WHAT MOVES HEALTH INSURANCE IN BOSNIA AND HERZEGOVINA? AN ARDL HEALTH INSURANCE PREMIUMS STUDY

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Abstract
This study examines how macroeconomic conditions and demographic change shape the highfrequency dynamics of Bosnia and Herzegovina's health insurance market. The focus is on quarterly premium flows for health lines. The empirical design combines quarterly health insurance premium flows, along with macroeconomic and demographic factors, which, according to the literature, shape insurance and its premium flows. Annual variables are disaggregated to quarterly frequency using the Chow-Lin disaggregation method; DentonCholette is used as a robustness alternative, where no adequate driver is found. After stationarity diagnostics, and dimensionality reduction through principal component analysis, the Autoregressive Distributed Lag (ARDL) bounds-testing approach to cointegration with an associated error-correction form is employed, to quantify the long- and short-run links between premiums/claims and macro-demographic drivers. Data on quarterly premiums are compiled from national insurance market reports, while other variables are taken from aggregated data banks, comprising datasets created by the International Monetary Fund, World Bank, and World Health Organization. The observed data range is from 2011 to 2023. The study is implemented at the national level with a goal of providing a compact, data-driven assessment of how growth, inflation, and population factors move together with health insurance flows in Bosnia and Herzegovina.
Keywords: Bosnia and Herzegovina; health insurance; ARDL bounds cointegration testing; PCA; macroeconomic and fiscal factors; demographic factors
JEL: C32, C38, I13, H51, O52

1. INTRODUCTION

Over the last two decades, global health financing systems have undergone significant transformation, marked by rising health expenditures, demographic transitions, and the diversification of financing sources. According to the OECD, health spending has increased both in absolute terms and as a share of GDP, driven mostly by population ageing, technological change, and demand for higher quality care (OECD, 2023). In more developed and high-income countries, private health insurance often complements mandatory public schemes by covering extra services or reducing out-of-pocket expenditure (OOP) stress and costs. In contrast, in low- and middle-income economies, private health insurance plays a very limited role, but surprisingly enough, a growing one. However, as per World Health Organization (2022), such insurance unfortunately still dominates across high-income societal groups. These global patterns underline the importance of distinguishing between mandatory contribution-based systems and voluntary premium-based systems when analysing these dynamics.

Health insurance plays a dual role in transition economies. Mandatory schemes dominate across coverage, private health insurance, and health insurance outside of the mandatory framework in general, still represents a smaller but dynamic market segment. In Bosnia and Herzegovina (BiH), mandatory health insurance is publicly organised and covers the majority of the population, but its premium and claims data are neither transparent nor available directly, and at high frequency. On the other hand, health insurance outside of these flows, though modest in size, provides premium data through insurance companies regulated by the entity agencies[25], and overseen by the Agency of Supervision of Insurance (AZOBIH).

This paper therefore focuses on health insurance premiums outside of the mandatory framework (HIPREM), and thus uses them as a proxy for high-frequency dynamics in the health insurance market. The empirical analysis narrows down to health insurance premiums outside of the mandatory framework (i.e. private health insurance). These will be referred to in the paper mostly as health insurance premiums Bosnia and Herzegovina presents a specific case of a largely decentralised health financing system, which is the result of the country's political and economic decentralised shape. The mandatory scheme is thus decentralised across the Republic of Srpska, the Federation of Bosnia and Herzegovina (further decentralised through cantons), and Brčko Distrikt. This being said, another reasoning is that the institutional framework of mandatory health insurance in Bosnia and Herzegovina does not operate on the basis of risk-rated insurance premiums but on payroll-based contributions (Law on Contributions, 1998, Law on Contributions, 2017). These contributions are set administratively, usually as a fixed percentage of gross

[25] In Bosnia and Herzegovina, the private insurance market is supervised by the Insurance Agency of the Republic of Srpska and the Insurance Supervision Agency of the Federation of BiH, with overall coordination by the Agency for Supervision of Insurance of Bosnia and Herzegovina (AZOBiH). See: https://www.azobih.gov.ba/.

wages, and do not vary with the market conditions. As such, they cannot be directly observed or modeled as "premiums" in the actuarial and econometric sense. WHO/IRIS reports confirm that contributions for mandatory health insurance are collected and pooled at the canton/Entity level, further reinforcing that "premiums" under this framework are not applicable in the public sector of BiH (World Health Organization, 2022).

By employing an Autoregressive Distributed Lag (ARDL) model with quarterly data from 2011 to 2023, the paper examines whether private health insurance premiums in BiH are cointegrated with macroeconomic, fiscal-demographic, and labor-financial factors, while also accounting for structural breaks such as the COVID-19 pandemic and the conflicts in Ukraine.

2. LITERATURE REVIEW

A substantive literature has explored so far plenty of key factors influencing health expenditures and has used a variety of econometric methodologies to analyze their relationships. However, much less attention has been devoted to health insurance premiums specifically, especially in the Western Balkans, where empirical evidence remains scarce. This has been used to provide a contextual background for studying health insurance premiums and claims, particularly through an Autoregressive Distributed Lag (ARDL) lens.

Global health expenditures have seen a rapid increase over the last two decades, driven by widespread environmental deterioration, extensive economic activities, and other social factors (Usman et al., 2019). Bosnia and Herzegovina is no exemption, given the structural shocks and breaks these factors have had worldwide. On the other had, not only rising economic activity has produced effects, but also shocks within, affecting stable and sustainable growth.

With that in mind, it is valid to observe the economic growth in that context. Since economic growth is vastly measured by gross domestic product (GDP), and its derived indicators, it is mandatory to highlight the fact that the GDP consistently shows a positive and statistically significant impact on both government and private health expenditures (Usman et al., 2019). This suggests that as economies grow, there is a corresponding increase in demand for medical services and a greater ability to finance healthcare (Ozyilmaz et al., 2022). Usman et al. (2019) found that a one percent increase in GDP can lead to a considerable rise in both government and private health expenditures. This relationship has been observed across various country groups, including emerging economies (Usman et al., 2019; Ozyilmaz et al., 2022) and OECD countries (Yang et al., 2022).

However, the nature of this causality varies in the literature. Some of these studies found a unidirectional causality from economic growth to health expenditures, like the study by Usman et al. (2019), whereas others found a bidirectional causality, indicating that health expenditures also stimulate economic growth through improved labor productivity and quality of life (Yang et al., 2022; Ozyilmaz et al., 2022).

Demographics and social factors have been proven to have a significant effect on health expenditures. Population aging is a strong determinant of increasing health expenditures for both government and private systems in emerging economies (Usman et al., 2019; Chen and Wang, 2023). Both authors found that the elderly often require more costly medical facilities and treatments for chronic illnesses and multiple morbidities. Usman et al. (2019) project the acceleration of the trend, posing a significant challenge for healthcare budgets. Studies in China, for instance, specifically highlight that aging population induces a relatively strong reaction from health expenditure per capita (Lopreite and Zhu, 2020). The aging index can significantly influence health spending (Chang et al., 2022; Yang et al., 2022).

What is also found is that private insurance and out-of-pocket payments can be alternative approaches for reducing the impact on public financing (Usman et al., 2019). This distinction is particularly relevant for BiH, where the data capture the private segment only. Because private premiums are priced and adjusted in competitive markets, they tend to be more income- and price-sensitive than contribution-financed public coverage, making them a suitable dependent variable for short- and long-run macro-linkages. Private insurance, though smaller in scope, often reacts more quickly to macroeconomic shocks and thus serves as a useful dependent variable for studying short- and long-run dynamics.

In China, aging induces more elderly people to be covered by health insurance, which consequently decreases private health expenditure (Chen and Wang, 2023). This suggests an important link between demographic trends, insurance coverage, and private spending. Moreover, expenditures related to health insurance in general are found to positively affect economic growth within Chinese provinces (Ozyilmaz et al., 2022).

When it comes to voluntary (private) health insurance, and health insurance outside of the mandatory systems, a significant impact of income on the demand was found, which consequently raises equity concerns regarding access to healthcare, especially in the context where voluntary private health insurance predominantly covers the wealthy (Ćurak et al., 2021). The present study builds on this literature by empirically testing the macroeconomic and demographic determinants of private health insurance in BiH, filling a gap where comparable analyses have been more or less absent.

In European countries, the demand for voluntary private health insurance (VPHI) has significantly increased, with density growing by 122.3% from 2004 to 2017 (Ćurak et al., 2021). Comparable statistics for Bosnia and Herzegovina are lacking, even that of health

insurance in general, which further underlines the contribution of this study in addressing this empirical gap. On the other hand, in Nigeria, for example, only 5% of the population has any health insurance, with 70% still relying on out-of-pocket payments (Alawode and Adewole, 2021), mostly due to its non-mandatory nature.
When it comes to what drives health insurance in general, within the observed literature, it has been shown that GDP, both in nominal, and real terms, significantly reflects overall economic capacity and is positively associated with health expenditure and insurance uptake (Atılgan et al., 2017; Qehaja et al., 2022; Stepović et al., 2020). Still, the literature largely focuses on health expenditures rather than health insurance premiums, and no study has explicitly modelled these relationships in the BiH context using an ARDL approach. Prices, proxied by consumer price index (CPI) and a narrower term - Health CPI, have been proven to capture inflationary pressures, which directly affect the cost structure of health services and insurance (Stepović et al., 2020). Additionally, Stepović et al. (2020) showed that money, defined by M2 and its ratio to GDP, serves as a proxy for financial sector development and liquidity, influencing both public and private health insurance markets.
When it comes to methodological approaches, the study of long-run relationships between economic variables often employs cointegration techniques to account for non-stationary time series data (Nkoro and Uko, 2016). The Autoregressive Distributed Lag (ARDL) cointegration technique, or bounds testing approach, is particularly favored due to its flexibility and robustness (Chang et al., 2022; Nkoro and Uko, 2016).
The ARDL model, additionally, has seen extensive and continuous usage in health economics for analyzing both short- and long-term relationships among variables with mixed integration orders, i.e. I(0) and I(1), making it ideal for time series data in transitional economies like Bosnia and Herzegovina (Atılgan et al., 2017). To the best of author's knowledge, no previous study has employed the ARDL bounds testing approach to investigate the determinants of private health insurance premiums in Bosnia and Herzegovina. This focus is partly a matter of data availability, but it also reflects the fact that private health insurance premiums are more sensitive to economic dynamics, unlike mandatory schemes that are politically and institutionally rigid. This study therefore contributes to filling this gap by combining PCA-based factor aggregation with ARDL modeling.
Some of the key advantages found in ARDL are that, first, it is insensitive to variable integration order. It can be applied irrespective of whether the underlying variables are integrated of order I(0) or I(1). However, it cannot be applied if variables are integrated of order I(2) (Chang et al., 2022; Nkoro and Uko, 2016; Chen and Wang, 2023). Secondly, the ARDL models allow for a simple linear transformation to an Error Correction Model (ECM), which reflects the speed of adjustment to equilibrium (Nkoro and Uko, 2016). Finally, the ARDL error correction representation is relatively more efficient for small or finite sample sizes ($n \leq 30$) (Chang et al., 2022; Nkoro and Uko, 2016).

3. DATA AND METHODOLOGY

3.1. Defining Hypotheses

Drawing on the theoretical arguments and the empirical evidence discussed in the previous section, this study formulates a set of testable hypotheses concerning the determinants of private health insurance premiums in Bosnia and Herzegovina. The purpose of explicitly defining the hypotheses is to provide a transparent empirical framework and to link the literature review with the econometric methodology applied in the subsequent sections. All hypotheses are stated in both null (H0) and alternative (H1) form to enable directional testing within the ARDL–ECM framework. For hypotheses that specify an expected sign (e.g., positive long-run effect), I evaluate one-sided p-values in line with the directional alternative.
Prior to elaborating on the five hypotheses, a precondition hypothesis is introduced. This tests whether a stable long-run equilibrium exists between private health insurance premiums and their key determinants. In econometric terms, this involves assessing cointegration among the macroeconomic, fiscal-demographic and labor-financial variables that in synergy, influence the insurance market. The null assumes that no such long-run relationship is present, whereas the alternative posits that at least one cointegrating vector exists, meaning that despite short-term fluctuations, premiums and their determinants share a common stochastic trend that binds them together in the long-run.
The first hypothesis (H1) concerns the macroeconomic environment. Previous studies consistently find that economic growth, liquidity, and price dynamics are among the strongest drivers of health-related expenditures and insurance demand. Consequently, it is hypothesised that macroeconomic performance (GDP, money supply, and consumer prices) exerts a positive long-run effect on private health insurance premiums. The null hypothesis assumes that such a relationship does not exist.
The second hypothesis (H2) addresses fiscal and demographic pressures. Population ageing and increased public health expenditure are expected to create structural demand for complementary private health insurance. While the null hypothesis assumes no influence of these factors, the alternative hypothesis states that fiscal-demographic pressures positively affect private health insurance premiums in the long run.
The third hypothesis (H3) focuses on labour-market and financial constraints. Since voluntary health insurance

is typically a discretionary financial product, it is sensitive to short-term shocks in income, employment, and access to credit. The null hypothesis postulates no effect, whereas the alternative hypothesis expects a negative short-run impact of labour-financial stress on insurance premiums.

The fourth hypothesis (H4) is methodological in nature and relates to the stability of the estimated system. Within the ARDL–ECM framework, stability is confirmed if deviations from the long-run equilibrium are corrected over time. The null hypothesis assumes the absence of such adjustment ($\pi \geq 0$), while the alternative hypothesis postulates a negative and statistically significant error-correction term ($\pi < 0$), indicating convergence to equilibrium.

Finally, the fifth hypothesis (H5) accounts for extraordinary structural shocks, such as the COVID-19 pandemic and the war in Ukraine. Both events are expected to reduce the uptake of voluntary insurance by increasing uncertainty and reducing household disposable income. The null hypothesis assumes no effect of these shocks, while the alternative hypothesis anticipates a significant negative effect. COVID-19 and Ukraine dummies are treated as exogenous structural breaks. Given the predicted negative direction, one-sided tests are used for H5.

3.2. The Autoregressive Distributed Lag

Autoregressive Distributed Lag model (ARDL) is a technique for detecting and estimating long-run relationships among time series variables. It is particularly useful, as stated before, for not requiring pretests for unit roots. Its key advantage is robustness when there is a single long-run relationship between the variables, especially with small sample sizes.

The ARDL cointegration technique (i.e. bound test of cointegration) was proposed by Pesaran and Shin (1995) following Pesaran's earlier work on ARDL models.

General ARDL model can be written as:

$$Y_t = \alpha + \sum_{i=1}^{p} \gamma_i Y_{t-i} + \sum_{j=1}^{k} \sum_{i=0}^{q_j} \beta_{j,i} X_{j,t-i} + \delta W_t + \epsilon_t$$

where Y_t is the dependent variable at time t; $X_{j,t-i}$ represents the j-th independent variable at lag i; α is a constant term; γ_i are coefficients for the lagged dependent variable; $\beta_{j,i}$ are coefficients for the lagged independent variables; p is the optimal lag order for the dependent variable; q_j is the optimal lag order for the j-th independent variable; k is the number of independent variables; W_t is a vector of deterministic variables (e.g. intercept, time trends, seasonal dummies, or exogenous variables with fixed lags); u_t (or ϵ_t) are error terms, assumed to be identically and independently distributed (iid) (0, δ^2; $\Phi(\mathrm{L},\mathrm{p}) = 1 - \Phi_1\mathrm{L} - \Phi_2 L^2 - \cdots - \Phi_p L^p$; $\beta(\mathrm{L},\mathrm{q}) = \beta_0 - \beta_1\mathrm{L} - \cdots - \beta_q L^q$. For the bounds test (Case 3), k denotes the number of level regressors included in the UECM, which determines the lower/upper critical bounds.

3.3. Data

In order to equate variable frequency, temporal disaggregation had to be conducted. This was done through CRAN's tempdisagg package, utilising Chow-Lin and Denton-Cholette methods where applicable. As the dataset is quarterly in frequency, variables such as public health expenditure (PUBHEXP) and Age 65+ to working population ratio (AGE65_PC), being published with yearly values, had to be disaggregated to quarterly frequency.

Let's assume that y_t is the goal high-frequency series, in this case quarterly series; \Delta^2y_t is the second difference of the series, and T is the total number of high-frequency periods.

The minimum problem function presents as:

$$\min_{y_t} \sum_{t=3}^{T} (\Delta^2 y_t)^2 = \sum_{t=3}^{T} (y_t - 2y_{t-1} + y_{t-2})^2.$$

Pluging in the variables, I create a solid same-frequency dataframe.

Seasonality-wise, the seasonal components were omitted from the affected variables using CRAN's package seasonal and its included X-13ARIMA-SEATS methods developed by U.S. Census Bureau. After the dataset was uniformed, variables were natural log-transformed. Conducting test-regressional models, a high multicollinearity was noticed among the variables. This is expected when it comes to such set of macroeconomic and demographic variables given their simultaneous movements and inter-dependence. Plugging these variables in the model as such, would produce a biased output with deviated and untrustworthy results, providing us with zero to little economic meaning and interpretation. This was tested through variance inflation factor (VIF) tests, where the results yielded whopping VIFs, which is an indicator of high multicollinearity.

To account for multicollinearity, the model needed contraction in terms of dimensionality. This is best achieved through principal components extraction via principal component analysis, which is common preprocessing technique prior to conducting ARDL models (Hooley et al., 2022).

What was conducted can be summarized as follows. First, I summarized and ordered the dataframe, along with all the operations done prior (disaggregation, log-transformation, seasonal adjustment, etc.). After that, potential factors were identified and grouped in the following order. Macro_vars dataframe was created consisting of nominal gross domestic product (GDP_NOM), industrial production index (IPI), consumer price index (CPI), and broad money monetary aggregate - M2 (M2). This group is to account for macroeconomic surroundings. Next, a fiscal_vars group was formed consisting of public health expenditure (PUBHEXP) and a demographic variable presenting a ratio of those aged 65+ in the BiH population (AGE65_PC). This is to account for fiscal environment along with the demographic control, given the interdependence of these two variables. Finally, thelabfin_vars frame was constructed consisting of net wage variable (NET_WAGE), a number of those actively seeking work (ACTIVE_UNEM), and finally the yearly average lending rate given at 3-month levels, to account for banking and the sector depicting liquidity in the economy (LEND_RATE).

All data was collected from official statistical and institutional sources. Specifically, health insurance premiums were obtained from the Agency for Supervision of Insurance of BiH (Agency for Supervision of Insurance of BiH, 2024). Macroeconomic indicators (GDP, CPI, M2, IPI) were drawn from the World Bank (2024), IMF Data Portal which aggregates IMF databases (International Monetary Fund, 2024), and the Central Bank of Bosnia and Herzegovina (Central Bank of Bosnia and Herzegovina, 2024). Public health expenditure and demographic indicators (share of population aged 65+) were taken from the World Bank Database (World Bank, 2024), and Agency for Statistics of BiH (2024) respectively. Labor market and financial sector variables (wages, unemployment, lending rates) were also provided by Agency for Statistics of BiH (2024) and Central Bank of Bosnia and Herzegovina (2024). A full overview of data sources is presented in Table 1.

Table 1: Data sources

Variable	Source
Health insurance premiums	AZO BiH (2024)
GDP, CPI, M2, IPI	IMF (2024); CBBiH (2024)
Public health expenditure	BHAS (2024)
Population aged 65+	World Bank (2024)
Net wages, unemployment	BHAS (2024)
Lending rates	IMF (2024)

Source: Author's aggregation

For each block of variables (macroeconomic, fiscal-demographic, and labour-financial), two principal components were initially extracted. To facilitate interpretability, the components were subjected to orthogonal varimax rotation. Following rotation, only one component per block was retained, selected on the basis of the largest explained variance and consistent factor loadings. These retained components were then labelled as MACRO, FISCAL, and LABFIN, and subsequently used as composite regressors in the ARDL framework. Prior to extraction, all variables were standardised using Z-score transformation. The dependent variable - health insurance premiums (HIPREM) was not subject to principal component analysis, but was rather only standardized.

Table 2: Varimax – rotated loadings for extracted components

Variable	MACRO (F1)	FISCAL (F2)	LABFIN (F3)
GDP_NOM	0.992		
IPI	-0.425		
CPI	0.833		
M2	0.941		
PUBHEXP		0.991	
AGE65_PC		0.991	
NET_WAGE			-0.947
ACTIVE_UNEM			0.786
LEND_RATE_PC			0.348

Source: Author's calculation

Factor 1 (MACRO) is defined by very high loadings on nominal GDP, consumer prices, and the monetary aggregate M2. This cluster reflects the broad stance of the economy, capturing both real activity and monetary conditions. GDP_NOM shows the scale and performance of output, CPI points to inflationary dynamics, while M2 embodies liquidity and money supply growth. Together, these variables indicate that the first component can be interpreted as a macroeconomic performance and stability factor, highlighting how output, prices, and money circulate jointly in the system. Factor 2 (FISCAL) is dominated by public health expenditure and the share of the population aged 65 and above, both with nearly perfect loadings. This dimension captures the dual challenge of aging societies: rising demand for healthcare and increasing fiscal burdens on government budgets. The tight correlation between demographic pressures and public spending justifies interpreting this factor as a fiscal-demographic burden, where structural shifts in population age drive long-term sustainability concerns for public finances. Factor 3 (LABFIN) is shaped by a strong negative loading on net wages, a strong positive loading on unemployment, and a weaker positive contribution from lending rates. This suggests a dynamic where labor market distress, low wages and high unemployment coincides with financial constraints. The factor therefore combines both social and financial vulnerabilities, making it best understood as a labor market and financial pressure component, pointing to the interdependence of employment conditions and access to credit in shaping economic stress.

These factors as such played a key role in creating a model. A model was created using these three factors as explanatory variables. To account for structural breaks and shocks, COVID and UKR step dummies were introduced, addressing the global health pandemic starting in Q2 2020, ending in late 2022, and the war in Ukraine starting 2022 Q1 and lasting ever since.

Unit-root diagnostics (ADF/PP) indicate a mix of I(0)/I(1) processes across series, with no I(2), which is compatible with the ARDL framework.

Note on Software Utilisation. All estimations and diagnostic checks were conducted in the R statistical environment (R Core Team, 2024). For stability analysis, I employed the strucchange package (Zeileis et al., 2002). Robust covariance matrices were obtained using sandwich (Zeileis, 2004), while inference relied on lmtest functions for coefficient and diagnostic testing (Zeileis & Hothorn, 2002). Dynamic regression specifications were facilitated by dynlm (Zeileis, 2019), factor extraction and psychometric routines by psych (Revelle, 2024), and autoregressive distributed lag estimation by the ARDL package (Pesaranghader & Wagner, 2021).

4. RESULTS

4.1. Base ARDL and Model Selection

Lag orders were selected by a bounded grid search with quarterly-appropriate caps (up to four lags for the dependent variable to allow one-year persistence, up to two lags for each regressor to capture short-run propagation), while COVID/UKR indicators entered in levels without lags to represent discrete shocks. Among candidates, I minimized AIC and retained the most parsimonious specification within Δ AIC ≤ 2 that was meant to satisfy multicollinearity and other diagnostics.

ARDL lag order specifications ranked by the Akaike Information Criterion. The ARDL(1,2,2,1) model achieved the lowest AIC value (−8.015), indicating that one lag of HIPREM, two lags of MACRO, two lags of FISCAL, and one lag of LABFIN provide the most adequate specification. Simpler models, such as ARDL(1,0,1,1), yield considerably higher AIC values, suggesting that some short-run dynamics in the macroeconomic and fiscal factors are required to adequately capture the data-generating process.

After settling most of the data-characteristic issues, finally I create the model equation:

$$\begin{aligned}\mathrm{HIPREM} = {} & \mathrm{L(HIPREM,1)} + \mathrm{MACRO} + \mathrm{L(MACRO,1)} + \mathrm{L(MACRO,2)} + \mathrm{FISCAL} + \mathrm{L(FISCAL,1)} \\ & + \mathrm{L(FISCAL,2)} + \mathrm{LABFIN} + \mathrm{L(LABFIN,1)} + \mid \mathrm{COVID} + \mathrm{UKR}\end{aligned}$$

One thing to note, before conducting the ARDL model, is the interpretation of this regression. Given that ARDL allows for I(0) and I(1) to be combined in the model without difference transformation, it is safe to assume that I might receive a spurious output as a result. Other authors that utilized this model received high, irrational $R2$ results along with issues like heteroskedasticity of residuals, multicollinearity issues, and even non-normal residuals.

This is, in a way, expected, with all duly addressing and precautionary steps to avoid such issues. As many state, it is completely normal to expect such output given the integration order of the variables. Given the mixed integration orders, I report levels-ARDL primarily as a staging model for the bounds test, with HAC (Newey–West) standard errors to address residual heteroskedasticity and autocorrelation. Inference on long-run coefficients is based on the UECM/RECM representation conditional on cointegration.

4.2. Diagnostic Tests

In the following section, diagnostics tests are conducted, prior to proceeding with the bounds test and other procedures, in order to ensure that the coefficients from the primary model can be safely interpreted as such. The diagnostic tests can be summarized below.

Table 3: Residual diagnostic tests for the ARDL model

Test	Statistic	df	p-value
Phillips–Perron (stationarity)	Z(α) = −71.094	—	< 0.01
Jarque–Bera (normality)	$\chi2$ = 1.569	2	0.456
Breusch–Pagan (heteroskedasticity)	BP = 26.923	11	0.0047
Breusch–Godfrey (serial corr., lag 4)	LM = 21.013	4	0.0003

Table notes: Residuals are stationary and normally distributed. However, heteroskedasticity and autocorrelation are present according to BP and BG tests. These issues are addressed by reporting HAC (Newey–West) robust standard errors.
Source: Author's calculation

The initial residual diagnostics (see Table 3) revealed evidence of heteroskedasticity (Breusch–Pagan test, p < 0.01) and autocorrelation (Breusch–Godfrey test, p < 0.001), despite residuals being both stationary and normally distributed. These violations imply that the classical OLS standard errors are unreliable, potentially biasing inference about coefficient significance. To address this issue, I re-estimated the model with heteroskedasticity- and autocorrelation consistent (HAC) standard errors, following the Newey and West (1987) procedure. I use HAC (Newey-West) with prewhitening and finite-sample adjustment. Importantly, the point estimates of the coefficients remain unchanged, but the standard errors were adjusted upward or downward to correct for serial correlation and heteroskedasticity in the residuals. This ensures that the reported t-statistics and p-values are valid even in the presence of such problems.

Table 4: Coefficient estimates with HAC (Newey–West) robust standard errors

Variable	Estimate	Std. Error (NW)	t	p
Intercept	0.268	0.131	2.053	0.047*
L(HIPREM,1)	0.270	0.079	3.402	0.0016**
MACRO	0.485	0.303	1.604	0.117
L(MACRO,1)	1.114	0.469	2.376	0.023*
L(MACRO,2)	-0.414	0.241	-1.717	0.094
FISCAL	1.019	0.696	1.464	0.151
L(FISCAL,1)	-3.052	1.263	-2.416	0.021*
L(FISCAL,2)	1.929	0.805	2.397	0.022*
LABFIN	-1.024	0.347	-2.955	0.005**
L(LABFIN,1)	1.121	0.399	2.812	0.008**
COVID	-0.585	0.242	-2.413	0.021*
UKR	-1.118	0.277	-4.034	0.0003***

Notes: Standard errors are heteroskedasticity- and autocorrelation-consistent (HAC, Newey–West). Significance levels: ***p<0.01, **p<0.05, *p<0.1.
Source: Author's calculation

After the Newey–West correction (Table 4), changes can be observed. Key findings remain robust, meaning that COVID and UKR dummies remain highly significant and negative; the lagged dependent variable indicates persistence in levels-ARDL (L(HIPREM,1)=0.270, p=0.0016). Some coefficients lose precision; for example, the direct effect of MACRO becomes insignificant at the 5% level, though its lag structure still shows significance. Fiscal variables show mixed evidence, with significance mainly in lagged terms, consistent with slower adjustment dynamics. Ultimately, LABFIN effects remain robust: the contemporaneous negative effect and lagged positive effect are still significant.

Overall, the substantive interpretation of the model is unchanged, but the inference is now statistically reliable. By applying Newey–West HAC standard errors, the issues of heteroskedasticity and autocorrelation have been explicitly addressed, ensuring that the reported significance levels are not artificially inflated.

To assess parameter stability, both recursive CUSUM and OLS-based CUSUM tests were applied (Brown et al., 1975) using the strucchange package in R. The recursive CUSUM statistic (S = 0.525, p = 0.571) and OLS-CUSUM statistic (S0 = 0.357, p = 0.999) fail to reject the null of stable coefficients. The fluctuation processes remain within the 95% confidence bounds, indicating no evidence of structural instability in the estimated ARDL/ECM model. The recursive MOSUM test indicated potential local instabilities around 2016Q2 and 2020Q1. However, Bai–Perron breakpoint selection based on BIC did not confirm these as statistically warranted structural breaks, as the inclusion of exogenous COVID-19 and Ukraine dummies already captures such shocks. This suggests that while short-lived shifts are visible in the fluctuation process, they do not undermine the overall stability of the ARDL/ECM specification.

4.3. Bounds Tests and ECM – Observing Long-Term Relationship

The relationship between health insurance premiums and the selected macroeconomic and health indicators was estimated using a clean ARDL approach. However, this approach can be reformulated to an Error Correction Model (ECM) form. This form allows for separation of shortand long-run dynamics, while explicitly including the error correction term (ectt−1) to capture deviations from long-run equilibrium.

The general restricted ECM can be written as:

$$\Delta y_t = c_0 + c_1 t + \sum_{i=1}^{p-1} \psi_{y,i} \Delta y_{t-i} + \sum_{j=1}^{k} \sum_{l=1}^{q_j-1} \psi_{j,l} \Delta x_{j,t-l}$$

$$+ \sum_{j=1}^{k} \omega_j \Delta x_{j,t} + \pi_y ECT_t + \epsilon_t$$

$$\psi_{j,l} = 0 \quad \forall q_j = 1, \qquad \psi_{j,l} = \omega_j = 0 \quad \forall q_j = 0$$

where ect_{t-1} denotes the error correction term obtained from the long-run cointegration equation, and under case 3 (no deterministic trend):

$$c_1 = 0$$

$$ECT = y_{t-1} - \left(\sum_{j=1}^{k} \theta_j x_{j,t-1} \right)$$

The general unrestricted ECM can be written through a formula of an unrestricted ECM conditional to an ARDL(p,q1,...,qk) is:

$$\Delta y_t = c_0 + c_1 t + \pi_y y_{t-1} + \sum_{j=1}^{k} \pi_j x_{j,t-1} + \sum_{i=1}^{p-1} \psi_{y,i} \Delta y_{t-i}$$

$$+ \sum_{j=1}^{k} \sum_{l=1}^{q_j-1} \psi_{j,l} \Delta x_{j,t-l} + \sum_{j=1}^{k} \omega_j \Delta x_{j,t} + \epsilon_t$$

$$\psi_{j,l} = 0 \quad \forall q_j \leq 1, \qquad \psi_{y,i} = 0 \quad \text{if } p = 1$$

In addition, if qj = 0 then $x\_\{j, t-1\}$ and \ Delta $x\_\{j,t\}$ cancel out, becoming simply $x\_\{j,t\}$.

First, however, the presence of a long-run relationship between the variables must be verified.

The Bounds F-test introduced by H. Pesaran and Shin (1995) was used and applied under Case 3 (unrestricted intercept, no trend). Long-run multipliers are obtained from the UECM as θˆ = −πˆx/πˆy, where πˆy denotes the coefficient on L(HIPREM,1) and πˆx the coefficient on L(x,1). Standard errors for θˆ are computed via the delta method using the HAC variance–covariance matrix. The results were as depicted in the Table 5.

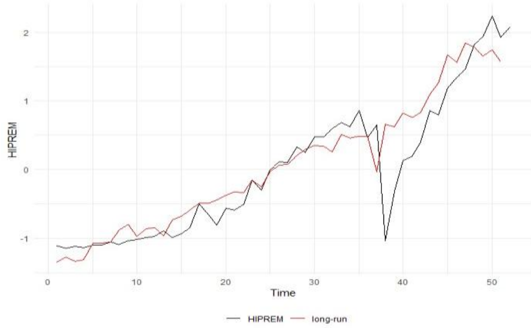
Table 5: Bounds F-Test for Cointegration

Test	Statistic	p-value
Bounds F-test (Wald)	4.774	0.015
Null hypothesis	No cointegration	
Alternative	Possible cointegration	

Source: Author’s calculation

The bounds F-test statistic is 4.774 (Case 3, k = 3). Since F exceeds the 5% upper critical value, I reject the null of no cointegration. This provides evidence of a long-run relationship between HIPREM and the levels regressors (MACRO, FISCAL, LABFIN). COVID and UKR are treated as exogenous shock dummies and do not form part of the long-run cointegrating vector. Stability of adjustment is confirmed separately by a negative and significant error-correction term in the error-correction model. Economically, this result implies that changes in these explanatory variables do not only affect health insurance premiums in the short run, but that they move together in the long run (see Figure 1). For instance, macroeconomic performance and demographic shifts eventually align with premium dynamics, suggesting structural co-movement.

Figure 1: Long-run Dynamics



Source: Authors' visualisation using CRAN's ggplot2 and ARDL packages

However, one potential limitation is that the test is sensitive to lag selection and small sample size. With relatively few observations (T close to 50), critical values may not be perfectly reliable, and the evidence of cointegration should be supported by additional diagnostics (e.g., stability tests, robustness checks).

Table 6: ECM Estimates with HAC (Newey–West) Standard Errors

Variable	Estimate	Std. Error	t value	Pr(>|t|)
	Restricted ECM			
(Intercept)	0.191	0.028	6.919	0.000***
ΔLABFIN	-0.856	0.457	-1.872	0.068*
COVID	-0.635	0.096	-6.607	0.000***
UKR	-0.912	0.140	-6.490	0.000***
ECT	-0.734	0.056.	13.106	0.000***
	Unrestricted ECM			
(Intercept)	0.191	0.054	3.542	0.001***
L(HIPREM, 1)	-0.734	0.060	-12.307	0.000***
MACRO	0.937	0.459	2.042	0.047*
FISCAL	0.063	0.389	0.161	0.873
L(LABFIN, 1)	0.063	0.094	0.673	0.504
ΔLABFIN	-0.856	0.343	-2.499	0.016*
COVID	-0.635	0.166	-3.820	0.000***
UKR	-0.912	0.305	-2.990	0.005**

Notes: HAC (Newey–West) robust errors. Significance codes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Author's calculation

4.3.1. Short-Run Interpretation

The restricted ECM estimates indicate that a 1% increase in the LABFIN factor immediately reduces private health insurance premiums by about 0.86%. This suggests that labor market and financial stress, characterized by lower wages, higher unemployment, and tighter lending conditions, directly suppress person's ability or willingness to pay for voluntary (i.e. private) health insurance. At the same time, the coefficients for COVID-19 (-0.63) and Ukraine (-0.91) dummies show that these structural shocks significantly reduced premiums, by around 0.6–0.9 units on average, highlighting the sensitivity of the private insurance market to extraordinary crises.

4.3.2. Long-Run Adjustment Dynamics

The bounds test confirms the existence of a cointegration relationship, allowing the estimation of both short-run and long-run coefficients. In the unrestricted ECM, the coefficient on the lagged dependent variable implies a long-run multiplier for the macroeconomic factor of θMACROˆ = 1.275 (SE 0.653; one-sided $p = 0.025$). This confirms that stronger macroeconomic performance significantly increases voluntary health insurance premiums in the long run, supporting H1. The contemporaneous short-run effect of MACRO is smaller in magnitude (≈ 0.9), indicating that the full adjustment materialises gradually through the error-correction mechanism. By contrast, the fiscal–demographic factor is economically small and statistically insignificant (θFISCALˆ = 0.085, $p = 0.436$), so H2 cannot be confirmed. The adjustment coefficient is negative and significant ($\pi = -0.734$, $p < 0.001$), implying rapid convergence towards the long-run equilibrium within two to three quarters, consistent with the typical renewal cycle of insurance contracts.

4.4. Hypotheses Testing and Interpretation

Precondition – Cointegration. The ARDL bounds test confirms the presence of at least one long-run cointegration relationship between private health insurance premiums and their macroeconomic, fiscal-demographic, and labour-financial determinants. Accordingly, the null hypothesis (H0A) is rejected, and the alternative (H1A) is accepted.

The long-run multiplier for the macroeconomic factor is positive and statistically significant (θˆMACRO = 1.275, one-sided $p = 0.025$), implying that stronger macroeconomic conditions are associated with higher private health insurance premiums. For these reasons, I reject the null hypothesis of no long-run cointegration, and accept the alternative hypothesis.

The long-run multiplier for the fiscal–demographic factor is small and not statistically significant (θˆFISCAL = 0.085, one-sided $p = 0.436$). For these reasons, H0 cannot be rejected, implying that the evidence for H2 is inconclusive at conventional levels. The contemporaneous short-run effect of LABFIN is negative and statistically significant (ΔLABFINd = − 0.856, one-sided $p = 0.003$), indicating an immediate contraction in voluntary insurance during periods of labour–financial stress. As stated, I can reject the null hypothesis, and accept the alternative.

The error-correction term is negative and highly significant ($\hat{\pi} = -0.734$, one-sided $p < 0.001$), which confirms stable adjustment toward the long-run equilibrium. Again, I reject the null.

The COVID-19 and Ukraine-war dummy coefficients are negative and statistically significant (both one-sided $p < 0.001$), consistent with a contraction in voluntary insurance during crises. Consequently, H0 is rejected and H1 accepted.

In summary, the results reject null-hypotheses of H1, H3, H4, and H5. Evidence for H2 is inconclusive at conventional levels (one-sided $p = 0.436$), so I fail to reject H0.

5. DISCUSSION

A stable long-run cointegrating relation is found between private health insurance premiums and macro-/labor-financial factors, with swift error-correction over two to three quarters. The long-run MACRO effect is positive, which is consistent with evidence that output, prices and liquidity scale health spending/insurance demand (Atılgan, 2017; Stepović, 2020; Chang, 2022). Short-run dynamics are dominated by LABFIN: adverse wage/unemployment/credit conditions immediately depress premiums, with a delayed rebound as conditions normalise, which can be noticed as a pattern aligning with demand elasticity and postponable personal expenditures (Casabianca, 2022). In contrast, the fiscal-demographic factor is not significant in the long run over the sample, which diverges from studies on public health expenditures (Usman 2019; Yang 2022) but is plausible for private premiums priced in competitive markets and over a shorter horizon. COVID-19 and the Ukraine war exert significant negative shocks, reinforcing international evidence that crises curtail voluntary insurance uptake. The significant intercept and fast ECT suggest a drifting equilibrium level and quick mean-reversion in premium flows. Together, results position private premiums as an income-sensitive barometer of household financial resilience rather than a proxy for structural ageing/fiscal pressures.

These results have important policy implications. They suggest that macroeconomic stability and labor market recovery play a decisive role in sustaining the voluntary insurance segment, while fiscal and demographic pressures act mainly through public channels. Strengthening household financial security and improving credit access may, henceforth, enhance insurance penetration more effectively than demographic subsidies or short-term fiscal measures.

5.1. Limitations

Although the ARDL-ECM approach offers robustness in modelling short- and long-term dynamics, several limitations should be considered. First, the sample covers a relatively short time span (2011–2023) and a small cross-sectional domain limited to Bosnia and Herzegovina, which constrains the generalisability of results. Quarterly frequency improves estimation efficiency but inevitably compresses structural shifts that may have occurred gradually, such as regulatory reforms or market liberalisation episodes.

Second, data availability restricts the inclusion of microeconomic and institutional variables that could better capture the behavioural dimension of health insurance demand. Aggregate macro-fiscal indicators proxy these mechanisms only imperfectly, leaving potential measurement bias in unobserved heterogeneity, especially across income and age groups. Moreover, the absence of continuous data on public health expenditure and demographic transitions narrows the ability to fully identify long-run fiscal–demographic pressures.

Third, the estimation strategy assumes linear and symmetric adjustment, whereas the insurance market may respond non-linearly to shocks and policy uncertainty. Future research could explore asymmetric ARDL specifications, regime-switching, or structural break modelling to capture such complexities. Additionally, the analysis excludes cross-country spillovers and regional integration effects, which could be relevant given the common exposure of Western Balkan economies to EU monetary and financial conditions.

5.1.1. Factor Limitations

A further limiting concern might be the construction of the FISCAL factor. The almost perfect correlation between public health expenditure and the share of population aged 65+ (reflected in their near identical factor loadings), implies that this composite variable might not be able to separate fiscal from demographic effects. As a consequence, the insignificance of FISCAL in the long-run equation might reflect two points: either a genuine absence of fiscal-demographics pressures on private health insurance premiums, or on the other hand, an identification problem which arises from the collinearity of the two variables. The aggregation into a factor was necessary however, due to the fact that, when these two regressors were included separately, the model exhibited clear signs of overfitting, including unstable coefficient estimates, inflated standard errors and weak out of sample performance, again due to high collinearity among the two, combined with limited sample size. Collapsing them into a single factor was therefore necessary in terms of contractionary mathematical rigor. Future research with longer time series, higher-frequency fiscal data, or cross-country panel structures could attempt to separate these channels more precisely.

Moreover, the negative IPI loading in the MACRO factor (IPI=-0.425) requires clarification. This pattern might indicate what structural composition of BiH economy looks like, where, as expected, the services sector accounts for the dominant and growing share of GDP, while industrial output has exhibited flat or countercyclical movements relative to aggregate economic expansions (predominantly in nominal terms), e.g. COVID and Ukraine-Russia conflict. During the observed period (2011–2023), nominal GDP growth was driven primarily by trade, financial services, and public administration, whereas industrial production remained subdued due to the legacy of deindustrialisation and limited manufacturing investment. Similar patterns have been documented in other post-transition economies of the Western Balkans, where service-led growth coexists with stagnant industrial bases (Bartlett, 2009; Uvalić, 2010). The negative loading thus captures the divergence between overall macroeconomic performance and industrial activity, rather than a statistical anomaly. As a robustness consideration, an alternative MACRO factor excluding IPI was tested, yielding qualitatively similar results.

Concludingly, the interpretation of PCA-derived factors requires caution, as these are latent constructs rather than directly observable economic variables. Each factor represents a weighted combination of its constituent indicators, and the factor scores capture common variation rather than the isolated effect of any single variable. The MACRO factor should therefore be understood as a composite index of macroeconomic conditions (output, prices, and liquidity) rather than as a proxy for GDP alone. Similarly, LABFIN reflects the joint state of labour market tightness and credit conditions. This aggregation approach, while necessary to address multicollinearity, implies that coefficient estimates represent the response of health insurance premiums to broad economic dimensions rather than to specific policy instruments. The trade-off between multicollinearity reduction and interpretability is a recognised limitation of factor-augmented regression approaches (Stock & Watson, 2002).

Despite these limitations, the present model offers a coherent empirical framework for understanding the macro-fiscal linkages of private health insurance premiums and establishes a foundation for subsequent regional and micro-level extensions.

6. CONCLUSION

In conclusion, this study demonstrates that private health insurance premiums in Bosnia and Herzegovina are not merely passive indicators but are statistically and economically linked to broader fundamentals. Macroeconomic performance (H1), labour-financial stress in the short run (H3), stable adjustment (H4), and extraordinary shocks (H5) emerge as robust determinants of private health insurance premiums. Fiscal-demographic pressures (H2) are not statistically significant in the long run over the sample period. The existence of a long-run cointegration relationship underscores that premiums adjust over time to changes in macroeconomic conditions (GDP and prices). Demographic–fiscal pressures do not exhibit a statistically significant long-run effect over the sample. This resonates with evidence from OECD countries, where insurance markets are found to co-evolve with growth. (Younsi et al., 2024). Additionally, the significant negative impacts of COVID-19 and the Ukraine conflict illustrate that external shocks materially affect the premium setting in the private insurance sector, reinforcing that this market is sensitive to both domestic and global disturbances. Similar results were observed in international contexts, where voluntary insurance uptake decreased under adverse economic shocks or crises (Casabianca et al., 2022). The robustness of these findings, confirmed via ARDL-ECM modeling, HAC standard errors, and diagnostic plus stability testing, suggests that private premiums could function as an early warning mechanism for deteriorations in economic or labor-financial health. This is consistent with recent meta-analytical evidence showing that health insurance and broader economic performance are tightly interlinked, with causality running in both directions (Fan et al., 2024).

For policy, this suggests that regulators, insurers, and government actors should not ignore the trajectories in private insurance premium data when crafting health financing strategies. Measures such as transparency in premium setting, support for financial resilience among households, and macroeconomic stability policies could help mitigate adverse impacts of shocks. It may also be useful to develop data collection mechanisms for the mandatory contribution-based side of health insurance so that future studies and policy decisions can consider the system as a whole.

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