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The link between life insurance activities and economic growth: Some new evidence

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This paper applies the panel seemingly unrelated regressions augmented Dickey-Fuller (SURADF) test to re-investigate the stationarity properties of real life insurance premiums per capita and real gross domestic product (GDP) per capita for 41 countries within three levels of income covering 1979–2007. Our empirical results first reveal that the variables in these countries are a mixture of I(0) and I(1) processes, and that the traditional panel unit-root tests could lead to misleading inferences. Second, for the estimated half-lives, the degrees of mean reversion are greater in high-income countries. Third, there is concrete evidence favoring the hypothesis of a long-run equilibrium relationship between real GDP and real life insurance premiums after allowing for the heterogeneous country effect. The long-run estimated panel parameter results indicate that a 1% increase in the real life premium raises real GDP by 0.06%. Finally, we determine that the development of life insurance markets and economic growth exhibit long-run and short-run bidirectional causalities. These findings offer several useful insights for policy-makers and researchers.

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1. Introduction

The importance of the insurance-growth relationship has risen over the past few decades due to the bigger makeup of insurance within the financial sector. The global insurance industry has seen an annual growth rate of over 10% since 1950, far exceeding that of global economic development (Dowling, 1982; Swiss Reinsurance Company, 1990; UNCTAD, 1972, 1991). This greater significance is also reflected in the business volume of life insurers.[1] The rapid growth of life insurance premiums not only increases insurers' role as providers of risk transfer, but also raises their importance as institutional investors. In addition, a number of international life insurance firms, such as American International Group (AIG) and International Netherlands Group (ING), have experimented with various industrial and banking linkages (Wilkins, 2009). Such developments have a profound influence in that, while they may promote economic activities, they give rise to risk in financial markets at the same time. These ideas prompted the initial motivation for this study, where we investigate the link between life insurance activities and economic growth.

It is essential to consider the relationship between life insurance and economic growth from both theoretical and empirical aspects. From a theoretical point of view, the relationship between life insurance and economic growth may run in either or both directions. The '*supply-leading*' and '*demand-following*' views as presented by Patrick (1966) postulate that economic growth (real income) can be enhanced either through growth in financial systems, or alternatively through growth in the economy, which brings about the development of financial activities. Based on the 'supply-leading' view, financial development enhances economic growth by transferring resources from traditional sectors to modern sectors and by promoting an entrepreneurial response in these modern sectors. In contrast, the 'demand-following' view indicates that a lack of financial development or institutions is due to a lack of demand for financial services. Thus, as the growth rate of real income rises, investors' and savers' demands for various new financial services materialize, hence leading to the creation of modern financial institutions, the supply of their financial assets and liabilities, and related financial services.

For the empirical aspect, previous studies mainly utilize time-series or cross-sectional datasets to investigate the relevant issues of life insurance premiums and macroeconomics, e.g., Ward and Zurbruegg (2000) and Kugler and Ofoghi (2005), to mention a few. These empirical works concentrate on a small group of countries over fairly short or distant time spans and conceivably suffer from the "small sample" problem.[2] However, researchers have recently been implementing panel data to analyze related issues (Beck and Webb, 2003; Arena, 2008; Haiss and Sümegi, 2008; Han et al., 2010; Lee, 2011, forthcoming; Chen et al., forthcoming).[3] Therefore, this study employs panel unit root, panel cointegration, and panel causality tests to explore the relationship between per capita real gross domestic product (hereafter RPGDP) and per capita real life insurance premiums (hereafter RPLIP; insurance density). Previous studies lack a diagnostic analysis of the order of integration for variables entering a long-run relationship between one another, which could lead to spurious regression bias. The presence of a unit root in real income and life insurance premiums has crucial implications for modeling the insurance-growth nexus.

Existing panel studies on insurance premiums do warn about the adverse effects of imposing homogeneity across countries and have employed a panel unit-root test combined with a panel

[1] For the period 1997–2007, the world's total written real insurance premiums increased approximately 5.5 times from US$0.63 trillion to US$4.13 trillion, while life insurance premiums rose approximately 7.5 times. In 2009, insurance companies worldwide wrote US$4.07 trillion in direct premiums, indicating the equivalent of about 7.0% of global GDP was used to purchase insurance products. In the same year, insurance companies in developing countries generated premiums worth US$0.53 trillion (13.1% of global insurance premiums), with the share of emerging and developing economies continuing to increase (9.3% of total business in 2000 and 12.0% in 2008). Such data from *Simga* (Swiss Reinsurance Company, 1980–2010) suggest that the development of the insurance market plays an increasingly substantial role within the financial sector.

[2] Odedokun (1996) found that panel data estimation yields more robust effects of financial development on economic growth than time-series estimation by individual countries.

[3] They used the GMM dynamic panel estimator, but did not consider the integration and cointegration properties of the data. Thus, it is not clear whether they represent a structural long-run equilibrium relationship or a spurious one (Christopoulos and Tsionas, 2004; Lee, 2011, forthcoming).

cointegration test to exploit the extra power derived from combining cross-sectional and time-series data. Through understanding the order of integration of variables and controlling for country heterogeneity using the panel data approach, we believe this study contributes to providing not only a clear picture of the interrelationships between the life insurance market's development and economic growth, but also presents a more accurate inference than would be shown by time-series or cross-country data alone.

The main purpose of this paper is to re-investigate the stationarity properties of RPLIP and RPGDP for 41 countries within three levels (high-, middle-, and low-) of income covering 1979–2007.[4] We apply the panel seemingly unrelated regressions augmented Dickey-Fuller (Panel SURADF) test developed by Breuer et al. (2001, 2002), which allows us to account for possible cross-sectional effects and to identify how many and which countries within the panel contain a unit root.[5] We proceed by measuring the half-lives and the corresponding confidence intervals when a variable's stationarity (integrated of degree zero; I(0)) is confirmed. Finally, after confirming the stationary process for the two series, we then investigate the relationships between the series based on the different orders of integration. For countries in which the two series are I(1), we employ the panel cointegration method to evaluate whether real GDP and real insurance premiums are cointegrated. In addition, we identify those countries for which the two series have different properties, i.e., where one is I(1) and the other is I(0). Under these conditions, real GDP and insurance premiums cannot be related to each other in the long-run and the panel cointegration analysis is inappropriate. To this end, we distinguish between long-run and short-run causalities.

Why should we take different income levels into account? First, while insurance companies play an increasingly key role in the financial sector and their increased importance as providers of financial services and investment funds in capital markets is quite pronounced in developed economies,[6] there are striking differences in many developing economies where insurance premiums remain low. Second, previous studies have argued that an increase in income results in greater demand for insurance, which then leads to higher premiums. Thus, an increase in income tends to raise people's purchasing power and living standards, thus resulting in growing demand for economic security, i.e. saving insurance and annuity. To partially resolve the problem of homogeneity in the panel data, we classify the panel data into three different sub-panels – namely, high-, middle-, and low-income groups – made up of 26, 8, and 7 countries, respectively (see Table A1 in Appendix). Another reason for this classification is that the relationships between the investigated variables under consideration could be sensitive to rich and poor countries.

The remainder of this paper is organized as follows. Section 2 provides a literature review. Section 3 discusses the econometric methods. Section 4 presents the empirical results. Finally, Section 5 reviews our conclusions, while also outlining some of the implications.

2. Literature review

There is a growing strand of theoretical and empirical literature assessing the linkage between the activities of the insurance market and real income as a proxy for economic growth, but along with many disparate channels there is a variety of conflicting results. Many earlier studies investigated the insurance-growth nexus from one of two perspectives: the demand side (or the insurance demand function) and the supply side (or the aggregate production function). As to the effect of economic

[4] In practice, we do not present that much of a time dimension. A common problem encountered in empirical research on insurance premiums is finding data covering a short time dimension.

[5] Such a procedure has already extensively been used in applied works, e.g., Rapach and Wohar (2004) and Koedijk et al. (2004, 2011), to mention a few. Our research is novel in that it is the first study that models RPLIP and RPGDP within a panel SURADF framework.

[6] The increased importance of the insurance sector is recognized by both insurers and the insured, because of the rise in risks and uncertainties in most societies. Because the life insurance sector has been growing rapidly, as mentioned earlier, its role is not just limited to providing protection to the insured, as it has a more prominent role in financial intermediation (Ching et al., 2010). In 2005, for example, the total assets of insurance companies were about US$17 trillion and institutional investors, such as insurance companies and mutual and pension funds, managed around 44% of an average household's holdings (Haiss and Sümegi, 2008).

growth (or real income) on the insurance market, several theoretical models of life insurance demand have been developed by Yarri (1965), Hakansson (1969), Fischer (1973), Karni and Zilcha (1985), and Lewis (1989). The pioneering study of Yarri (1965) pointed out that the demand for life insurance is attributed to an individual's desire to provide income for retirement and to bequeath funds to dependents, and it is derived from maximizing the individual utility function. In this framework, life insurance demand is a function of wealth, income stream, interest rates, prices including insurance premium rates, the consumer's subjective discount function for current consumption, future consumption, and wealth. In short, these theoretical models demonstrate that an individual's wealth and income affect life insurance demand, suggesting that a causal relationship may run from economic growth to life insurance activities.

For an empirical analysis, Fortune (1973), Lewis (1989), Beck and Webb (2003), Li et al. (2007), and Lee et al. (2010) all ascertained that life insurance demand is positively related to income. In assessing whether the life insurance demand behavior is sensitive to income growth, Hammond et al. (1967) used cross-sectional data classified into three income groups – i.e., low-income (US$3500 or less), middle-income (US$3501–US$6750), and high-income (US$6751 or more) – to examine the relationships between life insurance premium expenditures and various economic and demographic characteristics of households. They found life insurance premium expenditures to be inelastic with respect to income for the low- and high-income groups, while the middle-income group's estimated elasticity indicates that life insurance premium expenditures are highly responsive to changes in income. For the case of young newly-married couples' life insurance purchasing behavior, Anderson and Nevin (1975) noted that middle-income couples purchase less life insurance than lower- and upper-income couples, suggesting that young married couples do not uniformly purchase more life insurance as income increases.

Enz (2000) argued that supply and demand factors may limit insurance penetration, such that the income elasticities of insurance penetration are not constant, indicating an S-curve relationship between insurance and real income. Indeed, the income elasticity of insurance demand is one at both low- and high-income levels, but two or more at intermediate-income levels. Feyen et al. (2011) noted that insurance demand generally varies with income, showing that the very rich groups may not need life insurance, because they have surplus/excess assets, whereas the very poor groups do not have the means to buy life insurance products. They further found that the coefficient of the log proportion of total income by the top quintile is significantly positive, suggesting that life insurance is a relative luxury good. Lee and Chiu (2012) employed a panel smooth transition regression (PSTR) model to examine the non-linear insurance-income nexus. They presented that there are different income elasticities of life and non-life insurance premiums for countries with high- and low-income levels and concluded that the relationships between real GDP on life and non-life insurance premiums follow J-shaped and U-shaped patterns, respectively.

On the supply side of the insurance-growth nexus, attention has been paid to insurance's effect on economic growth. A large body of the theoretical literature (for example, Skipper, 1997; Rejda, 2005; Skipper and Kwon, 2007; Dorfman, 2008; Njegomir and Stojić, 2010) has discussed the insurance industry's influence on the economy and society, showing that it: (1) enhances financial stability; (2) facilitates the development of trade and commerce by increasing creditworthiness, lowering the total necessary amount and cost of capital, and decreasing total risk; (3) mobilizes domestic savings; (4) substitutes for government security programs; (5) fosters the efficiency of capital allocation; (6) facilitates loss mitigation; and (7) manages risk more efficiently. In the macroeconomic aspect, the insurance industry can also contribute to the formation of national income by creating added value. The service offered by the insurer is that of an intermediary, while knowledge of the insurance cost helps measure the effort made by the community to provide itself with an insurance system. On the basis of premiums collected and less liabilities incurred, this value added is apportioned for the payments of salaries and commissions, dividends, and indirect taxes.

The United Nations Conference on Trade and Development (UNCTAD) at its first session in 1964 formally acknowledged that "a sound national insurance and reinsurance market is an essential characteristic of economic growth". In general, insurance firms can stimulate economic growth by improving the functions of the financial system as providers of risk transfer and indemnification and as institutional investors (Ward and Zurbruegg, 2000; Haiss and Sümegi, 2008). More specifically, as discussed in Brainard (2008), insurance plays an important role in risk management, and insurers contribute their specialized expertise in the identification and measurement of risk. The indemnification and risk pooling characteristics of insurance facilitate both transactions and the provision of credit by alleviating losses as well as by assessing and managing non-diversifiable risk more often.

MacMinn (1987) showed that insurance helps mitigate agency problems, i.e. the underinvestment incentive due to a large asset-loss, thus leading to increased economic activity. Ward and Zurbruegg (2000) and Kugler and Ofoghi (2005) argued that in offering risk transfer, indemnification for unexpectedly large losses, real services, and financial intermediary services, insurance firms have had a crucial productive impact within economies. In this view, insurance likely helps banks mitigate their credit risk, invest in potentially high yielding projects, and increase corporate and private lending. Thus, the development of insurance contributes to economic growth.

From the empirical aspect of the aggregate production function, Soo (1996) found that growth in the life insurance industry contributes to both productivity growth and economic growth, and that the variance of life insurance-growth explains approximately 14% of the variance in economic growth. He showed that in the short-run and long-run, a unit shock in life insurance-growth has a positive impact on economic growth. Haiss and Sümegi (2008) further applied panel data to examine the relationship between insurance and economic growth based on an endogenous growth model for the whole EU, the EU-15+ (including Switzerland, Norway, and Iceland) and the emerging CEE countries (new EU Member States from Central and Eastern Europe). They found an insignificant effect from life and non-life insurance on economic growth for the whole EU and the EU-15+ countries, while non-life insurance seems to have a positive effect on economic growth in the CEE+ countries.

Most existing studies, however, only investigate either the impact of economic growth on insurance or conversely the impact of insurance on economic growth, while only a few studies examine the direction of causality between insurance market activities and economic growth. Ward and Zurbruegg (2000) investigated the causal relationship between the insurance industry growth and economic growth. They observed that in the long-run there is a bidirectional causal relationship between real insurance premiums and real GDP for Australia, Canada, Italy, and Japan, but a unidirectional causality from real GDP to real insurance premiums for France. Kugler and Ofoghi (2005) used nine different types of life insurance premiums to examine the causal relationship between life insurance and GDP for the U.K. and found evidence of a long-run causality from insurance premiums to GDP for six cases, a long-run causality from GDP to insurance premiums for one case, and a long-run bidirectional causality between them for the remaining two cases. Vadlamannati (2008) applied a vector error correction model (VECM) to investigate the short-run causality between India's life and non-life insurance sectors and its economic growth, indicating there is a bidirectional causality between life insurance sector growth and economic growth. Adams et al. (2009) provided evidence of unidirectional causality running from insurance to economic growth, but with no reverse effect, in the case of Sweden.

Concerning econometric methodology, the existing literature has investigated the insurance-growth nexus using either small, short-term cross-sectional samples, or relied upon panel techniques that do not account for the presence of potential endogeneity among variables. To address the endogeneity problem, Ćurak et al. (2009) applied a two-stage least squares technique (2SLS) and found that insurance sector development has a positive impact on economic growth. Arena (2008) used a generalized method of moments (GMM) approach to investigate whether insurance market activity promotes economic growth. The results indicated that life and non-life insurance both have different impacts on economic growth for different levels of economic development. Life insurance has a bigger impact on economic growth at low levels of economic development, whereas non-life insurance has a larger effect at middle levels of economic development.

Han et al. (2010) also employed the GMM approach to examine the relationship between insurance development and economic growth for a panel dataset of 77 countries. Their empirical results showed evidence of the positive impact of insurance development on economic growth, and that insurance development for developing countries plays a more important role than it does for developed countries. In discussing the direction of causality on the same topic, Webb et al. (2005) used the three-stage least squares instrumental variable approach (3SLS-IV) to reduce the endogeneity problem and examined the causal relationships between banks, life insurance, and non-life insurance on economic growth based on a revised Solow-Swan neoclassical economic growth model. The evidence exhibited bidirectional causality between life insurance and economic growth, yet no significant relationship between non-life insurance and economic growth in any direction.

Following this vein, we expect that there is a potential endogeneity between insurance market activities and real income. Thus, to understand the direction of causality between life insurance and economic growth, we examine the long-run and short-run causal relationships between the two. Real income should be treated as an endogenous variable when investigating the interrelationship between it and the insurance sector in order to reduce the bias of the estimated results. To eliminate endogeneity bias, this paper utilizes the dynamic ordinary least squares (DOLS) approach to estimate the relationship between them.

3. Methodology

Our econometric methodology proceeds in five steps. First, we implement the SURADF panel unit-root test to ascertain the order of integration for both variables. The sample countries are classified into two groups – I(1) and mixed – according to the results of the SURADF test. Second, to provide a complete analysis of the short-run adjustments and the mean reversion process of the insurance premium, we measure the half-lives and the corresponding confidence intervals when stationarity is confirmed. Third, conditional upon finding that both variables are I(1), we test for panel cointegration using the approach suggested by Pedroni (1999, 2004). Fourth, conditional upon finding a cointegration relationship, we calculate panel DOLS estimates of the coefficients of real GDP per capita. Fifth, we specify and estimate the most appropriate dynamic error correction model (ECM) for heterogeneous panels. Sixth and finally, from the perspective of insurance market activities and economic development policies, it is important to confirm any causality running from the insurance market to economic growth or from economic growth to the insurance market. Thus, we examine the long-run and short-run causal relationships between life insurance and economic growth.

3.1. The panel SURADF unit-root tests

It has been suggested that one feasible way to increase power when testing for a unit root is to use panel data. One of the most notable works is that of Breuer et al. (2001, 2002) who showed that the recent methodological refinements to the Levin et al. (2002, LLC) test fail to fully address the “all-or-nothing” nature of the test. Therefore, LLC test’s rejection signifies that at least one panel member is stationary, with no information about how many series are stationary.

Expanding upon this issue, Breuer et al. (2001, 2002) developed a panel unit-root test based on the augmented Dickey-Fuller (ADF; Dickey and Fuller, 1979) regression estimation in a seemingly unrelated regressions (SUR) framework and then tested for an individual unit root within the panel members. This procedure has several advantages. First, these multivariate tests use the information content in the variance–covariance matrix, thereby avoiding the unrealistic assumption of cross-sectional independence made in the panel tests. Second, conventional univariate unit-root tests not only fail to consider information across regions, but are also restricted in regard to the problem of a small sample, leading to less efficient estimations. In this regard, the multivariate ADF-type unit-root tests, by exploiting the information from the error covariance and allowing for an autoregressive process, produce more efficient estimators than the single-equation methods. Third, the estimation tests also allow for an important degree of heterogeneity in the lag structure across the panel members, in that the augmented test’s lag order varies among the individuals and the autoregressive parameter also differs for every cross section. Fourth, the panel SURADF unit-root test allows us to identify how many and which members of the panel contain a unit root.

The unit-root test of the panel SURADF for N countries and T time periods is based on the system of ADF equations, which can be represented as:

$$
\begin{aligned}
\Delta Z_{1,t} &= \alpha_1 + \beta_1 Z_{1,t-1} + \eta t + \sum_{j=1}^{k1} \varphi_{1,j} \Delta Z_{1,t-j} + u_{1,t} \qquad t = 1,2,\ldots,T \\
\Delta Z_{2,t} &= \alpha_2 + \beta_2 Z_{2,t-1} + \eta t + \sum_{j=1}^{k2} \varphi_{2,j} \Delta Z_{2,t-j} + u_{2,t} \qquad t = 1,2,\ldots,T \\
&\vdots \qquad\qquad\qquad\qquad\qquad\qquad\qquad\qquad \vdots \\
\Delta Z_{N,t} &= \alpha_N + \beta_N Z_{N,t-1} + \eta t + \sum_{j=1}^{kN} \varphi_{N,j} \Delta Z_{N,t-j} + u_{N,t} \qquad t = 1,2,\ldots,T,
\end{aligned} \tag{1}
$$

where Z denotes RPLIP (or RPGDP), and $u_{i,t}$ ($i = 1, 2, \ldots, N$) is an error term. Coefficient α_i is the heterogeneous constant term, $\beta_i = \rho_i - 1$, and ρ_i is the autoregressive coefficient for the ith cross-sectional member of the series, while t denotes the deterministic time trend.

Eq. (1) tests the null hypothesis of a unit root against trend stationarity. The model allows for heterogeneous fixed effects, heterogeneous trend effects, and heterogeneous lags for each cross-sectional unit in the panel. The flexibility to test for a unit root within each cross-sectional unit is especially beneficial for applied work where mixed stationary and non-stationary series are likely. This system is estimated by the SUR procedure, and we test the N null (H_o^i) and alternative hypotheses (H_A^i) individually as:

$$
\begin{aligned}
H_0^1 &: \beta_1 = 0; H_A^1 : \beta_1 < 0 \\
H_0^2 &: \beta_2 = 0; H_A^2 : \beta_2 < 0 \\
&\vdots \qquad\qquad \vdots \\
H_0^N &: \beta_N = 0; H_A^N : \beta_N < 0,
\end{aligned} \tag{2}
$$

with the test statistics being computed from SUR estimates of system (1), while the critical values are generated by Monte Carlo simulations.

3.2. The analysis of half-lives

The unit-root tests are even more importantly uninformative as to the speed of mean reversion. Alternatively, the "half-life" of deviation – defined as the number of periods required for a unit shock to dissipate by one half – measures the degree of mean reversion and the speed of adjustment back toward the long-run equilibrium. To examine this measure, suppose that the deviations of the series $Z_{i,t}$ from its long-run value $Z_{i,0}$ follow an AR(1) process:

$$Z_{i,t} - Z_{i,0} = \alpha(Z_{i,t-1} - Z_{i,0}) + \varepsilon_{i,t}, \tag{3}$$

where ε is white noise. The half-life deviation h is defined as the horizon at which the percentage deviation from the long-run equilibrium is one half – that is:

$$\alpha^h = \frac{1}{2} \Rightarrow h = \frac{\ln\left(1/2\right)}{\ln(\alpha)}. \tag{4}$$

A conventional 95% confidence interval associated with the above half-life statistic based on normal distributions can thus be defined as:

$$\widehat{h} \pm 1.96 \widehat{\sigma}_{\widehat{\alpha}} \left(\frac{\ln(0.5)}{\widehat{\alpha}} \left[\ln\left(\widehat{\alpha}\right) \right]^{-2} \right) \tag{5}$$

Here, $\hat{\sigma}_{\hat{\alpha}}$ is an estimate of the standard deviation of α. Since h cannot be negative, we impose a lower bound of zero (see Rossi (2005) for more details).

3.3. The panel cointegration tests

Pedroni (2004) considered the following panel data regression:

$$RPGDP_{it} = \alpha_i + \delta_i t + \beta_i RPLIP_{it} + e_{it}, \tag{6}$$

where subscripts i ($i = 1, 2, \ldots, N$) and t ($t = 1, 2, \ldots, T$) indicate respectively the individual country and the time period. Fixed country (α) and unit-specific trend effects are assumed, e_{it} is an error term, and δ_i is a country-specific deterministic trend effect. Pedroni (2004) developed asymptotic and finite–sample properties of the test statistic and proposed seven different statistics based on either the within-dimension approach or the between-dimension approach to examine the null hypothesis of non-cointegration in the panel.

In the presence of the unit-root variables, the effects of super-consistency may not dominate the endogeneity effects of the regressors when employing conventional OLS. To deal with the endogeneity bias in the regressors, we consider the DOLS estimation method. The DOLS method introduces a parametric bias correction, while the fully-modified OLS (FMOLS; Pedroni, 2000) uses non-parametric correction terms in the estimation to eliminate endogeneity bias. However, Kao and Chiang (2000) showed that the DOLS estimator outperforms the FMOLS estimator in the estimation of cointegrated panel regressions.

In our empirical analysis we implement two sets of panel cointegration test methods. The first set of tests is from Pedroni (1999, 2004), where the method involves the null hypothesis of no cointegration and uses the residuals derived from a panel regression to construct the test statistics and to determine the distributions. After appropriate standardization, the test statistics have an asymptotic distribution. Second, we perform Fisher's test to aggregate the p-values of the individual Johansen maximum likelihood cointegration test statistics (see Maddala and Wu, 1999). The Fisher test is a non-parametric method. When p_i denotes the p-value of the Johansen statistic for the ith country, we then have the statistic $-2\sum_{i=1}^{N} \log p_i \sim x^2_{2N}$. The test statistic is easy to compute and does not assume homogeneity in the coefficients of different countries.

3.4. The panel causality test

Following Engle and Granger (1987), we employ a two-step procedure to estimate a panel-based error correction model, which accounts for the long- and short-run dynamic relationships between variables. The first step estimates the long-run parameters in Eq. (6) in order to obtain the residuals corresponding to the deviation from equilibrium. The second step estimates the parameters related to the short-run adjustment. The resulting equations are used in conjunction with panel Granger causality testing:

$$\Delta RPGDP_{it} = \theta_{1i} + \lambda_1 ECT_{it-1} + \sum_{k=1}^{m} \theta_{11k} \Delta RPGDP_{it-k} + \sum_{k=1}^{m} \theta_{12k} \Delta RPLIP_{it-k} + v_{1it} \tag{7}$$

$$\Delta RPLIP_{it} = \theta_{2i} + \lambda_2 ECT_{it-1} + \sum_{k=1}^{m} \theta_{21k} \Delta RPGDP_{it-k} + \sum_{k=1}^{m} \theta_{22k} \Delta RPLIP_{it-k} + v_{2it}. \tag{8}$$

The term Δ denotes first differences, θ_{ji} ($j = 1, 2$) represents the fixed country effect, k ($k = 1, \ldots, m$) is the optimal lag length determined by the Schwarz Information Criterion, and ECT_{it-1} is the estimated lagged error correction term derived from the long-run cointegrating relationship of equation (6). The term λ_j ($j = 1, 2$) is the adjustment coefficient, and v_j ($j = 1, 2$) is the disturbance term assumed to be uncorrelated with mean zero. The short-run adjustment coefficients are constrained to be the same for all countries (Coiteux and Olivier, 2000; Lee, forthcoming). Moreover, a widely-used estimator for the system in Eqs. (7) and (8) is the dynamic panel GMM estimator proposed by Arellano and Bond (1991). This method resolves the possible simultaneity and heteroskedasticity problems.

The directions of causality can be identified by testing the significance of the coefficient estimate for each of the dependent variables in Eqs. (7) and (8). First, we consider the short-run effects to be transitory. For short-run causality, the following are tested: H_0:$\theta_{12k} = 0$ for Δ RPGDP for all k in Eq. (7) and H_0:$\theta_{21k} = 0$ for Δ RPLIP for all k in Eq. (8). Second, we evaluate long-run causality by looking at the estimate for the speed of adjustment parameter λ_j, which is the coefficient of the estimated error correction term ECT_{it-1}. The coefficients of the significance of ECT_{it-1} represent how rapidly deviations from the long-run equilibrium are eliminated following changes in each variable. Thus, we conduct a joint test of ECT_{it-1} and the respective interactive terms to check for long-run causality. The joint test shows which variables bear the burden of a short-run adjustment in order to re-establish a long-run equilibrium after a shock to the system.

4. Empirical results

4.1. Unit-root tests and the estimation of half-lives

This study uses panel data covering 41 countries for the period 1979–2007.[7] We obtain annual data for real GDP per capita from the *World Development Indicators* (WDI, 2007). The data for the real life insurance premiums per capita (i.e. life insurance density) are taken from various issues of *Simga*, a publication of the Swiss Reinsurance Company (1980–2010).[8] Insurance density indicates the average annual per capita premium within a country expressed in US dollars and shows how much on average each inhabitant of a country spends on insurance, although currency fluctuations can affect comparisons. Premiums per capita are converted using Purchasing Power Parity (PPP) values rather than US dollar exchange rates, and the PPP correction can be significant. All variables are expressed in logarithmic form, and the unit is expressed in constant 2000 US dollars.

Table 1 provides a summary of the life insurance sector of all 41 countries with average values of premium volume, market share, insurance density, and insurance penetration for the period 1979–2007. Premium volume represents life insurance premiums written in the reporting country and is a main indicator of the life insurance industry's importance in that country's economy. The market share of a country is the ratio of that country's premiums to the total premiums from all 41 countries. Table 1 shows that Japan ranked first in premium volume with a 32.59% share, followed by the United States (25.60%) and the United Kingdom (9.61%). Compared to the average values for the period 1979–2007, the statistics of life insurance density in 2007 indicate that recent life insurance market activities have risen, which can also be confirmed by the upward trend of insurance density from 1979 to 2007 in Fig. 1. According to Fig. 2, we find a positive relationship between life insurance density and real GDP per capita, and this relationship is roughly consistent with Enz's (2000) S-curve hypothesis for the insurance-income nexus. Countries at a high-income level have larger life insurance density than do countries with middle- and low-income levels, suggesting that as an economy develops, the development of life insurance markets matures.

Fig. 3 shows the growth rate trends of insurance density and real GDP, for which the economies of the countries roughly present stable growth during the period 1979–2007, while only occasionally turning negative. More specifically, the growth in life insurance density generally follows the growth pattern of GDP, with a higher magnitude under most occasions. For instance, GDP grew at an annual growth rate of above 10% in real terms, as did the growth of real life insurance density, which was over 30% in 1987. When comparing life insurance density across different income groups, the development process of the life insurance market can be divided into an approximate delineation of three rapid growth stages for life insurance. The rapid growth stage for life insurance, as described by Enz (2000), seems to be around the period 1981–1985 for the high-income countries, the period 1986–1998 for the middle-income countries, and after 1999 for the low-income countries. This indicates that the development of a life insurance market starts in high-income countries, then spreads to middle-income countries, and finally moves to low-income countries.

[7] The potential limitations of insurance data are that many countries have less than ten years of insurance data and that only 50 countries have insurance data covering about 30 years in Swiss Re's Sigma database. The lack of insurance data may not truly reflect the characteristics of global insurance market activities, making it unable for us to examine the content of life insurance, i.e. whole life insurance, universal life insurance, limited-pay life insurance, accidental death, annuity, and so on. Our study determines the number of countries and period by the availability of insurance and real GDP data. Since the SURADF test can just be used in the "*balance*" panel data, we select 41 countries to obtain the same estimated period for each country. Annual data are used in order to capture the long-run relationship between insurance and real GDP.

[8] According to Swiss Reinsurance Company (2008), life and non-life insurance activities are categorized according to standard EU and OECD conventions, whereas accident and health insurance are both counted as part of non-life insurance. Annuity is included when discussing life insurance. Therefore, life insurance premiums include both the businesses of life insurance and annuity.

Table 1
Life insurance in the 41 countries with average values for the period 1979–2007.

	Premiums (US$ millions)	Market share (%)	Rank	Insurance density	Insurance penetration	Insurance density in 2007
Panel A. Individual country						
Algeria	31	0.0029	41	1.28	0.0005	1.51
Argentina	886	0.0829	31	24.33	0.0041	52.40
Australia	13,698	1.2818	11	729.52	0.0329	1652.39
Austria	3768	0.3526	21	470.40	0.0185	1187.43
Belgium	8322	0.7787	16	805.16	0.0297	2860.14
Canada	17222	1.6116	8	578.11	0.0270	1375.03
Colombia	330	0.0309	35	8.08	0.0040	26.14
Denmark	5349	0.5005	20	1003.71	0.0303	3309.20
Egypt	135	0.0126	38	2.12	0.0022	6.84
Finland	6333	0.5926	19	1232.19	0.0491	3087.70
France	62396	5.8387	4	1055.86	0.0405	3073.00
Germany	47528	4.4474	5	584.94	0.0254	1240.90
Greece	889	0.0832	30	81.93	0.0063	276.50
India	7945	0.7435	18	7.74	0.0144	44.62
Indonesia	831	0.0777	33	4.01	0.0039	20.95
Ireland	8331	0.7796	15	2075.96	0.0718	12111.24
Israel	1785	0.1670	27	297.88	0.0201	591.08
Italy	36356	3.4020	6	479.15	0.0251	1486.41
Japan	348230	32.5856	1	2762.21	0.0835	2324.79
Kenya	68	0.0064	40	2.46	0.0059	6.74
Korea, South	29041	2.7175	7	629.44	0.0675	1660.04
Luxembourg	3332	0.3118	23	7436.66	0.1192	31369.21
Malaysia	1688	0.1579	28	74.66	0.0192	212.89
Mexico	2376	0.2223	26	24.64	0.0047	73.00
Morocco	186	0.0174	37	6.78	0.0052	23.20
Netherlands	15419	1.4429	9	981.23	0.0391	2205.71
New Zealand	858	0.0803	32	235.99	0.0179	241.94
Norway	3372	0.3156	22	754.42	0.0207	2431.01
Philippines	516	0.0483	34	7.44	0.0084	15.14
Portugal	2807	0.2627	24	270.07	0.0192	1221.78
Singapore	2418	0.2263	25	601.53	0.0278	2188.25
South Africa	12938	1.2107	12	312.90	0.0876	719.53
Spain	11033	1.0324	14	267.77	0.0160	709.62
Sweden	8066	0.7547	17	912.90	0.0322	2620.11
Switzerland	14019	1.3118	10	1978.78	0.0556	3168.58
Taiwan	12155	1.1374	13	547.40	0.0434	2169.71
Thailand	1437	0.1344	29	23.73	0.0110	70.83
Turkey	286	0.0268	36	4.21	0.0012	13.79
United Kingdom	102646	9.6051	3	1732.23	0.0728	6946.45
United States	273533	25.5959	2	1000.22	0.0353	1920.34
Venezuela	104	0.0098	39	5.50	0.0014	7.13
Panel B. Income group						
High income	39957.92	3.7391	–	1134.83	0.0395	3593.41
Middle income	2329.88	0.2180	–	56.95	0.0153	138.30
Low income	1588.29	0.1486	–	7.75	0.0073	26.90

Notes: These figures represent weighted averages. Premium volume represents life insurance premiums written in the reporting country. The share of a country is the ratio of that country's premiums to all countries' premiums taken together. Insurance density is calculated by dividing direct gross premiums by the population. Insurance penetration is the ratio of direct gross premiums to GDP.

Most countries' economic growth rates as well as the growth rates of insurance density fell in the early 1980s and early 1990s. This drop may reflect the energy crises triggered by the 1978 Iranian revolution, the 1990 Gulf War, and the failures from operating traditional life insurance products by U.S. life insurance firms during the late 1970s and early 1980s due to the recession of global economies. For Asian countries, i.e. Indonesia, Japan, and Thailand, their average economic growth rates fell to less

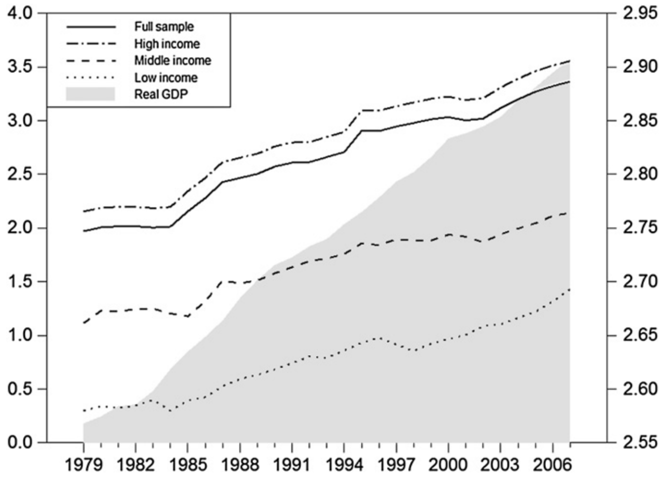

Fig. 1. The trend of insurance density under different income levels and real GDP (in logs).

than 1% during the period 1996–2000 (see Table A2 in Appendix), reflecting the 1997 Asian financial crisis. During this period, Asian insurance market activities were seriously impacted. Household purchasing power declined in many regional markets, and thus the growth rate of insurance density presents a decreasing trend. In Indonesia, the sales of new life insurance policies fell by 12.6% in 1998 and the cancellation rates of insurance contracts increased from 2.6% in 1997 to 12.8% in 1999. In 1998,

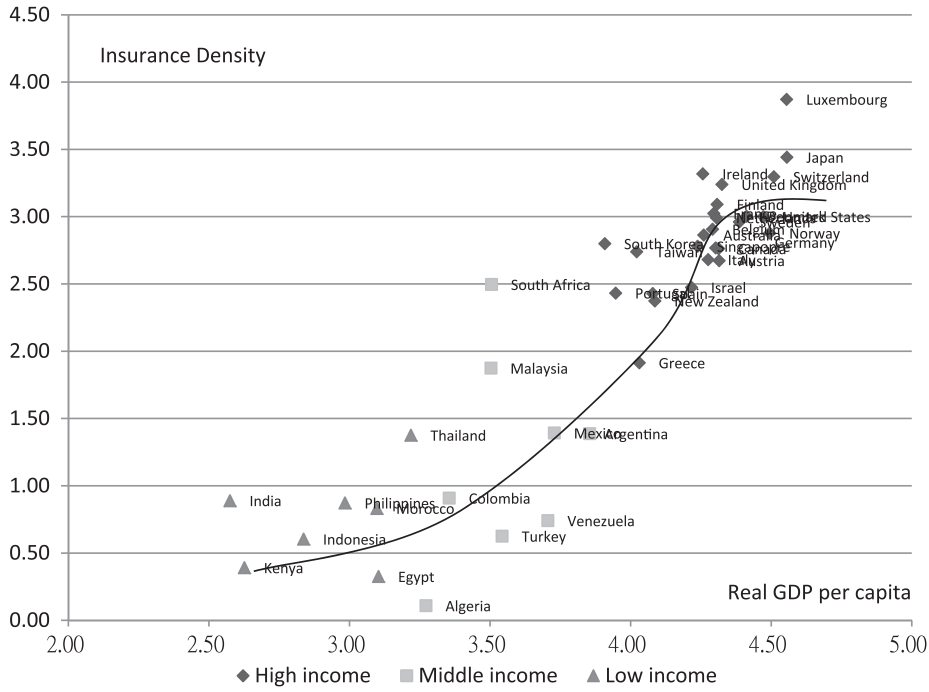

Fig. 2. The relationship between insurance density and real GDP per capita (in logs).

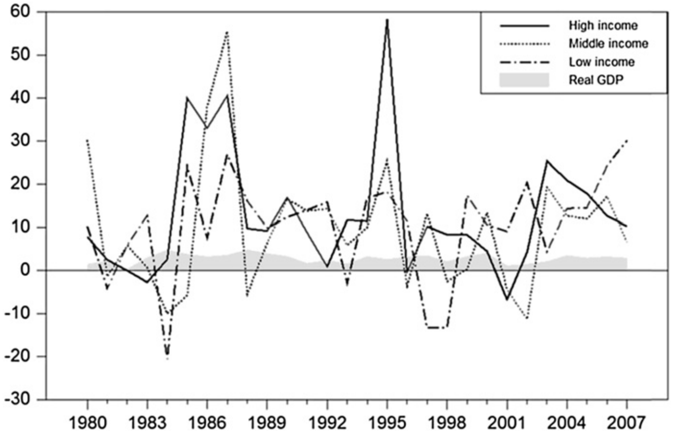

Fig. 3. The trend of insurance density under different income levels and real GDP (growth rates).

Thailand's life insurers suffered not only from a drop of 29% in new business, but also from a threefold increase in cancellations of insurance contracts. In 1999, Japan's central bank cut interest rates seven times, leading to large losses for life insurance companies, and thus five life insurance companies went bankrupt in Japan.

There are strong reasons, i.e. differences in mortality and income levels across countries, to believe that considerable heterogeneity exists in these countries under investigation, and thus the traditional panel unit roots applied herein may present misleading inferences. To provide the number of unit roots or stationary cross-sectional elements, we utilize the Breuer et al. (2001, 2002) test that individually examines the null of non-stationarity in the SUR grid. Tables A3 and A4 provide the results of the panel SURADF unit-root tests and the critical values for three different levels of income (high-, middle-, and low-income groups). As the SURADF test has non-standard distributions, the critical values need to be obtained via simulations. In the simulation's data generation phase, the intercepts and the coefficients on the lagged values for each series are set equal to zero. We obtain the estimated 1%, 5%, and 10% critical values from Monte Carlo simulations based on 29 observations for each series and 10000 replications by using the lag and covariance structure from the panel of life insurance premiums. Since the SUR estimation takes into account error correlation, which is different for different data series, the critical values for the Panel SURADF are different for each series.

We report the results for three sub-panels classified according to World Bank estimates of 2008 GNI per capita (see World Bank Atlas method). The groups are: low income, US$3855 or less; middle income, US$3856–$11905; and high income, US$11906 or more. Our results reveal that the two variables have a mixture of I(0) and I(1) processes. According to Table A3, the null hypothesis of non-stationarity for real life insurance premiums per capita can be rejected for Argentina, Colombia, Israel, Japan, Mexico, New Zealand, and South Korea. Table A4's results indicate that except for Italy and Japan, the real GDP per capita is stationary, i.e., I(0). We summarize the results of the panel unit-root tests for RPGDP and RPLIP based on different I(1)/I(0) properties in Table A5.

The unit-root tests alone may not be sufficient to justify the adjustment dynamics of a long-run equilibrium for insurance premiums and are uninformative as to the degree of mean reversion. Nevertheless, it is likely that although the unit-root hypothesis is rejected, deviations are still persistent. What we are interested in is the variable for the speed of convergence to the long-run equilibrium. One measure of the degree of mean reversion that has attracted much attention in the literature is the half-life. Recent point estimates of the half-lives alone provide an incomplete description of the speed of convergence toward the long-run equilibrium. To this end, we compute the corresponding confidence intervals to provide better indications of uncertainty around the estimates of half-lives.

Table 2 presents the half-lives and their confidence intervals, showing that the half-lives of real insurance premiums per capita (Panel A) have an extensive range from 1.24 to 6.40 years among high-

Table 2
Estimated half-lives and confidence intervals.

Country	β	Estimated half-life (year)	Confidence interval at 95%
Panel A. Real per capita of life insurance premiums			
(i) High income			
Israel	−0.1909	3.27	[0.75, 5.79]
Japan	−0.1026	6.40	[0, 18.68]
Korea, South	−0.1332	4.85	[1.01, 8.69]
New Zealand	−0.4277	1.24	[0, 2.49]
(ii) Middle income			
Argentina	−0.1298	4.98	[0, 10.78]
Colombia	−0.1505	4.25	[0, 11.72]
Mexico	−0.1596	3.99	[0.20, 7.78]
Panel B. Real GDP per capita			
(i) High income			
Italy	−0.1341	4.81	[0,14.37]
Japan	−0.0943	7.00	[0, 15.41]

Notes: Rossi (2005) proposed the method of estimation of the confidence intervals for half-lives; see Rossi (2005) for more details.

income level countries versus a narrower range (3.99–4.98 years) among middle-income level countries. For real GDP per capita (Panel B), the half-lives in Italy and Japan are respectively 4.81 and 7 years. The results imply that the degree of mean reversion is small for high-income countries, suggesting that rich countries' economies grow slowly, which is consistent with the conditional convergence hypothesis widely suggested in the literature (Ford et al., 2008; Fung, 2009).[9] As Romer (1987) pointed out, the reason for economic convergence can be attributed to the law of diminishing returns to capital. In sum, the results of the half-lives roughly reflect Enz's (2000) S-curve hypothesis, which states that at higher levels of income insurance market activity becomes less sensitive to income growth, hence implying varying income elasticities. In middle-income countries, the speeds of both insurance and economic development are faster than those in high-income countries.

4.2. The panel cointegration and causality test results

Through our SURADF test results, Table A5 summarizes the results of the panel unit-root tests for RPGDP and RPLIP. Among them, both RPGDP and RPLIP are I(0) in Japan, while the two series have different properties in 7 countries. Both series are I(1) in 33 countries, and we thus conduct panel cointegration tests for these 33 countries.

We follow Ward and Zurbruegg (2000) and proceed to test the two variables for cointegration in order to determine whether there is a long-run relationship to control for in the econometric specification. Table 3 reports the results of the Pedroni (2004) heterogeneous panel tests (the corresponding p-values are reported in parentheses) for Eq. (6). Except for the group r-statistic, the null hypothesis of no cointegration is rejected in six cases at the 5% significance level.[10] The results of the Johansen-Fisher panel cointegration test in Table 4 are fairly conclusive: Fisher's tests (regardless of the trace test statistics or max-eigen test statistics) support the presence of a cointegrating relationship between the two variables. After analyzing sensitivity and robustness, the test results show that there is a panel cointegration relationship between the two variables for all model specifications. Therefore, the next step is to estimate this relationship.

Table 5 provides the country-by-country results based on the panel DOLS model. Eq. (6) is estimated through the panel DOLS, where the dependent variables are RPGDP. As shown at the bottom of Table 5, the panel parameter is 0.0625 for real life insurance premiums, and because the cointegration

[9] Fung (2009) demonstrated that the mutually reinforcing relationship between financial development and economic growth is stronger in the early stage of economic development, and this relationship decreases when supported economic growth gets under way.

[10] Pedroni (1999) indicated that the panel-ADF and group-ADF tests have better small-sample properties than the other tests, and hence they are more reliable.

Table 3
Pedroni's residual cointegration test results.

Model	Test statistic	Prob.
Panel v-stat	2.556**	0.005
Panel r-stat	−3.511**	0.000
Panel PP-stat	−3.831**	0.000
Panel ADF-stat	−5.630**	0.000
Group r-stat	−1.299	0.096
Group PP-stat	−3.270**	0.001
Group ADF-stat	−5.387**	0.000

Notes: The null hypothesis is that the variables are not cointegrated. Under the null tests, all the statistics are distributed as normal (0, 1). ** indicates that the parameter is significant at the 5% level.

Table 4
Panel cointegration test results of a Fisher-type test using an underlying Johansen's methodology.

	Fisher stat. (From trace test)	Prob.	Fisher stat. (From max-eigen test)	Prob.
None	137.3**	0.000	137.3**	0.000
At most 1	66.94	0.445	66.94	0.445

Notes: Asymptotic p-values are computed using a Chi-square distribution. ** indicates that the parameters are significant at the 5% level. Fisher's test applies regardless of the dependent variable.

Table 5
Dynamic OLS estimates (dependent variable is RPGDP).

Country	RPLIP	
	Coefficient	t-statistic
Algeria	0.106**	26.424
Australia	−0.074**	−3.127
Austria	−0.025**	−2.190
Belgium	−0.028**	−7.359
Canada	0.451**	3.908
Denmark	−0.082**	−5.049
Egypt	0.010**	2.064
Finland	−0.109	−1.489
France	−0.064**	−6.279
Greece	−0.357	−1.651
Germany	−0.039	−0.659
India	0.026	0.403
Indonesia	0.134	1.226
Ireland	0.364**	15.325
Kenya	0.136**	9.030
Luxembourg	0.069**	8.158
Malaysia	0.347**	12.384
Morocco	−0.252**	−6.197
Netherlands	0.001	0.041
Norway	0.093	1.247
Philippines	0.177**	8.182
Portugal	0.036**	4.137
Singapore	0.187**	30.185
South Africa	0.454**	10.816
Spain	0.027**	3.582
Sweden	0.015	0.151
Switzerland	−0.244**	−6.124
Taiwan	0.186**	2.083
Thailand	0.473**	43.833
Turkey	0.006	1.770
United Kingdom	0.010	0.571
United States	−0.025	−0.270
Venezuela	0.056**	10.208
Panel	**0.0625****	**27.087**

Notes: The asymptotic distribution of the t statistic is standard normal as T and N go to infinity. ** indicates that the parameters are significant at the 5% level.

coefficients are statistically significant at the 5% level, the effect is positive, showing that a 1% increase in life insurance premiums raises real GDP by around 0.06%. On a per country basis, RPLIP has a significantly positive impact on RPGDP in 15 of the 33 countries at the 5% level. Furthermore, our results indicate a significantly negative impact of RPLIP on RPGDP in 7 of the 33 countries. Since insurance markets are prone to moral hazard, adverse selection problems, and high information and transaction costs, the losses from rent collection in these 7 countries may outweigh the efficiency gains. Therefore, a cointegrating relationship might still exist even though it is masked by these other factors.

The DOLS estimates of the positive coefficients of life insurance premiums with respect to real GDP range from 0.010 (Egypt) to 0.473 (Thailand). As for the majority of time series-based tests, after factoring in that the panel unit-root test and cointegration test utilize the data in a more efficient way, the panel results provide clear evidence that there is a fairly strong long-run relationship between real GDP and real insurance premiums. When comparing our findings with those of previous works, we are able to support Ward and Zurbruegg (2000) and Beck and Webb (2003) with rather positive findings – that is, there is a positive impact of life insurance on economic growth. This result favors the supply-leading hypothesis that states growth in the insurance sector can stimulate different countries' economic growth.

Once the two variables are cointegrated, we perform a panel-based error correction model to examine the short-run and long-run causalities between life insurance and economic growth. Table 6 shows the *F*-test results of the panel causality tests for both the long-run and short-run for Eqs. (7) and (8). The results present that there are both long-run and short-run bidirectional causalities between life insurance markets and economic growth, indicating that a high level of economic growth leads to a high level of insurance premiums and vice versa.

Our result of the panel causality test supports the viewpoint of Patrick (1966), who argued that the direction of causality between financial activities and economic growth varies over the course of an economy's development. Before sustained modern economic growth gets under way, financial development can induce real innovation-type investment and lead to economic growth (the supply-leading phenomenon). When real economic growth occurs, the supply-leading impetus gradually becomes less important and the demand-following financial response turns dominant. Thus, in the short-run, economic growth is supply-led through growth in life insurance activities, while there is an interaction of supply-leading and demand-following phenomena for the life insurance-economic growth nexus.

4.3. Further discussions and implications

Many previous studies on the life insurance-growth nexus rely only on a single country or on a few countries over fairly short or distant time spans. Not surprisingly, the results on a country-by-country basis are ambiguous and depend on country-specific circumstances. Unlike those studies, our investigation uses the panel data model to obtain additional information. In fact our work provides an individual country's life insurance-growth nexus via a panel data model. From this, we obtain more accurate estimations, which are substantial in the academic literature and which can provide effective policies for the development of the life insurance market.

There are four main reasons for the importance of testing the unit-root null hypothesis and long-run cointegration. First, this hypothesis test proffers knowledge on whether a shock has a permanent or a transitory effect on real life insurance premiums and real GDP. Rejecting the unit-root null implies that the series is either mean reverting or trend reverting. Over the time period for which the null is rejected, any economic policies and/or financial news are interpreted as having a transitory effect on the two series. Second, the unit-root hypothesis has clear implications for econometric modeling. One branch of the literature has examined the long-run relationship between insurance premiums and real GDP. In examining such a relationship using cointegration analysis, a pre-requisite is that both variables contain a unit-root – in other words, accepting the unit-root null hypothesis is a primary step toward conducting a test for a long-run relationship. Third, the unit-root null hypothesis has direct implications for the convergence hypothesis of real insurance premiums (or real income), meaning that in a time-series framework the difference in insurance premiums (or real income) between any two countries should not contain a unit root. Fourth and finally, if the two series are erroneously treated as I(1) and the causality tests for insurance premiums and macroeconomics are applied to the first difference, then the result will entail spurious causality.

Table 6
Panel causality test results.

Dependent variable	Source of causation (independent variable)			
	Short-run		Joint (short-run/long-run)	
	ΔRPGDP	ΔRPLIP	ΔRPGDP, ECT	ΔRPLIP, ECT
ΔRPGDP	–	10.234**	–	39.893**
ΔRPLIP	10.133**	–	35.629**	–

Notes: Figures denote *F*-statistic values. ECT indicates the estimated error-correction terms. ** indicates that the parameter is significant at the 5% level.

We apply the newly-developed panel SURADF approach to identify which series within the panel contain a unit root and which can confirm the stationary process for both real insurance premiums per capita and real GDP per capita. Therefore, we can draw more accurate inferences when using the panel cointegration method to evaluate whether real GDP and real insurance premiums are cointegrated for countries in which the two series follow an I(1) process. This avoids inappropriate panel cointegration analysis when we identify the two series as having different orders of integration. Our empirical work therefore draws a clear and precise picture of the relationship between the life insurance market's development and economic growth.

Our empirical results with respect to the insurance–growth relationship indicate that life insurance premiums have a significantly positive impact on real GDP. This conforms to the theory in the literature which indicates that life insurance contributes to economic growth by improving the sound functions of financial systems, as noted in Ward and Zurbruegg (2000), Haiss and Sümegi (2008), Han et al. (2010), and Lee (forthcoming).

The empirical evidence based on the causality test supports not only the supply-leading hypothesis, but also the demand-following hypothesis. The results suggest that insurance markets and economic growth are endogenous, and therefore any single equation forecast of one or the other could be misleading. At the same time, economic growth in the long-run must be based on an effective insurance market, and economic growth can facilitate contiguous development of an insurance market. Furthermore, life insurance activities and real GDP are both endogenous variables, meaning that they mutually influence each other in the long-run. Thus, any previous studies that focused on unidirectional analysis – namely, either on life insurance demand (Beck and Webb, 2003) or economic growth – ignored another direction in the insurance-growth nexus. We conclude that there exists a theoretical view for life insurance premiums influencing economic growth and vice versa, which is consistent with the works of Ward and Zurbruegg (2000) for Australia, Canada, Italy, and Japan and Webb et al. (2005). This empirical finding that life insurance premiums and economic growth reinforce each other has important implications for the conduct of economic or financial policies.[11]

5. Conclusions

This paper's techniques are econometrically a significant improvement over existing studies on the link between life insurance activities and economic growth. This paper adopts the panel SURADF unit-root test of Breuer et al. (2001, 2002) to re-examine 41 countries' life insurance premiums and real GDP within high-, middle-, and low-income panel sets following the classification criterion of the World Bank and covering 1979–2007. Our results illustrate that the per capita real insurance premiums and

[11] Since life insurance market activities and economic development are not independent of each other, countries cannot make economic or financial policies that do not affect the other sector. Policies, for example, favoring the growth of life insurance activities positively influence economic growth. Thus, when making economic or financial policies, policy-makers should simultaneously consider the impacts of these policies on the developments of the insurance market and the economy.

real GDP per capita in the sample countries are a mixture of I(0) and I(1) processes, and that the generally used panel root tests could lead to misleading inferences.

For the persistence measures, the half-lives of real insurance premiums per capita tend to cluster in the range of 1.24–6.40 years among high-income level countries, or much slower than the range of 3.99–4.98 years among middle-income level countries. For real GDP per capita the half-lives in Italy and Japan are respectively 4.81 and 7 years. These results suggest that high-income countries' economies grow slower, which is consistent with the convergence hypothesis.

The panel cointegration test results herein provide substantive evidence of a fairly strong long-run cointegration relationship between real GDP and real life insurance premiums. The long-run estimated panel regression parameter results indicate that a 1% increase in real life premium raises real GDP by 0.06%.

Finally, the causality results signify that there is a positive bi-causal relationship between the level of economic activity and life insurance markets in the long-run. In this sense, a high level of economic growth leads to a high insurance premium level and vice versa. In high-income countries, there seems to be a tendency to depend on insurance markets and that a sufficiently large amount of insurance activity seems to ensure a higher level of economic growth.

This paper provides an empirical justification for strengthening both the insurance market and economic development. First, to achieve sustainable economic growth, it is advantageous to further reform the insurance market in order to improve information flow and to enhance competition. Second, to take advantage of the positive interaction between insurance market activities and economic growth, authorities should liberalize the economy while liberalizing the insurance sector – that is to say, strategies that promote development in the real economy should also be emphasized (Calderon and Liu, 2003; Lee, forthcoming). If policy-makers wish to promote economic growth, they should focus attention on long-run policies – for example, the creation of modern financial institutions, i.e. banks, insurance firms, and so on, that will foster the most efficient allocation of resources through better risk management. Countries will benefit from strengthening their regulatory framework by creating a sound environment that facilitates insurance markets' development, which further stimulates economic growth.

Acknowledgements

We would like to thank the Editor, Professor Lothian, Professor McCarthy, and the anonymous referees for helpful comments and suggestions.

Appendix

See Tables A1–A5.

Table A1
List of countries classified into the three income groups for 41 countries.

High-income countries (26 countries)	Middle-income countries (8 countries)	Low-income countries (7 countries)
Australia	Algeria	Egypt
Austria	Argentina	India
Belgium	Colombia	Indonesia
Canada	Malaysia	Kenya
Denmark	Mexico	Morocco
Finland	South Africa	Philippines
France	Turkey	Thailand
Germany	Venezuela	
Greece		
Ireland		
Israel		
Italy		
Japan		
Korea, South		
Luxembourg		

(*continued on next page*)

Table A1 (*continued*)

High-income countries (26 countries)	Middle-income countries (8 countries)	Low-income countries (7 countries)
Netherlands		
New Zealand		
Norway		
Portugal		
Singapore		
Spain		
Sweden		
Switzerland		
Taiwan		
United Kingdom		
United States		

Notes: Classified according to World Bank estimates of 2008 GNI per capita. The GNI of the high- and low-income countries are above US$11906 and below US$3855, respectively. The GNI of the middle-income countries falls in-between these two figures.

Table A2
The 5-year average rates of economic growth of the 41 countries.

Country	1979–1985	1986–1990	1991–1995	1996–2000	2001–2007
Algeria	4.62	0.78	0.28	3.14	4.21
Argentina	0.32	−0.33	6.70	2.66	4.12
Australia	3.04	3.92	2.44	4.32	3.26
Austria	2.07	2.89	2.07	2.99	2.19
Belgium	1.65	3.09	1.60	2.86	1.93
Canada	2.85	2.89	1.75	4.14	2.54
Colombia	2.96	4.95	4.14	1.25	4.53
Denmark	2.45	1.46	2.36	2.86	1.63
Egypt	7.12	4.23	3.41	5.20	4.51
Finland	3.75	3.39	−0.54	4.81	3.27
France	1.85	3.27	1.16	2.81	1.85
Germany	1.77	3.31	2.22	2.01	1.29
Greece	0.67	1.27	1.26	3.45	4.21
India	3.90	5.97	5.11	5.84	7.72
Indonesia	6.31	7.14	7.87	0.99	5.07
Ireland	2.71	4.75	4.67	9.62	5.54
Israel	4.00	4.34	6.50	5.05	3.13
Italy	2.55	3.14	1.28	1.91	1.14
Japan	4.24	5.01	1.41	0.97	1.56
Kenya	3.69	5.64	1.61	2.16	4.51
Korea, South	6.35	9.65	7.82	4.55	4.68
Luxembourg	2.25	7.50	3.98	6.16	4.29
Malaysia	6.08	6.91	9.47	4.99	5.16
Mexico	4.15	1.72	1.61	5.46	2.51
Morocco	3.63	4.51	1.13	3.95	5.06
Netherlands	1.57	3.35	2.30	4.05	1.95
New Zealand	2.41	0.82	3.15	2.65	3.33
Norway	3.66	1.71	3.73	3.69	2.30
Philippines	0.73	4.74	2.19	3.96	4.98
Portugal	2.11	5.68	1.73	4.22	1.13
Singapore	7.34	8.53	8.87	6.40	6.47
South Africa	2.49	1.68	0.89	2.80	4.32
Spain	1.32	4.50	1.52	4.11	3.42
Sweden	2.11	2.58	0.72	3.53	3.01
Switzerland	2.10	2.92	0.10	2.05	1.98
Taiwan	6.78	8.88	7.23	5.25	4.23
Thailand	5.40	10.34	8.62	0.64	5.09
Turkey	3.04	5.67	3.32	4.13	5.01
United Kingdom	1.61	3.33	1.66	3.44	2.56
United States	2.75	3.22	2.54	4.35	2.38
Venezuela	−1.17	2.76	3.53	0.84	4.77

Notes: The values are measured in percentage terms.

Table A3
Panel SURADF tests and critical values for RPLIP.

Country panel label	SURADF	Critical values		
		0.01	0.05	0.1
Part A: High income				
Australia	−1.291	−3.865	−3.255	−2.934
Austria	−4.552	−5.653	−4.981	−4.646
Belgium	−1.794	−4.571	−3.957	−3.607
Canada	−0.259	−4.737	−4.155	−3.776
Denmark	−2.063	−4.991	−4.293	−3.950
Finland	−3.685	−5.106	−4.459	−4.084
France	−3.819	−5.260	−4.655	−4.285
Germany	−3.986	−5.143	−4.479	−4.118
Greece	−1.519	−4.628	−4.005	−3.658
Ireland	−1.928	−4.619	−3.968	−3.625
Israel	−11.001***	−4.380	−3.807	−3.469
Italy	−2.950	−4.568	−3.909	−3.591
Japan	−4.255***	−4.239	−3.627	−3.294
Korea, South	−6.246***	−4.423	−3.809	−3.470
Luxembourg	−0.082	−4.029	−3.466	−3.122
Netherlands	−2.181	−5.100	−4.454	−4.081
New Zealand	−4.312**	−4.657	−3.957	−3.596
Norway	−0.467	−4.216	−3.601	−3.235
Portugal	0.406	−4.777	−4.133	−3.768
Singapore	−1.342	−4.454	−3.789	−3.440
Spain	−1.976	−4.544	−3.945	−3.605
Sweden	−1.830	−4.593	−3.898	−3.531
Switzerland	−3.658	−5.625	−4.891	−4.497
Taiwan	−2.638	−4.293	−3.687	−3.326
United Kingdom	−0.670	−4.522	−3.882	−3.536
United States	−3.133	−4.069	−3.512	−3.170
Part B: Middle income				
Algeria	−2.816	−3.893	−3.287	−2.964
Argentina	−4.535***	−3.908	−3.365	−3.063
Colombia	−3.854**	−3.871	−3.290	−2.963
Malaysia	−0.839	−3.481	−2.934	−2.641
Mexico	−9.728***	−4.433	−3.839	−3.494
South Africa	−1.865	−4.036	−3.458	−3.142
Turkey	−0.711	−3.833	−3.256	−2.951
Venezuela	−2.145	−4.041	−3.416	−3.098
Part C: Low income				
Egypt	−1.732	−3.586	−3.002	−2.703
India	3.254	−3.707	−3.090	−2.793
Indonesia	−1.261	−3.578	−3.034	−2.752
Kenya	−2.081	−3.581	−3.002	−2.723
Morocco	0.874	−3.618	−3.064	−2.774
Philippines	−1.887	−3.596	−2.991	−2.686
Thailand	−1.984	−3.767	−3.188	−2.880

Notes: *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. Critical values are calculated using the Monte Carlo simulation with 10,000 draws, tailored to the present sample size (For details of this simulation, see Breuer et al., 2001).

Table A4
Panel SURADF tests and critical values for RPGDP.

Country panel label	SURADF	Critical values		
		0.01	0.05	0.1
Part A: High income				
Australia	−0.427	−4.529	−3.908	−3.553
Austria	−1.560	−4.840	−4.190	−3.826
Belgium	−2.226	−6.055	−5.406	−5.060
Canada	0.203	−4.125	−3.535	−3.221
Denmark	−2.294	−4.565	−3.873	−3.534
Finland	−0.079	−4.623	−3.982	−3.604
France	−1.832	−5.427	−4.761	−4.393
Germany	−2.622	−5.354	−4.764	−4.410
Greece	4.418	−5.050	−4.399	−4.023
Ireland	−0.660	−4.607	−4.001	−3.658
Israel	−0.322	−4.629	−4.013	−3.660
Italy	−4.798**	−5.239	−4.633	−4.273
Japan	−4.204**	−4.787	−4.164	−3.820
Korea, South	−3.284	−4.442	−3.752	−3.438
Luxembourg	−1.422	−4.226	−3.628	−3.290
Netherlands	−0.076	−5.089	−4.491	−4.122
New Zealand	−0.049	−4.436	−3.811	−3.473
Norway	−1.266	−4.560	−3.876	−3.540
Portugal	−1.098	−5.481	−4.787	−4.389
Singapore	−1.259	−4.513	−3.899	−3.568
Spain	−0.207	−5.511	−4.866	−4.494
Sweden	−0.053	−4.540	−3.911	−3.564
Switzerland	0.244	−4.653	−4.047	−3.706
Taiwan	−3.631	−4.790	−4.169	−3.800
United Kingdom	−0.813	−4.203	−3.607	−3.256
United States	−1.423	−5.313	−4.662	−4.325
Part B: Middle income				
Algeria	−1.306	−3.737	−3.191	−2.862
Argentina	−0.877	−3.845	−3.269	−2.963
Colombia	−0.043	−3.967	−3.413	−3.102
Malaysia	−1.183	−3.647	−3.104	−2.815
Mexico	−0.541	−3.669	−3.114	−2.808
South Africa	−0.580	−4.100	−3.450	−3.104
Turkey	−1.221	−3.770	−3.210	−2.927
Venezuela	−1.710	−4.097	−3.498	−3.178
Part C: Low income				
Egypt	−0.925	−3.837	−3.219	−2.905
India	2.746	−4.116	−3.568	−3.240
Indonesia	−1.511	−4.009	−3.415	−3.108
Kenya	−0.563	−4.054	−3.473	−3.179
Morocco	0.435	−3.740	−3.171	−2.860
Philippines	−0.179	−3.895	−3.309	−2.982
Thailand	−2.769	−4.165	−3.630	−3.295

Notes: *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. Critical values are calculated using the Monte Carlo simulation with 10,000 draws, tailored to the present sample size (For details of this simulation, see Breuer et al., 2001).

Table A5
List of selected countries by different properties of both variables.

Both series are I(1): 33 Countries	Both series have different properties or are I(0): 8 Countries
Algeria	Argentina
Australia	Colombia
Austria	Israel
Belgium	Italy
Canada	Japan
Denmark	Korea, South
Egypt	Mexico
Finland	New Zealand
France	
Greece	
Germany	
India	
Indonesia	
Ireland	
Kenya	
Luxembourg	
Malaysia	
Morocco	
Netherlands	
Norway	
Philippines	
Portugal	
Singapore	
South Africa	
Spain	
Sweden	
Switzerland	
Taiwan	
Thailand	
Turkey	
United Kingdom	
United States	
Venezuela	

Notes: Classified according to the results of Tables A3, A4.

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