country	mtzq models	sfdq models	Country	mtzq models	sfdq models	Country	mtzq models	sfdq models
AT	410	4138	ES	240	40459	LV	3282	4268
BE	1725	1034	FI	568	3278	MT	254	660
BG	3670	1849	FR	840	935	NL	415	4745
CY	7321	3830	HR	9736	520	PL	2834	25335
CZ	3343	4589	HU	46868	1003	PT	451	1553
DE	12698	5199	IE	3719	713	RO	622	851
DK	7662	2619	IT	3335	1250	SE	605	719
EE	918	2088	LT	1809	2479	SI	397	2423
EL	2580	2119	LU	8934	525	SK	2650	3572
						UK	134	1818

For the mtzq models, the necessary simulation operations vary from only 134 (United Kingdom) to 46,868 (Hungary), while for sfdq, they vary from 520 iterations (Croatia) to 40,459 (Estonia) iterations. The lengths of the post-sample simulations are also significantly different for the mtzq and sfdq models within the same country. The key results of the simulations are presented graphically in the following figures.

Figures 6a and 6b display the statistics and the 50 post-sample simulations for each series of mtzq and sfdq; they disclose how the original data configuration impacts the pattern of the subsequent simulations.

Figure 6a: The mtzq Statistical Series Prolonged by 50 Post-sample Simulations

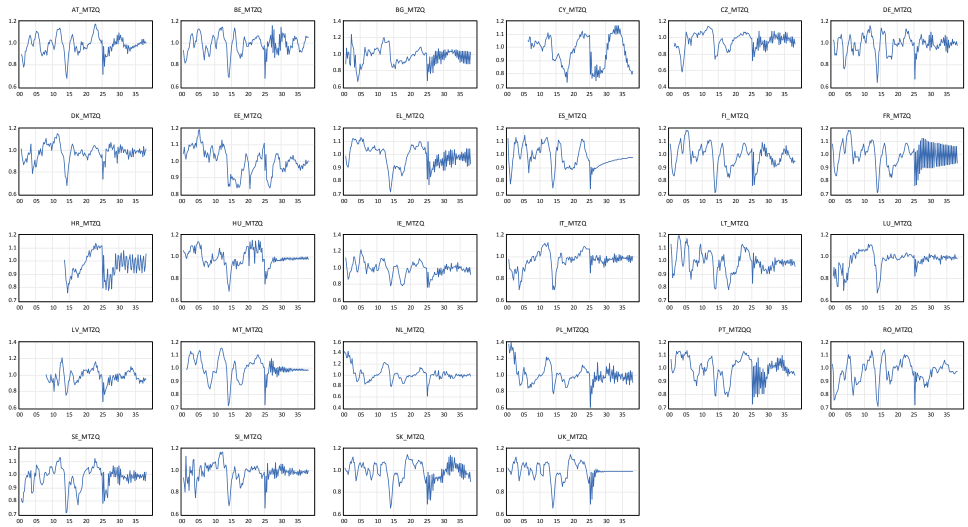
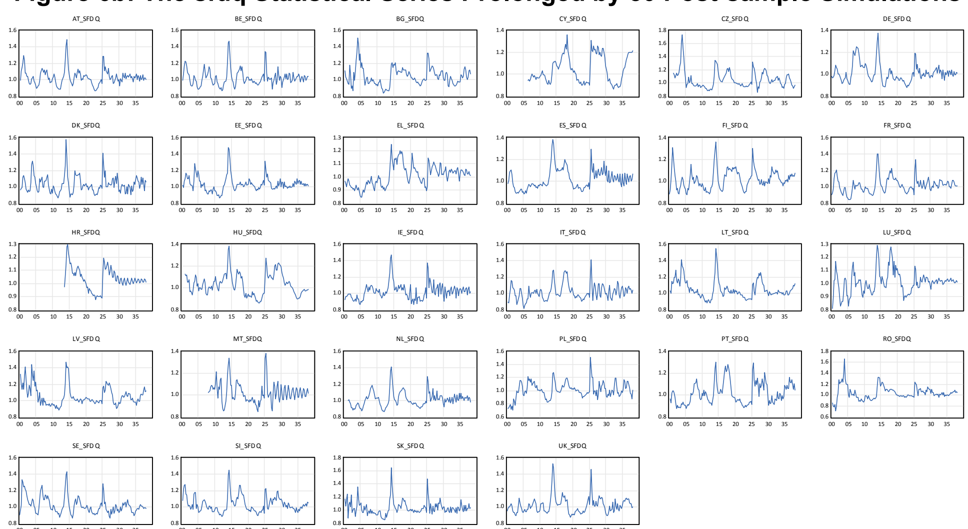


Figure 6b: The sfdq Statistical Series Prolonged by 50 Post-sample Simulations



Therefore, successive iterations most frequently describe an irregularly oscillating pattern.

The post-sample iterations are carried on until the mtzq and sfdq computational results become compatible with restriction (□); the values observing it are denoted by smtzq and ssfdq. Figures 6c and 6d display the first hundred smtzq and ssdfq estimations.

Figure 6c: First Hundred smtzq Series

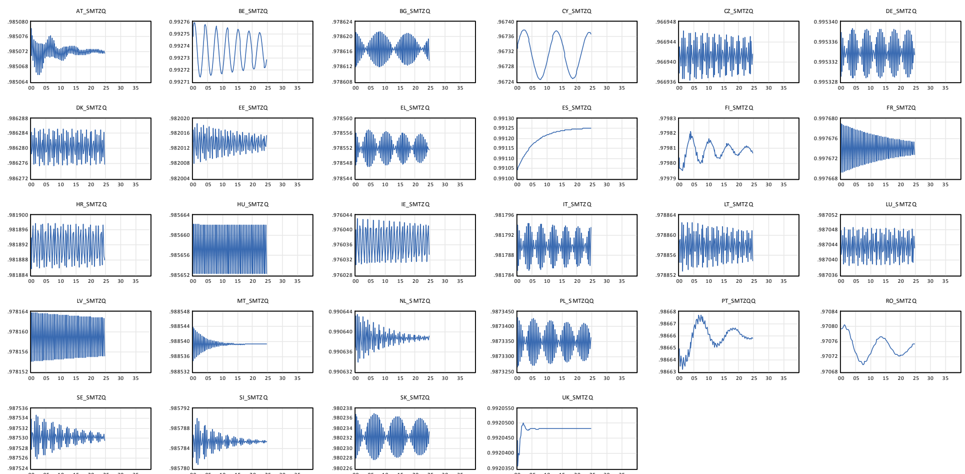
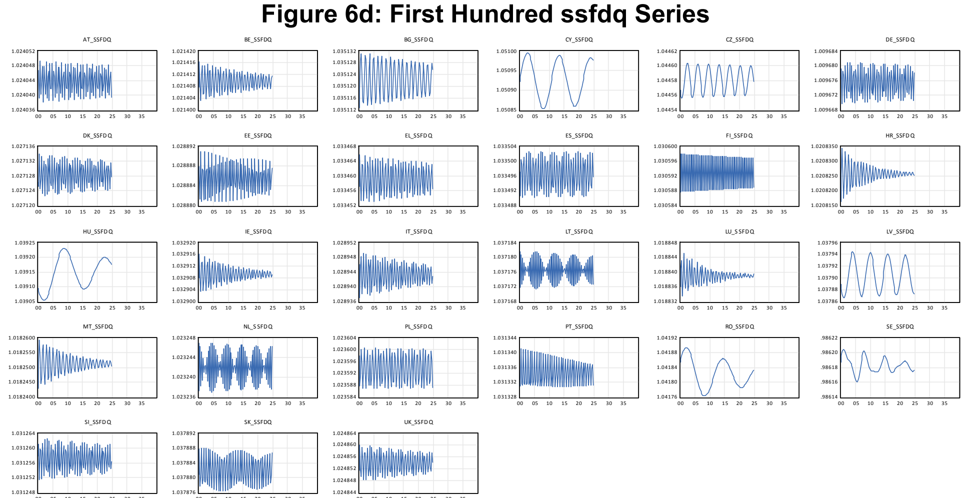


Figure 6d: First Hundred ssfdq Series



Figures 6c and 6d are contrasted by their dynamic regularity with the pictures described in Figures 6a and 6b.

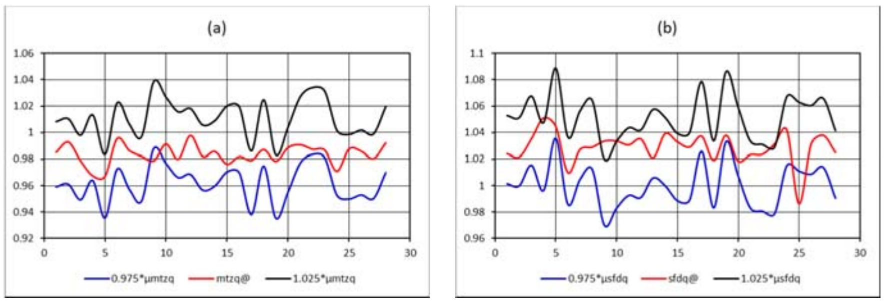
The post-sample simulations are completed by applying the averaging operator on the series given in Figure 6c and 6d to the approximation of the attractor levels (mtzq@ and sfdq@). The approximated attractors are detailed in Table 2.

Table 2. Estimated mtzq@ and sfdq@ Attractors

Country	mtzq@	sfdq@	Country	mtzq@	sfdq@	Country	mtzq@	sfdq@
AT_mtzq	0.98507201	1.024043694	ES_mtzq	0.991204414	1.033496366	LV_mtzq	0.978158923	1.037903959
BE_mtzq	0.99273617	1.021409607	FI_mtzq	0.979808124	1.030592795	MT_mtzq	0.98853934	1.018251147
BG_mtzq	0.97861661	1.035122219	FR_mtzq	0.997674015	1.035082052	NL_mtzq	0.990638726	1.023242013
CY_mtzq	0.96731757	1.050929016	HR_mtzq	0.981891737	1.020825562	PL_mtzq	0.987334766	1.023593572
CZ_mtzq	0.96694115	1.044579466	HU_mtzq	0.985657163	1.039147115	PT_mtzq	0.986658471	1.031334519
DE_mtzq	0.9953336	1.009675404	IE_mtzq	0.976036707	1.032909246	RO_mtzq	0.970745667	1.041823567
DK_mtzq	0.98628029	1.027128162	IT_mtzq	0.981789647	1.028943253	SE_mtzq	0.987530232	0.986182997
EE_mtzq	0.98201339	1.0288855	LT_mtzq	0.978857754	1.037176473	SI_mtzq	0.985785277	1.031257028
EL_mtzq	0.97855192	1.033459113	LU_mtzq	0.987043597	1.018838909	SK_mtzq	0.980232289	1.037882004
						UK_mtzq	0.992047737	1.024853905

The closeness of all the attractors to the corresponding sample means is remarkable. (Figures 7a and 7b).

Figure 7. Comparison of Attractors with Corresponding Sample-means



Note: (a) Comparison of mtzq attractors (mtzq@) with corresponding sample-means (μmtzq);
(b) Comparison of sfdq attractors (sfdq@) with corresponding sample-means (μsfdq).

Except for a few deviations, the attractors of the EU countries are placed within the band of ±2.5% around the sample mean. This can be considered as a notable confirmation of the stability of the VAR condition.

The sensitivity of expectations and their degree of self-fullfilment are intimately linked by the factor referred to as the expectational impulse (eimq); this is interpreted as the modification of the speed or direction registered by the process as a result of interactions of the corresponding expectations with reality. Applied to the attractor values of mtzq and sfdq (estimated for database of the European Union), Table 3 presents the values of eimq@ by country.

Table 3. Expectational Impulse (eimq@) Conforming to the mtzq@ and sfdq@ Attractors by Country

Country	mtzq@	sfdq@	eimq@	Country	mtzq@	sfdq@	eimq@
AT	0.985072	1.024044	0.008757	IE	0.976037	1.032909	0.008157
BE	0.992736	1.02141	0.01399	IT	0.98179	1.028943	0.010206
BG	0.978617	1.035122	0.012988	LT	0.978858	1.037176	0.015248
CY	0.967318	1.050929	0.016582	LU	0.987044	1.018839	0.005638
CZ	0.966941	1.044579	0.010047	LV	0.978159	1.037904	0.015235
DE	0.995334	1.009675	0.004964	MT	0.988539	1.018251	0.006581
DK	0.98628	1.027128	0.013036	NL	0.990639	1.023242	0.013663
EE	0.982013	1.028886	0.010379	PL	0.987335	1.023594	0.01063
EL	0.978552	1.033459	0.011293	PT	0.986658	1.031335	0.017575
ES	0.991204	1.033496	0.024406	RO	0.970746	1.041824	0.011346
FI	0.979808	1.030593	0.009783	SE	0.98753	0.986183	-0.02611
FR	0.997674	1.035082	0.032674	SI	0.985785	1.031257	0.016598
HR	0.981892	1.020826	0.00234	SK	0.980232	1.037882	0.017365
HU	0.985657	1.039147	0.024243	UK	0.992048	1.024854	0.016704

Therefore, from the viewpoint of the expectational impulse leading to the estimated attractors (of the sensitivity and self-fullfilment of economic expectations) within the European Union, five classes can be identified — negative (eimq@<0); relatively stable (eimq@ between 0 and 0.01); slightly positive (eimq@ between 0.01 and 0.015); moderately positive (eimq@ between 0.015 and 0.02); and accentuated positive (eimq@ over 0.02). Only Sweden belongs to the first class. Seven countries (Croatia, Germany, Luxemburg, Malta, Ireland, Austria, and Finland) are found in the second class, while the third category is the most populated (Czechia, Italy, Estonia, Poland, Greece, Romania, Bulgaria, Denmark, Netherland, and Belgium). The fourth class includes seven countries (Latvia, Lithuania, Cyprus, Slovenia, the United Kingdom, Slovakia, and Portugal), while the last one contains three countries (Hungary, Spain, and France). As they reveal the deep state of the operators' market perceptions, expectational drifts can represent useful insights for the macroeconomic management.

4. Conclusions

Our research report is a continuation of the debates over an extremely important issue: how the expectations of market operators act as the core of their decisional behavior and, implicitly, of the entire economic life mechanism. Our main objective was to identify whether the steady-state theorem is applicable to this modern topic by focusing on the reactivity of the expectations against the economy's factual dynamics and the extent to which these expectations are translated into reality. Unlike many previous studies that focused on the specificity of such phenomena in different business segments — in the context of prices, interest rate, investment, inventories — we analyzed its configuration on macroeconomic grounds. We started from the statement that despite the heterogeneity of expectation forming mechanisms within the microenvironment (theoretically generalized as different random-walk variants, a large series of auto-corrective algorithms, and rational paradigm), expectations manifested at the macro level as a congruent enough economic zeitgeist. This coagulating impact predominantly arises from pre-contractual communications between market operators as well as the macroeconomic guidance exerted by governments, central banks, and international financial and commercial organizations. Two statistical aggregates allow us to examine the "real economy-expectations binomial" at a macro-level: the gross domestic product and the economic sentiment index.

Quantitatively, the sensitivity of expectations can be measured by the first order differences of the involved indicators (early Metzler definition), or by using the levels of these indicators. We preferred the second solution (named mtz), which usually generates positive time series of lower standard deviations, and consequently, more robust macroeconomic estimates of the sensitivity of expectations. Measuring the degree of self-fulfilment of expectations (sfd) by the ratio of the realized indicators to the expected ones, proved to have good modelling tractability.

The quarterly series of the entire European Union in pre-Brexit format (28 countries) for the years 1995–2020 ensured a solid informational foundation for the analysis. Not only was the dimension of this database important, but the diversity of the national economies studied was also significant (in the context of the development stage, sectoral structure, cultural traditions, institutional and behavioral peculiarities, etc.). Both these circumstances contributed to increasing the generality of the resulting conclusions. All the involved time series are characterized by a pronounced temporal auto-correlation, which statistically confirms the strong interdependence between the real economic processes and the corresponding expectations of the market operators.

A key issue in such a modelling approach is determining the optimal length of the AR process. From this viewpoint, our application was confronted with a double challenge. On one hand, the interpretation of the current expectations in connection with previous experience would imply a greater number of lags in the regression. This way, the past behavioral features of the expectation forming mechanisms would be counted most comprehensively. On the other hand, the study's objective is to pinpoint potential steady-states in the available data. To achieve this, two technical procedures could be considered: i) a unit root test for the stationarity of mtzq and sfdq series in terms of their levels; ii) the VAR stability condition, meaning that no root of the corresponding characteristic polynomial lies outside the unit circle. The second criterion proved to be more relevant, at least in the present application.

As previously shown, the largest part of the mtzq series is I(0). There are, nevertheless, two series (for Cyprus and Croatia) where the stationarity of the levels could hardly be established. For instance, in the case of CY_mtzq, eight tests (ADF, PP, ERS as such and in the DF-GLS version, three variants Ng-P, and BP) contested it, with only KPSS and one Ng-P variant being favorable. In the case of HR, all the ten applied tests rejected stationarity. Instead, CY_mtzq conforms to the VAR stability condition for 17 lag-lengths (from 2 to 18), and HR for 9 (from 2 to 10); both series generate steady states.

Consequently, the study admitted the maximal lag-length which observes the AR stability condition (LStAR) as optimal. All 28 AR models built on this principle successfully passed the residual tests for the normality distribution, serial correlation, and homoscedasticity. A question that remains to be studied concerns the priority of VAR stability condition against usual I(0) tests and if it is a peculiarity of examined database or a more general property of time series.

The most important finding of this study is that when submitted to successive post-sample simulations, all the experimented models generated steady state type estimations. The numerical verification of this hypothesis was guided by the exogenous restrictions Λ=0.00001 and ⲫ=100. These are an ad-hoc solution for the present application, but it should be noted that their eventual modification (within reasonable borders) would not switch the estimations of the attractors. As an illustration (see Supplementary Data SF), simulations were operated on FR-mtzq under constant Λ (at 0.00001) and variable ⲫ (between 50 and 4000). The resultant attractors remain in the proximity of the basic variant; the results were similar when fixed ⲫ (at 100) was combined with changing Λ (from 0.00005 to 0.0000001). This example shows that the study's option concerning Λ and ⲫ can be admitted as realistic enough for the analyzed database.

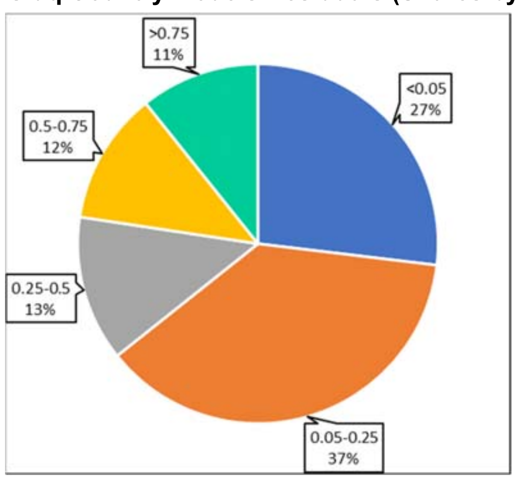
The attractor notion assigned to the steady-states revealed by the study's proposed modelling approach should be understood in a narrow, special sense. It is not a normatively desirable measure of the sensitivity of economic expectations or degree of self-fulfilment of expectations. Rather, it is about their presumptive levels concerning the actual dominant market behaviors being indefinitely perpetuated. In other words, the estimates of this study's attractors do not represent anything more than signals concerning the sensitivity of economic expectations and their degree of self-fulfilment under the conditions of a completely static society. Consequently, they must be considered with caution for the purposes of forecasting and only used for short-run predictions. Analyzing the expectational processes at the macro-level is still in a nascent stage; our analysis represents a very small step ahead in this field. Some issues remain to be addressed. We draw attention to two of them.

The first concerns the modelling specifications. As a starting attempt to approximate some attractors of the economic expectational mechanism, the present study used linear AR estimations. The VAR residual serial correlation LM tests were considered relevant. However, the involvement of more powerful tools to check the interdependencies among model residuals reveals some problems.

The BDS test (Broock *et al*., 1996) is such a technique (H_0: residuals are independently and identically distributed), as it can not only identify linear deviations of the tested series from independence, but also nonlinear deviations, or those associated with chaos-states. In our application, it was applied to five dimensions (2, 3, 4, 5, and 6) and three options concerning the testing of proximity of data points (fraction of pairs, standard deviations, and fraction of range). The probability of the null hypothesis was not only calculated for the sample data, but also for the bootstrapped series, whose results were counted as more dependable. The results of the latter are presented below.

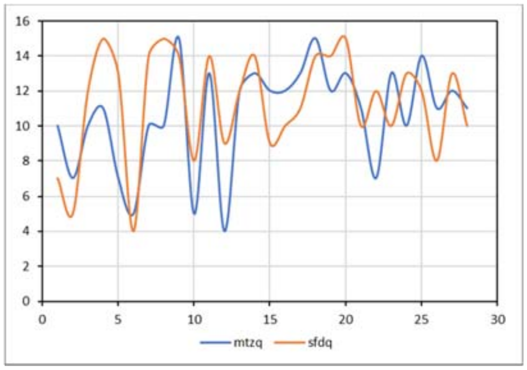
There were 15 bootstrapped BDS H_0 probabilities obtained for each mtzq and sfdq country-model; that is 840 estimations in all. These were grouped depending on the following thresholds: <0.05; 0.05–0.25: 0.25–0.5; 0.5–0.75; >0.75. The resultant shares are plotted in Figure 8.

Figure 8. Distribution of All BDS Bootstrapped Null Hypothesis Probabilities for mtzq and sfdq Country Models Residuals (Shares by Groups)



It appears that more than 73% of the probabilities of the bootstrapped BDS tests surpass the referential border of 0.05, which in principle is conclusive to validate the tested models. Nevertheless, the share of tests rejecting the BDS null hypothesis cannot be overlooked. The picture is more nuanced when organized by countries (Figure 9).

Figure 9. BDS Tests with the Null Hypothesis Probability over 0.05 (Country Models)



There are several country-models where more than half of the BDS residuals tests do not confirm the null hypothesis: Belgium, Czechia, Germany, Estonia, France, and Poland in the case of mtzq models, and Austria, Belgium, and Germany in the case of sfdq models. In such situations it would be worth trying nonlinear VAR adaptations.

Coming back to Table 3, the attractors for the degree of self-fulfilment of expectations exceed unity in most cases. Only Sweden's level is slightly inferior, while that of Germany is close to unity. The numerator of sfdq is represented by the effective data (IGDPq); therefore, the source of its higher-than-unity value must be examined in the denominator of the fraction — ESI, which could be under-sized from the very beginning. For example, we look at the case of the United Kingdom, with the estimated attractor for sfdq at 1.024854. Initial data for the economic sentiment index were multiplied by this value, correspondingly modifying the mtzq and sfdq series. Applying the LStAR methodology on the corrected data, a mtzq attractor of 1.016704 and a sfdq attractor equal to unity were obtained. The new resultant expectational impulse (1.016704*1-1) does not differ from that mentioned in Table 3 for the United Kingdom (0.016704).

The IGDPq>ESI inequality may result from many circumstances. Possible causes include a prolonged and chronically pessimistic economic zeitgeist, deficiencies in the building of survey questionnaires or in data collection. Some statistical methodologies and aggregative operations may also contribute to such a perturbation. Further research must be conducted to elucidate the sources of the signalled phenomenon.

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