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Godwin Olasehinde-Williams, Mehmet Balcilar

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The Long-run Effect of Geopolitical Risks on Insurance Premiums

Muhammad Shahbaz [a], **Godwin Olasehinde-Williams** [b], **Mehmet Balcilar**[a,b,c,*]

[a]Montpelier Business School, Montpelier, France
[b]Eastern Mediterranean University, Northern Cyprus, via Mersin 10, Turkey
[c]University of Pretoria, Pretoria, South Africa

Abstract

This paper deals with the estimation of the risk-insurance nexus. We specifically examined the effect of geopolitical risk on insurance premium in a panel of 18 countries while controlling for the effect of real income. We did this by employing second generation econometric methods. We found strong evidence of a positive impact of geopolitical risks on insurance premiums. We specifically found that the impact of geopolitical risks on non-life insurance premium is higher than the impact on life insurance premium. Real income was also found to have a significantly positive effect on insurance premiums, and the impact on non-life insurance premium similarly larger than the impact on life insurance premium. On the basis of the income elasticity, we found that insurance has the semblance of a luxury good, and on the basis of the panel causality tests, we confirmed the feedback hypothesis. We therefore conclude that exposure to geopolitical risks raises insurance premiums either as a result of increased insurance demand to hedge against geopolitical uncertainty or as a result of uncertainty tax factored into insurance premiums in unstable environments.

Keywords: Geopolitical risks, insurance market, panel data, panel bootstrap causality

JEL Codes: C33, G22, G32, O16

[*]Corresponding author. E-mail: mehmet@mbalcilar.net

1. Introduction

It is common knowledge that risks associated with wars, terrorism and tensions between states (geopolitical risks) exert powerful influences on macroeconomic and financial cycles. There is ample evidence that geopolitical disturbances create considerable amount of risks and uncertainties in the economy at large and in the financial sector in particular (Glick and Taylor 2010; Guidolin and La Ferrara 2010; Aghion et al. 2012; Christofis et al. 2013; Moretti et al. 2014; Apergis and Apergis 2016; Balcilar et al. 2016; Caldara and Iacoviello 2016; Balcilar et al. 2017; Drautzburg et al. 2017).

Moreover, because of recent geopolitical tensions (such as US-China and US-Russia tensions, escalations in the North-Korean and Syrian conflicts and fragmentation in Europe) and their potential adverse effects on the general economic outlook of the affected regions, concerns about geopolitical risks have heightened in recent times. According to the 'Global Risks 2015' report of the World Economic Forum, geopolitical risks currently rank very high on the list of issues that multinational businesses are concerned about, and this perception is not expected to change over the next decade.

A recent study by PwC (2018) shows that businesses are now more concerned about broader risks such as geopolitical risks than direct business risks. The study specifically shows that geopolitical risk is currently a top-five threat to businesses in every region of the world. Economic consequences of geopolitical risks are therefore an issue that is currently on the front burner of policy and academic discourse. Table 1 provides an insight into how high geopolitical risks rank on the list of threats faced by businesses around the world.

Table 1. Perception of top 5 business threats on regional basis

Africa	Asia-Pacific	Central & Eastern Europe	Latin-America
Social-instability	Availability of key skills	Availability of key skills	Populism
Increasing tax burden	Speed of technological change	Over-regulation	Inadequate basic infrastructure
Over-regulation	Geopolitical risks	Geopolitical risks	Increasing tax burden
Uncertain economic growth	Cyber-threats	Populism	Over-regulation
Geopolitical risks	Over-regulation	Changing workforce demographics	Geopolitical risks

Middle-East	North-America	Western Europe
Geopolitical risks	Cyber-threats	Populism
Cyber-threats	Over-regulation	Over-regulation
Over-regulation	Geopolitical risks	Geopolitical risks
Speed of technological change	Speed of technological change	Cyber-threats
Increasing tax burden	Increasing tax burden	Climate change

Source: PwC, 21st Annual Global CEO Survey.

The role of the insurance sector is broadly categorized into two; protection from adverse events (risk pooling and indemnification against losses) and financial intermediation (mobilization of savings and provision of investment opportunities). Since risk management is one of the primary functions of the insurance sector, it means that a strong correlation must exist between risks in general and the insurance sector. For example, Haiss and Sumegi (2006) posit that insurance lowers volatility and uncertainty in the economy and cushions the impact of crises at both micro and macro levels, thereby stabilizing the economic system.

Extant literature has also shown that major political events and their associated risks exert significant influence on the financial sector (Roe and Siegel 2011). For example, Ward and Zurbruegg (2002) show that increased political stability significantly raises life insurance consumption both in Asia and in OECD countries. According to Beck and Webb (2003), political instability can negatively affect the life insurance market due to its effect on the economic horizon of parties involved in life insurance transactions. The study by Lee et al. (2013) confirm that economic and political risks significantly affect the income elasticity of insurance demand. Furthermore, the insurance industry, in addition to being a key part of the financial sector, is also strongly intertwined with other sectors in the financial system (Billio et al. 2012). For these reasons, it is thus again logical to assume that a relationship exists between risks and the insurance sector because of its unique place in the financial sector.

The inter-linkage between risks and insurance has led to series of studies on the risk-insurance nexus. The larger parts of these studies are however micro-based analyses, focused on concepts such as information asymmetry and changes in consumer behavior (Lee et al. 2013). Macro-based studies on this relationship are still limited. There are two major reasons why this is so. The first is because over the years, the most significant risks that managers faced were direct business risks arising from issues such as moral hazard, adverse selection, new entrants into a particular sector, and changing consumer behavior. The second cause of this limitation is the lack of adequate macro-indicators for risks that are consistent over time.

The spread of globalization and liberalization has however changed the pattern of risks being experienced all around the world, with the nature of risks being experienced rapidly becoming macro-based as risks are becoming increasingly globalized and interconnected (Toma et al. 2012). A typical example of this new dimension of risks is geopolitical risks. Also, the advent of macro risk indicators such as the geopolitical risk index of Caldara and Iacoviello (2016) has made it possible to carry out macro-level studies on the risk-insurance nexus. This paper therefore examines the impact of geopolitical risks on insurance premiums (total, life and non-life).

In summary, this study for the benefit of the insurance sector and policy makers provides an evaluation of the effect of geopolitical risks on insurance premiums. In addition to the need for such a study, this paper also takes into account the role that income level plays in the insurance market. This is mainly because literature has confirmed real income as the most important stimulant of the insurance sector (Beck and Webb 2003; Li et al. 2007; Lee et al. 2010). Our contribution to literature includes; the use of the recently developed geopolitical risk index of Caldara and Iacoviello (2016), the application of second-generation panel tests, and the avoidance of pre-test bias that possibly exists in the existing risk-insurance literature through the use of the Emirmahmutoglu and Kose (2011) Panel Granger causality tests. The rest of this paper is organized as follows; section 2 presents the data and model used, section 3 describes the methodology adopted, section 4 presents the empirical results, and section 5 concludes the study.

2. Data and model specification

The study sample used includes the following 18 countries: Argentina, Brazil, China, Columbia, India, Indonesia, Israel, South Korea, Malaysia, Mexico, Russia, Saudi Arabia, South Africa, Philippines, Thailand, Turkey, Ukraine and Venezuela. The time period covered extends from 1985 to 2016. The choice of countries used and time frame selected is based on data availability.

To examine the effect of geopolitical risks on insurance premiums (total, life and non-life), we generate a panel data series for 5 variables—total insurance premium, life insurance premium, non-life insurance premium, geopolitical risk index and real gross domestic product (real income). Data on the insurance premiums was sourced from 'Swiss Re, Sigma database' and all three premiums are measured in millions of US dollars. Data for real GDP was retrieved from the World Development Indicators (http://data.worldbank.org) and is measured in 2000 US dollars. As for geopolitical risks, which is the variable of main interest, the monthly index for geopolitical risks constructed by Caldara and Iacoviello (2016) was downloaded from http://www.policyuncertainty.com/gpr.html and 12-month mean values were calculated.

The index is generated by searching of electronic archives of major newspapers for the following combination of words: geopolitical threats, nuclear threats, war threats, terrorist threats, war acts and terrorist acts. While the first 4 phrases focus on geopolitical threats and tensions, the last two are related to actual geopolitical occurrences. The number of articles containing these phrases is counted per month, a normalization process is then carried out to set the 2000-09 decade to an average value of 100. In this way, a value greater than 100 shows that newspaper mentions of risking political risks are higher than those recorded in the 2000-09 decade, while a value less than 100 means the opposite. The logarithmic forms of all the variables are used in the estimations.

Table 2 provides the summary statistics for total insurance premium, real GDP and geopolitical risk index. The reported values show that Philippines has the lowest mean total insurance premium (1504.83) while China has the highest mean total insurance premium (69926.21). Regarding the mean geopolitical risk index, South Africa is the riskiest (122.60) whereas Indonesia is the least risky (79.58). In terms of mean real GDP, Israel is the richest (26335.87) and India is the poorest (914.11). Philippines has the least variation in total insurance premium (1346.28). Mexico has the least variation in geopolitical risk (13.86) and real GDP (387.33). At the other extreme, China has the most variation in total insurance premium (95977.25), Ukraine has the most variation in geopolitical risk index (50.20) and South Korea has the most variation in real GDP (6223.59).

Table 2. Summary statistics.

Country	Mean	Std. Dev	Minimum	Maximum
Panel A: Total Insurance Premium (USDm)				
ARGENTINA	6120.20	4268.31	1394.59	16941.10
BRAZIL	26291.06	26873.69	2034.78	85444.46
CHINA	69926.21	95977.25	885.69	328439.10
COLUMBIA	3117.08	2841.91	488.59	10071.60
INDIA	24984.41	26469.05	2438.47	74570.51
INDONESIA	4663.26	5031.70	567.36	15560.18
ISRAEL	6360.14	3827.55	1474.01	14343.72
SOUTH KOREA	62648.23	44819.53	1152.47	159514.80
MALAYSIA	5979.91	4730.67	842.38	15863.51
MEXICO	10795.20	8115.08	1349.63	27358.29
RUSSIA	11888.77	8972.09	424.12	28421.31
SAUDI ARABIA	2428.88	2284.99	610.68	8128.27
SOUTH AFRICA	26433.66	15742.07	4183.00	54365.39
PHILLIPINES	1504.83	1346.28	315.77	5788.41
THAILAND	6212.82	6082.58	441.14	21696.12
TURKEY	4038.44	3928.88	248.69	12460.02
UKRAINE	1783.69	1445.48	165.88	4558.11
VENEZUELA	4677.88	5420.23	911.08	22664.57
Panel	16039.32	33802.54	165.88	328439.10
Panel B: Geopolitical risk index				
ARGENTINA	117.56	31.75	80.09	198.89
BRAZIL	103.94	17.92	70.17	140.86
CHINA	103.99	18.29	81.28	145.45
COLUMBIA	84.61	18.75	58.77	137.72
INDIA	97.16	20.49	70.17	148.38
INDONESIA	79.58	24.78	50.79	155.24
ISRAEL	86.66	15.20	66.81	127.77
SOUTH KOREA	110.78	24.12	80.13	174.77
MALAYSIA	95.48	17.76	70.52	139.83

MEXICO	97.01	13.86	78.55	127.84
RUSSIA	109.38	22.41	78.48	171.43
SAUDI ARABIA	94.90	23.44	62.77	132.16
SOUTH AFRICA	122.60	46.59	69.47	221.98
PHILLIPINES	103.99	26.79	49.99	154.97
THAILAND	100.37	22.15	66.90	154.27
TURKEY	113.13	30.62	73.54	179.06
UKRAINE	112.97	50.20	52.73	260.77
VENEZUELA	81.78	26.94	35.92	141.80
Panel	101.69	29.82	35.91	260.77
Panel C: Real GDP (constant 2000 dollars)				
ARGENTINA	8385.59	1485.59	5939.76	10780.02
BRAZIL	9383.14	1340.94	7796.84	11912.15
CHINA	2584.02	1965.44	538.69	6893.78
COLUMBIA	5273.05	1099.55	3751.15	7525.86
INDIA	914.11	420.52	446.99	1861.49
INDONESIA	2437.12	728.37	1386.48	3974.06
ISRAEL	26335.87	4378.16	18770.87	33673.13
SOUTH KOREA	15494.21	6223.59	5405.35	25458.89
MALAYSIA	7114.61	2185.00	3701.25	11031.82
MEXICO	8311.60	890.92	6893.92	9707.91
RUSSIA	8713.78	2151.01	5505.63	11615.70
SAUDI ARABIA	18836.34	1600.25	15608.75	21507.96
SOUTH AFRICA	6491.62	752.20	5516.76	7626.81
PHILLIPINES	1788.25	387.33	1380.68	2753.35
THAILAND	3806.99	1238.69	1666.28	5901.88
TURKEY	8925.56	2443.53	5659.22	14116.98
UKRAINE	2526.99	583.01	1686.94	3311.96
VENEZUELA	12634.48	1158.19	9710.27	14652.18
Panel	8348.41	6805.78	446.99	33673.13

3. Methodology

3.1 Preliminary tests

As a first step in our analysis, we test for the presence of cross-sectional dependence and slope heterogeneity in our data series so as to decide on the most appropriate unit root and cointegration tests to employ. Not controlling for the effect of these two issues when in fact they exist in the data series will cause us to obtain unreliable parameter estimates. The cross-sectional dependence tests and slope homogeneity tests conducted and the results are as follows;

Breusch-Pagan (1980) LM test: under a null hypothesis of no cross-sectional dependence, this LM statistic for dependence is given as:

$$LM = \sum_{i=1}^{N-1}\sum_{j=i+1}^{N} T_{ij}\hat{p}_{ij}^2 \rightarrow \chi^2 \frac{N(N-1)}{2} \qquad (1)$$

Where $\hat{p}_{ij}^2$ = correlation coefficients from residuals

Pesaran (2004) Scaled LM test: this test is more efficient than the Breusch-Pagan LM test for data series with large N. Also under the null of no cross-sectional dependence the LM statistic is:

$$LM_s = \sqrt{\frac{1}{N(N-1)}}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}(T_{ij}\hat{p}_{ij}^2 - 1) \rightarrow N(0,1) \qquad (2)$$

Pesaran (2004) CD test: a major flaw of the previous tests is that they suffer from size distortion. To address this shortcoming, Pesaran (2004) developed an alternative test statistic derived from averaging pairwise correlation coefficients $\hat{p}_{ij}$. Again for the null of no cross-sectional dependence the LM test statistic for dependence is specified thus:

$$CD_p = \sqrt{\frac{2}{N(N-1)}}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N} T_{ij}\,\hat{p}_{ij} \rightarrow N(0,1) \qquad (3)$$

Baltagi et al. (2012) Bias-corrected Scaled LM test: this test makes asymptotic bias corrections to the scaled LM test statistic, and under the same null of no cross-sectional dependence, the LM test statistic is specified as:

$$LM_{BC} = \sqrt{\frac{1}{N(N-1)}}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}(T_{ij}\hat{p}_{ij}^2 - 1) - \frac{N}{2(T-1)} \rightarrow N(0,1) \qquad (4)$$

For all four variations of the cross-sectional dependence test, the null of no cross-sectional dependence is represented as:

$$H_0: p_{ij} = corr(\mu_{it}, \mu_{jt}) = 0 \; for \; i \neq j \qquad (5)$$

As regards the issue of slope heterogeneity, we employ the Pesaran and Yamagata (2008) standardized version of the Swamy (1970) homogeneity test called the delta tests. A modified version of the Swamy (1970) test is first computed as:

$$\tilde{S} = \sum_{i=1}^{N}(\hat{\beta}_i - \hat{\beta}_{WFE})' X_I' \frac{M_\tau X_i}{\tilde{\sigma}_i^2}(\hat{\beta}_i - \hat{\beta}_{WFE}) \qquad (6)$$

Where $\hat{\beta}_i$=Pooled OLS estimator, $\hat{\beta}_{WFE}$ = weighted fixed effect pooled estimator and $\tilde{\sigma}_i^2$ is the estimator.

Next, the standard dispersion statistics is computed as:

$$\tilde{\Delta} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - k}{2k}\right) \qquad (7)$$

Alternatively, the bias adjusted version of the standard dispersion statistics may be computed:

$$\tilde{\Delta}_{Adj} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - E(\tilde{z}_{it})}{\sqrt{var\,(\tilde{z}_{it})}}\right) \tag{8}$$

Both statistics are tested under the null of slope homogeneity.

3.2 Panel unit root tests

Non-stationarity in data series has serious consequences on the behavior and properties of such series. It generally renders the standard assumptions of the asymptotic analysis invalid. Not properly handling the presence of unit roots in the data series will lead to the production of unreliable parameter estimates and result in spurious regression. We therefore carry out panel unit root tests to determine the order of integration of the variables used in our analysis.

Although panel unit root tests are regarded as statistically superior to time series unit root tests, the possible presence of cross-sectional dependence in panel data series is a challenge that is unique to panel data series. To deal with this issue, we employ the cross-sectional augmented Dickey Fuller (CADF) and cross-sectionally augmented IPS (Im et al. 2003) panel unit root tests.

The CADF unit root test statistic as specified by Pesaran (2007) is given as:

$$CADF_i = t_i(N,T) = \frac{\left(y_{i,-1}^T \bar{M} y_{i,-1}\right)^{-1}\left(y_{i,-1}^T \bar{M} \Delta y_i\right)}{\sqrt{\sigma_i^2 \left(y_{i,-1}^T \bar{M} y_{i,-1}\right)^{-1}}} \tag{9}$$

And the CIPS test statistic as specified by Pesaran (2007) for a null of unit root against a heterogeneous alternative is:

$$CIPS(N,T) = N^{-1}\sum_{i=1}^{N} t_i\,(N,T) = \frac{\sum_{i=1}^{N} CADF_i}{N} \tag{10}$$

Where $t_i(N,T)$ = ith cross-section CADF statistic.

The CIPS provides averages of the CADF test statistics for the entire panel and $t_i(N,T)$ is the cross-sectionally augmented Dickey-Fuller test statistic for the ith cross section unit in the CADF regression.

3.3 Panel cointegration test

In a case where the data series is plagued with cross-sectional dependence and slope heterogeneity, and the variables are integrated of a mixed order, it is possible to test for long-run relationship among the variables using the Durbin-Hausman cointegration tests of Westerlund (2008). The Durbin-Hausman tests are not only valid in the presence of cross-sectional dependence and slope heterogeneity, they are also applicable when the variables are integrated of a mixed order (I(0) and I(1)). The only condition required being that the dependent variable is non-stationary.

There are two types of the Durbin-Hausman tests. The first is the panel test (DH_p) in which the null hypothesis of no cointegration [$H_0: \varphi_i = 1$, for all I =1] is tested against an alternative of cointegration for all n units [$H_i^p: \varphi_i = \varphi, and\ \varphi < 1$]. The test statistic is specified as:

$$DH_p = \hat{S}_n(\tilde{\varphi} - \hat{\varphi})^2 \sum_{i=1}^{n} \sum_{t=2}^{T} \hat{e}_{it-1}^2 \tag{11}$$

The second is the group mean test (DH_g) in which the null hypothesis of no cointegration [$H_0: \varphi_i = 1$, for all I =1] is instead tested against the alternative hypothesis of cointegration in some of the cross sectional units. In this case the test statistic is:

$$DH_g = \hat{S}_i(\tilde{\varphi}_i - \hat{\varphi}_i)^2 \sum_{i=1}^{n} \sum_{t=2}^{T} \hat{e}_{it-1}^2 \tag{12}$$

3.4 Error-correction based panel estimations

The long-run impact of the regressors on insurance premium (total, life and non-life) is estimated via an error-correction form of the ARDL model. We begin by specifying ARDL models of the form:

$$TIP_{it} = \gamma_i + \sum_{j=i}^{p} \lambda_{ij}\, TIP_{it-j} + \sum_{j=0}^{q} \delta_{ij} GDP_{it-j} + \varepsilon_{it} \tag{13}$$

$$LIP_{it} = \gamma_i + \sum_{j=i}^{p} \lambda_{ij}\, LIP_{it-j} + \sum_{j=0}^{q} \delta_{ij} GDP_{it-j} + \varepsilon_{it} \tag{14}$$

$$NLIP_{it} = \gamma_i + \sum_{j=i}^{p} \lambda_{ij}\, NLIP_{it-j} + \sum_{j=0}^{q} \delta_{ij} GDP_{it-j} + \varepsilon_{it} \tag{15}$$

Here, i represents the number of groups (1,2,…,N) , t represents the number of periods(1,2,…,T), δ and λ represent coefficients and γ_i refers to the group specific effects.

Equations (13-15) are re-specified as error correction equations:

$$\Delta TIP_{it} = \emptyset_i(TIP_{it-1} - \theta_i GDP_{it}) + \sum_{j=i}^{p-1} \lambda_{ij}^{*}\, \Delta TIP_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta GDP_{it-j} + \varepsilon_{it} \tag{16}$$

$$\Delta LIP_{it} = \emptyset_i(LIP_{it-1} - \theta_i GDP_{it}) + \sum_{j=i}^{p-1} \lambda_{ij}^{*}\, \Delta LIP_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta GDP_{it-j} + \varepsilon_{it} \tag{17}$$

$$\Delta NLIP_{it} = \emptyset_i(NLIP_{it-1} - \theta_i GDP_{it}) + \sum_{j=i}^{p-1} \lambda_{ij}^{*}\, \Delta NLIP_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta GDP_{it-j} + \varepsilon_{it} \tag{18}$$

Where: $\emptyset_i = -\left(1 - \sum_{j=i}^{p} \lambda_{ij}\right)$ = speed of adjustment, if $\emptyset_i$ = 0, there is no proof of long-run relationship.

$\Theta_i = \frac{\sum_{j=0}^{q} \delta_{ij}}{1 - \sum_k \lambda_{ik}}$, $\lambda_{ij}^{*} = -\sum_{m=j+1}^{p} \lambda_{im}$ and $\delta_{ij}'^{*} = -\sum_{m=j+1}^{q} \delta_{im}$

To accommodate heterogeneous slopes, the Mean Group (MG) estimator of Pesaran and Smith (1995) which allows intercepts, slope coefficients and error variances to differ across cross-sections is used to estimate equations (16-18).

3.5 Panel causality tests

We further examine the causal links between the variables in order to determine their ability in forecasting one another with the aid of panel causality testing. We apply the Emirmahmutoglu and Kose (2011) panel causality test with bootstrapping. Emirmahmutoglu and Kose (2011) show that the Fisher (1932) test statistic may be used to test for panel Granger non-causality and specified thus:

$$\lambda = -2\sum_{i=1}^{N} ln_{(pi)} \ \ i = 1,2, \dots, N. \tag{19}$$

pi represents the p value for the ith cross section and the test statistic has a chi-square distribution with 2N degrees of freedom.

In the application of this test, it is not compulsory for all the series in the underlying VAR system to be stationary, it may thus be applied to panels composed of stationary, non-stationary, cointegrated and non-cointegrated series (Seyoum et al. 2014). In addition, the test is also suitable for panel data series affected by cross-sectional dependence and slope heterogeneity.

Following Emirmahmutoglu and Kose (2011), we adopt the lag augmented VAR (LA-VAR) model with L_y+ $dmax_i$ lags in heterogeneous mixed panels. It is specified as follows:

$$IP_{it} = a_{1i}^{IP} + \sum_{j=1}^{L_{IP}+dmax_i} B_{1ij} IP_{it-j} + \sum_{j=1}^{L_{GPR}+dmax_i} \gamma_{1ij} GPR_{it-j} + \varepsilon_{1it} \tag{20}$$

$$GPR_{it} = a_{2i}^{GPR} + \sum_{j=1}^{L_{IP}+dmax_i} B_{2ij} IP_{it-j} + \sum_{j=1}^{L_{GPR}+dmax_i} \gamma_{2ij} GPR_{it-j} + \varepsilon_{2it} \tag{21}$$

$$IP_{it} = a_{1i}^{IP} + \sum_{j=1}^{L_{IP}+dmax_i} B_{1ij} IP_{it-j} + \sum_{j=1}^{L_{GDP}+dmax_i} \gamma_{1ij} GDP_{it-j} + \varepsilon_{1it} \tag{22}$$

$$GDP_{it} = a_{2i}^{GDP} + \sum_{j=1}^{L_{IP}+dmax_i} B_{2ij} IP_{it-j} + \sum_{j=1}^{L_{GDP}+dmax_i} \gamma_{2ij} GDP_{it-j} + \varepsilon_{2it} \tag{23}$$

$$GPR_{it} = a_{2i}^{GPR} + \sum_{j=1}^{L_{GDP}+dmax_i} B_{2ij} GDP_{it-j} + \sum_{j=1}^{L_{GPR}+dmax_i} \gamma_{2ij} GPR_{it-j} + \varepsilon_{2it} \tag{24}$$

$$GDP_{it} = a_{2i}^{GDP} + \sum_{j=1}^{L_{GPR}+dmax_i} B_{2ij} GPR_{it-j} + \sum_{j=1}^{L_{GDP}+dmax_i} \gamma_{2ij} GDP_{it-j} + \varepsilon_{2it} \tag{25}$$

The null hypotheses for each pair of bivariate Granger causality tests are:

$$H_0: \gamma_{1i1} = \gamma_{1i2} = \dots = \gamma_{1ik_i} = 0 \ for \ i = 1,2, \dots, N \tag{26}$$

$$H_0: \beta_{2i1} = \beta_{2i2} = \dots = \beta_{2ik_i} = 0 \ for \ i = 1,2, \dots, N \tag{27}$$

4. Empirical results

4.1 Preliminary test results

The results from the 4 cross-sectional dependence tests are presented in table 3. We find evidence in favour of the rejection of the null hypothesis of no cross-sectional dependence at 1% significance level in all the cases. The presence of cross-sectional dependence is thus confirmed in our data.

Table 3. Cross-sectional dependence test results

CD Tests.	Test Stat. and Prob.				
	TIP	LIP	NLIP	GPR	GDP
LM (Breusch,Pagan 1980)	523.21***	556.89***	604.75***	201.27***	484.17***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CDlm (Pesaran 2004)	21.16***	23.08***	25.83***	2.76 ***	18.93***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CD (Pesaran 2004)	-2.99***	-2.58***	-2.31***	-2.89 ***	-2.51***
	(0.00)	(0.00)	(0.011)	(0.00)	(0.06)
LMadj (PUY, 2008)	2.01***	-0.03***	3.48***	2.59 ***	0.89***
	(0.02)	(0.51)	(0.00)	(0.00)	(0.17)

The significant test statistics for all the delta and adjusted delta tests presented in table 4 lead to the rejection of the null of slope homogeneity at 1%. We thus confirm the presence of slope heterogeneity.

Table 4. Slope homogeneity test results

Delta Tests.	Test Stat. and Prob.		
	TIP	LIP	NLIP
$\hat{\Delta}$	4.54***	4.48***	5.51***
	(0.00)	(0.00)	(0.00)
$\hat{\Delta}_{adj}$	4.77***	4.69***	5.78***
	(0.00)	(0.00)	(0.00)

4.2 Unit root test results

Table 5 presents the CADF and CIPS panel unit root test results with intercept and trend at levels. The CIPS results show that on the average, all the variables are stationary at level except for the geopolitical risk index. Table 6 reports the CADF and CIPS results with intercept and trend after first difference. The CIPS estimates show that on average, all the variables are stationary after first difference. We therefore conclude that while geopolitical risk index is integrated of order zero, all the other variables are integrated of order one.

Table 5. Panel unit root test results with intercept and trend at levels.

Countries	Test Statistics					Critical Values		
	TIP	LIP	NLIP	GPR	GDP	1%	5%	10%
ARGENTINA	-1.68	-3.38	-2.67	-2.92	-3.51	-4.97	-4.01	-3.65
BRAZIL	-1.50	-1.14	-3.33	-4.16^{**}	-0.47	-4.97	-4.01	-3.65
CHINA	-3.18	-4.14^{**}	-3.67^{*}	-3.13	-0.77	-4.97	-4.01	-3.65
COLUMBIA	-1.78	-2.48	-1.54	-1.81	-3.59	-4.97	-4.01	-3.65
INDIA	-2.72	-2.33	-1.69	-2.29	-1.79	-4.97	-4.01	-3.65
INDONESIA	-1.60	-1.51	-3.21	-1.57	-1.48	-4.97	-4.01	-3.65
ISRAEL	-1.15	-2.03	-4.06^{**}	-2.15	-2.53	-4.97	-4.01	-3.65
SOUTH KOREA	-2.44	-2.75	0.70	-2.86	-3.16	-4.97	-4.01	-3.65
MALAYSIA	-4.63^{**}	-2.43	-1.81	-1.97	-1.26	-4.97	-4.01	-3.65
MEXICO	-1.97	-1.80	-1.29	-2.97	-6.71^{***}	-4.97	-4.01	-3.65
RUSSIA	-3.36	-0.24	-4.80^{**}	-2.28	-2.14	-4.97	-4.01	-3.65
SAUDI ARABIA	-2.55	-2.18	-1.54	-4.08^{**}	-0.12	-4.97	-4.01	-3.65
SOUTH AFRICA	-2.50	-3.33	-4.46^{**}	-3.08	-6.07^{***}	-4.97	-4.01	-3.65
PHILLIPINES	-2.86	-1.58	-3.60	-3.39	1.24	-4.97	-4.01	-3.65
THAILAND	-1.77	-1.96	-2.20	-2.24	-6.48^{***}	-4.97	-4.01	-3.65
TURKEY	-3.61	-4.10^{**}	-3.00	-3.38	-6.42^{***}	-4.97	-4.01	-3.65
UKRAINE	-0.84	-2.01	-3.65^{*}	-2.50	-0.32	-4.97	-4.01	-3.65
VENEZUELA	-3.06	-4.33^{**}	-1.01	-1.71	1.49	-4.97	-4.01	-3.65
CIPS stat. for all countries (Panel)	-2.40	-2.43	2.47	-2.69^{**}	-2.54	-2.83	-2.67	-2.58

Table 6. Panel unit root test results with intercept and trend after first difference.

Countries	Test Statistics					Critical Values		
	TIP	LIP	NLIP	GPR	GDP	1%	5%	10%
ARGENTINA	-4.77^{**}	-6.13^{***}	-5.41^{***}	-6.60^{***}	-5.00^{***}	-4.97	-4.01	-3.65
BRAZIL	-5.40^{***}	-5.19^{***}	-5.27^{***}	-3.41	-9.71^{***}	-4.97	-4.01	-3.65
CHINA	-6.21^{***}	-4.29^{**}	-5.06^{***}	-2.67	-2.54	-4.97	-4.01	-3.65
COLUMBIA	-3.79^{*}	-6.07^{***}	-3.61	-2.23	-3.53	-4.97	-4.01	-3.65
INDIA	-5.35^{***}	-3.58	-6.21^{***}	-5.25^{***}	-3.17	-4.97	-4.01	-3.65
INDONESIA	-2.59	-6.89^{***}	-5.02^{***}	-2.76	-2.85	-4.97	-4.01	-3.65
ISRAEL	-4.71^{***}	-4.54^{**}	-6.28^{***}	-4.67^{**}	-5.31^{***}	-4.97	-4.01	-3.65
SOUTH KOREA	-6.08^{***}	-3.53	-4.96^{**}	-6.27^{***}	-6.21^{***}	-4.97	-4.01	-3.65
MALAYSIA	-5.67^{***}	-7.13^{***}	-6.75^{***}	-4.19^{**}	-11.88^{***}	-4.97	-4.01	-3.65
MEXICO	-6.23^{***}	-3.59	-3.26	-3.70^{*}	-7.88^{***}	-4.97	-4.01	-3.65
RUSSIA	-5.78^{***}	-5.93^{***}	-6.54^{***}	-3.16	-4.24^{**}	-4.97	-4.01	-3.65
SAUDI ARABIA	-6.26^{***}	-5.23^{***}	-6.92^{***}	-4.13^{**}	-0.10	-4.97	-4.01	-3.65
SOUTH AFRICA	-5.60^{***}	-9.32^{***}	-5.53^{***}	-5.21^{***}	-11.05^{***}	-4.97	-4.01	-3.65
PHILLIPINES	-8.77^{***}	-6.41^{***}	-5.39^{***}	-4.70^{**}	2.94	-4.97	-4.01	-3.65
THAILAND	-8.77^{***}	-5.58^{***}	-2.68	-4.63^{**}	-6.51^{***}	-4.97	-4.01	-3.65

TURKEY	-4.57**	-5.98***	-3.76*	-5.22***	0.92	-4.97	-4.01	-3.65
UKRAINE	-8.15***	-5.45***	-3.00	-4.48**	-2.29	-4.97	-4.01	-3.65
VENEZUELA	-3.76*	-4.18**	-4.64**	-4.10**	1.24	-4.97	-4.01	-3.65
CIPS stat. for all countries (Panel)	-5.69***	-5.50***	-5.02***	-4.29***	-4.29***	-2.83	-2.67	-2.58

4.3 Cointegration test results

The significant DH_g estimates reported in table 7 lead to our rejection of the null of no cointegration at 1% in all cases tested. We thus reach the conclusion that cointegration exists in at least some of the panel cross-sections.

Table 7. Westerlund (2008) panel cointegration test results.

TIP	
DH_g	-3.98***
	(0.00)
DH_p	-1.39
	(0.32)
LIP	
DH_g	-5.21***
	(0.00)
DH_p	-1.02
	(0.20)
NLIP	
DH_g	-4.89***
	(0.00)
DH_p	-1.36
	(0.33)

Notes: Probabilities are reported in parentheses. The criterion used in this paper is the BIC with maximum number of factors (K) set equal to two. For the bandwidth selection, M was chosen to the largest integer less than $4 * (t/100)^{2/9}$ ****, ** and * indicate the rejection of the null of no cointegration at 1%, 5% and 10%level of significance.*

4.4 Error-correction based panel estimation test results

Table 8 presents the MG long-run estimation results. In all 3 instances, the estimated adjustment coefficients are negative and highly significant, ranging between -0.259 and -0.232. This confirms the existence of a long-run relationship among geopolitical risk, GDP and insurance premiums. It also confirms the achievement of convergence in all the cases. The results show that the long-run impact of geopolitical risk on total insurance premium is positive and significant. One percent rise in geopolitical risk index causes total insurance premium to rise by 1.232 percent. This is significant at 5 percent significance level. Geopolitical risk also positively impacts both life and non-life insurance premiums. A percentage rise in geopolitical risk index results in 0.099 percent rise in life insurance premium and 3.795 percent rise in non-life insurance premium. The results are significant at 10 and 5 percent respectively.

The reported coefficients also confirm that GDP has a positive effect on insurance premiums (total, life and non-life). When GDP increases by 1 percent, total insurance premium increases by 2.418 percent, life insurance premium increases by 6.277% and non-life insurance premium increases 6.326%. The results are significant at 10, 1 and 5 percent significance levels respectively. This confirms the conclusion reached by Beenstock et al. (1988), Outreville (1990), Browne and Kim (1993), Beck and Webb (2003), Ching et al. (2010), and Pradhan et al. (2014).

Table 8. Mean Group long-run Estimation results

TIP	
Adjustment coefficient	-0.26^{***}
	(0.00)
Long-run coefficients	
GPR	1.23^{**}
	(0.03)
GDP	2.42^{*}
	(0.09)
LIP	
Adjustment coefficient	-0.23^{***}
	(0.00)
Long-run coefficients	
GPR	0.09^{*}
	(0.09)
GDP	6.28^{***}
	(0.00)
NLIP	
Adjustment coefficient	-0.24^{***}
	(0.00)
Long-run coefficients	
GPR	3.79^{**}
	(0.03)
GDP	6.33^{**}
	(0.02)

4.5 Panel causality test results

The panel causality test results among the insurance premiums (total, life and non-life), geopolitical risk index and GDP in bivariate models for 18 countries are reported in table 9. Results indicate the following; (i) the null that LIP does not Granger cause GPR and the null that GPR does not Granger cause LIP are both rejected at 1% significance level. We conclude that the relationship between both variables is bi-directional, (ii) the null that NLIP does not Granger cause GPR and the null that GPR does not Granger cause NLIP are both rejected at 1% and 5% significance level respectively. We similarly conclude that the relationship between both variables is bi-directional, (iii) the null that TIP does not Granger cause GPR and the null that GPR does not Granger cause TIP are rejected at 1% and 5% significance level respectively. We again conclude that the relationship between both variables is bi-directional, (iv) the null that LIP does not Granger cause GDP and the null that GDP does not Granger cause LIP are both rejected at 1% significance level. Thus, the verdict of bi-directional causality is reached, (v) the null that NLIP does not Granger cause GDP and the null that GDP does not Granger cause NLIP are both rejected at 1% significance level. The conclusion reached is that the relationship is bi-directional, (vi) the null that TIP does not Granger cause GDP and the null that GDP does not Granger cause TIP are both rejected at 1% significance level. Bi-directional causality is likewise inferred between the 2 variables, (vii) the null that GPR does not Granger cause GDP and the null that GDP does not Granger cause GPR are both rejected at 1% significance level. In summary, bi-directional causality is detected between all the variables.

Table 9. Results from Emirmahmutoglu-Kose Granger causality tests

Hypothesis	Statistic	P-Value	Conclusion
LIP→GPR	59.55 ***	0.008	Two-way causality between LIP and GPR
GPR→LIP	64.889 ***	0.002	
NLIP→GPR	292.923***	0.000	Two-way causality between NLIP and GPR
GPR→NLIP	56.205**	0.017	
TIP→GPR	63.430***	0.003	Two-way causality between TIP and GPR
GPR→TIP	56.673**	0.015	
LIP→GDP	508.463***	0.000	Two-way causality between LIP and GDP
GDP→LIP	134.203***	0.000	
NLIP→GDP	700.176***	0.000	Two-way causality between NLIP and GDP
GDP→NLIP	155.792***	0.000	
TIP→GDP	865.650***	0.000	Two-way causality between TIP and GDP
GDP→TIP	115.464***	0.000	
GPR→GDP	59.326***	0.009	Two-way causality between GPR and GDP
GDP→GPR	75.426***	0.000	

5. Conclusion

In this study, we empirically examined the impact of geopolitical risk on insurance premiums while controlling for the effect of income level in 18 countries for the period 1985-2016. The challenges posed by cross-sectional dependence and slope heterogeneity were factored into our estimations. For this reason, we first determined the order of integration of the variables using CADF and CIPS unit root tests of Pesaran (2007). Having established that the variables of interest are integrated of a mixed order, we tested whether cointegration exists between the variables using the Durbin-Hausman cointegration tests of Westerlund (2008). Upon confirmation of the existence of a long-run relationship, we estimated the long-run impact of geopolitical risk on insurance premiums (total, life and non-life) through the Mean Group estimator of Pesaran and Smith

(1995). The Mean Group estimation results showed that the impact of geopolitical risk and income level on the volume of insurance premiums is positive and significant.

Furthermore, to determine the direction of causalities in the presence of cross-sectional dependence and non-stationary variables, we performed the Emirmahmutoglu panel causality tests. The results confirmed bi-directional causality between geopolitical risk and the volume of the insurance premiums (total, life and non-life) and between income level and the volume of the insurance premiums (total, life and non-life). Based on our findings we come to the following conclusions.

The first is that the positive impact of geopolitical risk on the volume of insurance premiums recorded shows that rising geopolitical risks cause the volume of insurance premiums paid to rise. This may occur for two reasons. The number one reason is because insurance demand has risen in order to hedge against geopolitical uncertainty. This lends credence to the submission of Ward and Zurbruegg (2002) that political instability stimulates insurance consumption. The second reason may be because insurance plan premiums have risen in response to rising geopolitical risks. Insurers generally include some amount of uncertainty tax into insurance premiums charged in unstable environments. It is also possible that the increase in the volume of insurance premiums is also as a result of the combination of these two reasons. The second conclusion reached is that geopolitical risk has a bigger effect on non-life insurance premium than on life insurance premium.

We also infer from the positive impact of GDP on insurance premiums recorded that when income level is high, the volume of premiums recorded increases. Furthermore, from the size of the coefficients on GDP, we come to the conclusion that insurance has the semblance of a luxury good. Also, we find that the impact of income level is bigger on non-life insurance premium than on life insurance premium. Moreover, the bi-directional causality found between GDP and the insurance premiums (total, life and non-life) provides evidence in support of the feedback hypothesis on the relationship between the financial sector and economic growth.

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