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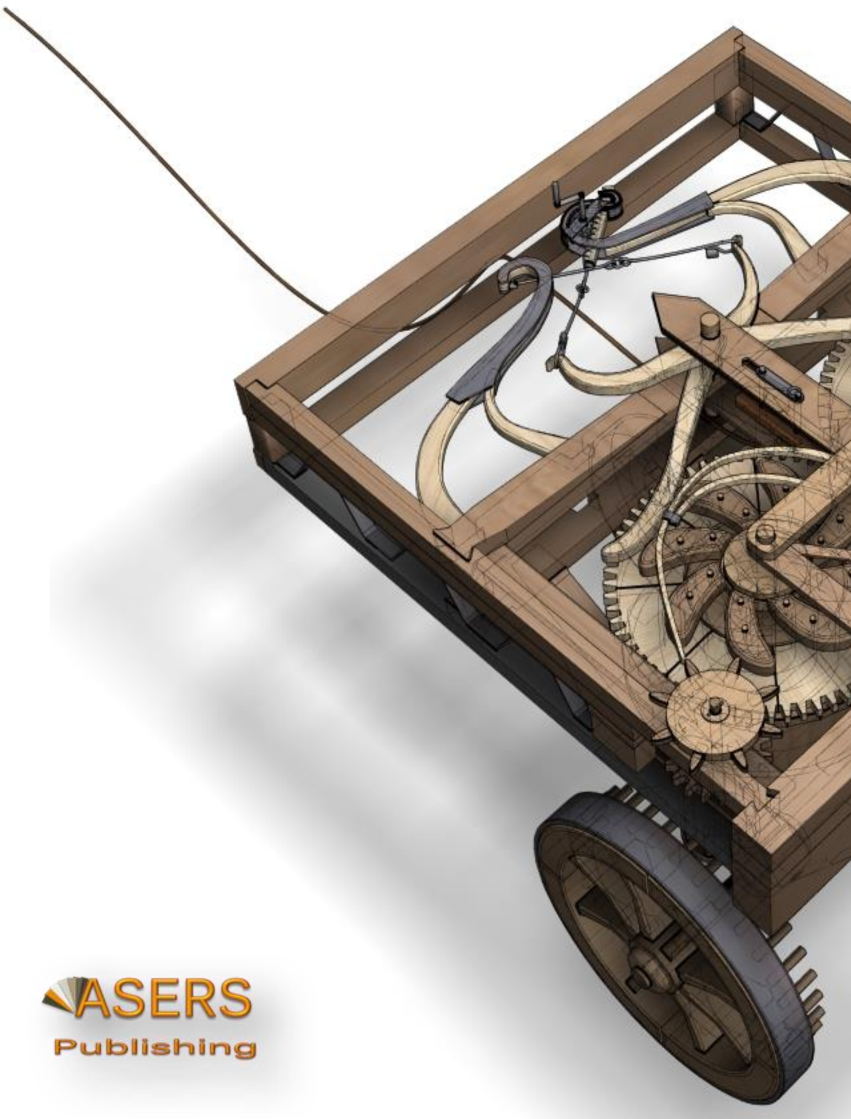

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GENERALIZED FISHER HYPOTHESIS VALIDITY FOR CANADA, UK, AND SUISSE STOCK MARKETS: EVIDENCE FROM PANEL ARDL MODELS

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Abstract:

In this paper we propose a decision support tool for the investor in terms of asset allocation. The key question is to know whether equities are perfect hedge against inflation if either we invest in only one market or if we go to all the considered markets. We chose three democratic countries having common monetary policy based on the Inflation rate stabilization targeting (including Canada, UK, and Suisse) over the period 1999M01-2018M04. We see how the stock return evolution is related to inflation rate Pre, during, and Post 2008 Global financial crisis (GFC). Then, some dynamic version of the Generalized Fisher hypothesis (GFH) models are explored by some univariate and panel autoregressive dynamic linear (ARDL) frameworks. We conclude that during crisis period, being on either Suisse or Canadian stock market, investors can have important abnormal gains. Then including the UK in a portfolio allows investors to limit losses caused by inflation in the UK stock market alone.

Keywords: GFH; GFC; panel and univariate ARDL models; MG; PMG; Canada; UK; Suisse.

JEL Classification: C23; G00; G14; G15.

Introduction

The original hypothesis that is attributed to the monetarist, Irvin Fisher offers the first preliminary study towards formalizing the relationship between asset returns and inflation. Fisher hypothesis assumes that nominal interest rate is expressed as the sum of real return and inflation rate.[2] Fisher, (1930) hypothesized that the expected real interest rate is determined by real factors and is *independent* of the *expected inflation rate*. This hypothesis was generalized to asset in the efficient stock markets context (Fama and Schwert 1977).

[2] Fisher (1930) asserted that the "nominal" interest rate consists of a "real" rate plus the expected inflation rate.

The generalized Fisher hypothesis (GFH) assumes independence between the expected real return and inflation. Invalidity of the GFH, that real returns on *financial assets* are likely to be dependent of inflation rates, has some implications. The more important implication is the uncertainty creation across financial markets, thereby adversely affecting investment and saving decisions in an economy.

The Fisher hypothesis has become the workhorse for motivating the inflation hedging question of any asset class including commodities (Arnold and Auer, 2015). However, existing empirical research on the relationship between stock returns and expected inflation hasn't reached a consensus yet.

The generalized Fisher hypothesis assumes the *independence* between the expected real return and inflation and a *positive* relationship between nominal stock returns and expected inflation. These conditions have been extensively explored for developing and advanced economies over the past three decades (Lintner 1973; Fama 1981; Geske and Roll 1983; Basse and Reddemann 2011; Arnold and Auer 2015; Baker and Jabbouri 2016; Baker and Jabbouri 2017; Adekoya, *et al.* 2021; and Sangyup and Junhyeok 2022). Some studies highlighted the existence of *positive and/or negative* associations (Hardin, *et al.* 2012; Hoesli, *et al.* 1997; Barnes, *et al.* 1999; Lee and Lee 2012), while others have detected *only a negative* relationship (Chatrath 1997 and Maysami and Koh 2000).

Two other important questions on the correlation between real stock returns and inflation rates are treated in the literature. The first is about the sign and the strength of the correlation that may depend on the frequency scale (price level vs index level). The second is about how the correlations can evolve heterogeneously overtime (Valcarcel 2012 and Antonakakis, *et al.* 2017).

Previous studies have dealt with different models and inferential (estimation and test) approaches in order to detect and explain the hedging inflation ability. Recently, for *robustness question*, the panel data-based approach was used in a few number of papers (Afees, *et al.* 2020; Afees, *et al.* 2019; and Halit 2016). For example, Afees *et al.* (2019) found that the GFH test results based on panel data (the price level data for the individual constituents of US stock returns) were opposite to those based rather on the index level data (univariate time series).

In this paper, GFH test will be verified within the Panel type data. We consider three developed countries having in common a monetary policy based on inflation rate targeting stabilization including Canada, the UK, and Suisse stock markets for the period from 1999M01 to 2018M04 covering 2008 GF crisis. The objective is to examine the inflation-hedging ability within each stock market and within the panel data of the considered three markets. We want to know if hedging ability results from each stock markets may be different from ones of the portfolio asset from the three stock markets. In addition, since the long run relationship between stock return and inflation can be instable through time, the analysis will be done for the following four periods: the full data set and the three sub periods: Pre the Global Financial Crisis (GFC), during the GFC period, and Post the GFC period. To the best of our knowledge, our paper is the first which uses a univariate and panel ARDL approaches to explore the GFH relationship that examining the inflation-hedging ability.

This study is organized as follows. After introduction, we give an empirical literature review. We mention then the required data and their sources and we give some descriptive analysis and present data analysis. After that, we outline the methodology used and we provide the empirical results and discussion. Concluding remarks will be given at the end.

1. Literature Review

During the 1970s, new evidence contradicted the economic GFH. More specifically, (Nelson 1976; Bodie 1976; Fama and Schwert 1977; and Modigliani and Cohn, 1979) reported a negative relationship between stock returns and inflation. Later, from the consequence of proxy hypothesis effects, Fama (1981) concluded also for the negative correlation between stock returns and inflation.

This proxy hypothesis garnered substantial support in some subsequent papers (Gultekin 1983; Geske and Roll 1983; and Erb *et al.* 1995).

The negative relationship between real stock returns and inflation rates has also been explained by four theories based on four hypotheses including Money Illusion Hypothesis (MIH), Tax Effect Hypothesis (TEH), Proxy Effect Hypothesis (PEH), and Reverse Causality Hypothesis (RCH) (Tiwari, *et al.* 2019).

The positive relationship between nominal stock returns and inflation rates was also explained by the Wealth Effect Hypothesis (WEH) since real stock returns can effect inflation rates through their impact on consumption and then on aggregate demand (Ando and Modigliani 1963). According to WEH, there are different channels through which stock prices can affect consumption such as the realized gain (higher future income and wealth) via the expectation that raising the current stock price, the liquidity constraint effect, and the stock option value effect. Based on these two hypotheses [GFH and WEH], a positive relationship between nominal stock returns and inflation rates can be observed in the data.

Empirically, the relationship between (nominal or real) stock returns and inflation has been analyzed in the literature for *short or long horizons*. For *short-run*, many have found a negative correlation (Bodie 1976; Fama and Schwert 1977; Fama 1981; Ghazali and Ramlee 2003; Koustas and Lamarche 2010; and Tsong and Lee 2013), while for *long-run*, the results are more likely to support the Fisher hypothesis (Schotman and Schweitzer 2000 and Lothian and McCarthy 2001).

More recent studies are based on recent models and techniques in order to detect the hedging inflation ability such as the NARDL model (Thi, *et al.* 2016), the time variation investigation (Salisu, *et al.* 2019 and Kuang 2017), the cointegration tests (Al-Nassar and Bhatti 2019), the comparative analysis (Akinsomi, 2020), the ARDL model (Afees, *et al.* 2020), the VAR model (Sangyup and Junhyeok 2022), etc.

Based on *markov-switching GRG copula model*, Kuang (2017) explored tail quantile dependences between the inflation rate and the real estate investment trust (REIT) return. Finding say that the positive and negative co-movements coexist. In the negative co-movement state, the REIT cannot hedge inflation risk, while in the positive co-movement state, the REIT has a partially hedging ability.

Later, Salisu, *et al.* (2019) examined the inflation hedging potential of the two most valuable precious metals namely gold and palladium. They employed both *time series and panel data techniques* for country-specific and group analyses. They concluded that both gold and palladium provide hedge against inflation in OECD countries notwithstanding the varying results across the individual countries. While the inflation-hedging potential of gold has been sustained, it only improves for palladium after the Global Financial Crisis. Their conclusions are sensitive to data frequency.

Recently, Akinsomi (2020) used a *comparative analysis of the year-to-date (YTD)* returns of global returns index and REITs sectors in the United States. They reveal that most sector REITs during the pandemic have lost considerable value based on YTD returns as at May 2020. Flight to quality is expected during this uncertain period to REITs such as data REITs, grocery-anchored REITs and storage REITs. These REITs are not as adversely affected by COVID-19 in comparison to other REITs.

Afees, *et al.* (2020) analyzed also asset-inflation hedging nexus for the US with the aim of determining inflation hedging characteristics of selected assets; stocks, gold, and real estates using the *bivariate and multivariate modelling* frameworks that taking into account of the asymmetry, the time-variation and the structural breaks. Founding say that inflation hedging tendencies of assets are heterogeneous across the considered assets. The real estates and stocks are proved to be good hedges against inflation, while gold investment defied Fisher's hypothesis. However, even the results are robust to alternative data frequencies, they are sensitive to the decomposition of data for pre- and post-GFC periods, indicating that asset-inflation hedging relationship for the US is time-varying.

Using a *Vector Autoregression* (VAR) model, Sangyup and Junhyeok (2022) provided systematic evidence on the relationship between inflation, uncertainty, and Bitcoin. Bitcoin appreciates against inflation (or inflation expectation) shocks, confirming its inflation-hedging property claimed by investors. The main findings hold with or without the COVID-19 pandemic episode.

To the best of our knowledge, only one study in the above literature has consider the *ARDL model* (Afees, *et al.* 2020) and only one which consider both univariate time serie and panel data analysis (Salisu, *et al.* 2019).

In this paper, we'll conduct an analyses on three developed countries including the United Kingdom, Canada and Switzerland for a period spanning from 1999 to 2018 covering the 2008 GFC period using univariate and panel ARDL models. We which to see if the asset-inflation hedging relationship for the considered sample is time-varying or not (say if results are sensitive to the decomposition of data for pre- during and post- GFC periods).

2. Models and Results

According to the GFH, in an efficient market, investors should be fully compensated for the increased price levels even if inflation decreases the value of money. Associated with perfectly competitive and informationally efficient capital markets in which investors are rational, the GFH postulates that stock prices should move *one-for-one* with goods prices to compensate investors for prices growth (inflation). This implies that stock returns should serve as a *hedge* against inflation, that is, *real* stock returns and inflation are independent. Consequently, we should observe a positive and *one-to-one* relationship between *nominal* stock returns and inflation rates.

GFH verification can be implemented in different specifications (static or dynamic). Dynamic specifications are considered and applied in the following sub-sections. Two type of data will be used: Time series and panel data.

2.1 The Panel ARDL Models

The framework and then methodology adopted in this paper are in two-fold; models with *heterogeneous slopes* and models with *homogeneous slopes.*

Case of heterogeneous Slopes

We consider a panel ARDL(p, q) framework formulating the Fisher dynamic equation as follows:

$$R_{it} = \alpha_i + \sum_{j=1}^{p} \delta_{ij} R_{i,t-j} + \sum_{j=0}^{q} \beta_{ij}\, INF_{i,t-j} + \varepsilon_{it} \quad (1)$$

We can reparametrize this model as the following ECM representation

$$\Delta R_{it} = \alpha_i + \varphi_i (R_{it-1} - \beta_i INF_{i,t-1}) + \sum_{j=1}^{p-1} \delta^*_{ij} \Delta R_{i,t-j} + \sum_{j=0}^{q-1} \beta^*_{ij}\, \Delta INF_{i,t-j} + \varepsilon_{it} \quad (2)$$

for i = 1, 2, N = 3 and t from 1999M01 to 2018M04 (TN = 696), where $\varphi_i = -\,(1 - \sum_{j=1}^{p} \delta_{ij})$, are the *speed of adjustment* to the long-run equilibrium, which is expected to be negative, $\gamma_i = \sum_{j=0}^{q} \beta_{ij}$, δ^*_{ij} and β^*_{ij} are the short-run coefficients (all are real parameters); $\delta^*_{ij} = -\sum_{m=j+1}^{p} \delta_{im}$, $j = 1, \dots, p-1$, $\beta^*_{ij} = -\sum_{m=j+1}^{q} \beta_{im}$, $j = 1, \dots, q-1$, the long-run coefficients $\beta_i = \frac{\gamma_i}{\varphi_i}$, and error-correction term $ECT_{it} = R_{it} - \beta_i INF_{i,t}$, ε_{it} is the error term which is independently distributed across i and t, while the term β_i are the *heterogeneous slopes.*

If $\varphi_i < 0$, then there is error correction, which implies that R_{it} and $INF_{i,t}$ are cointegrated, whereas if $\varphi_i = 0$, the error correction will be absent and there is no cointegration. This suggests that the null hypothesis of no cointegration for cross-sectional unit *i* can be implemented as a test of H_0: $\varphi_i = 0$ vs H_1: $\varphi_i < 0$.

Alternative methods of estimation to Fixed Effect (FE) and Random Effects (RE) estimators are suggested in (Pesaran, *et al.* 1999); henceforth PSS. The mean group (MG) estimator for MG model and the pooled mean group (PMG) estimator for PMG model.

Case of Homogeneous Slopes

We consider the model with elements β_i are *common* across countries:

$$\Delta R_{it} = \alpha_i + \varphi_i\left(R_{it} - \beta\, INF_{i,t}\right) + \sum_{j=1}^{p} \delta_{ij}^{*} \Delta R_{i,t-j} + \sum_{j=0}^{q} \beta_{ij}^{*}\, \Delta INF_{i,t-j} + \varepsilon_{it} \quad (3)$$

Pesaran, *et al.* (1999) refer to equation (5) as PMG model. The main characteristic of PMG model is that it allows short run coefficients (δ_{ij}^{*} and β_{ij}^{*}), the intercept (α_i), the error correction term (φ_i), and error variances (σ_i^2) to be *heterogeneous by country*.

PSS developed the PMG estimator, where the long-run parameters β_i are constrained to be the same (Belke and Dreger 2013).

To specify a model (either (2) or (3)), we use the (Hausman, 1978) type test, and we determine the most appropriate estimator either Pooled Mean Group (PMG) or Mean Group (MG) [or Dynamic Fixed Effect (DFE)].[3]

As diagnostic for the results, we perform several *causality tests*.[4] For the validity of considered models, there are several requirements. First, the coefficient on the error-correction term have to be negative and significant. Second, errors have to be White Noise (WN).

For the GFH to be hold, the slope restriction β = 1 should not be rejected (see, for example (Rushdi, *et al.* 2012 and Nassar and Bhatti, 2018)). Since the $\hat{\beta}$, estimate of the slope coefficient of the generalized Fisher relation may be less than 1 ($\beta < 1$) (Mundell, 1963 and Tobin 1965) or greater than 1 ($\beta \geq 1$) (Darby 1975), then common stocks will provide a partial or superior hedge against inflation. However, negative values of β suggest that the asset may act as a 'perverse hedge' against inflation.

Data Description

This paper uses a dataset for three (N = 3) countries, including Suisse, UK, and Canada over the period from 1999M01 to 2018M04 (T = 232). The stock price SP data is obtained from the investing.com while the consumer price CPI series is obtained from OCDE. We use a large sample that includes both the pre- and post-2008-2010 periods of the Global Financial Crisis (GFC).

Data will be explored separately for time series (Panel A) and for Panel context (Panel B). In the first steps, descriptive statistics (average value, Median, Maximum, Minimum, standard deviation, Skewness, Kurtosis, Jarque & Bera (J-B) statistic and its p-value) will be presented. Results for both cases are given at Table 1.

[3] We test the null hypothesis of homogeneity through a Hausman-type test. Under the null hypothesis of long-run slope homogeneity, both the PMG and MG estimators are consistent; however, only the PMG estimator is efficient. In other words, the Hausman test is used to compare the PMG and MG estimators. However, if the parameters are in fact homogeneous, the PMG estimates are more efficient. If we cannot reject the null hypothesis of homogeneity, data supports the PMG estimator to analyze the model.

[4] Causality can be then determined using the significance of (i) Error correction term (ECT) for joint causality (H_0: $\varphi_i = 0$), (ii) Long run coefficients for long run causality (H_0: $\beta = 0$), (iii) Short run coefficients for short run causality (H_0: $\beta_{ij} = 0$), and (iv) the simultaneous significance of ECT and long- and short-run coefficients for strong causality (H_0: $\beta_{ij} = \beta = \varphi_i = 0$).

Table 1. Descriptive Statistics

Panel A: Time series Data for full period and by country.

	Suisse		UK		Canada	
	R	INF	R	INF	R	INF
Mean	**0.004488**	**-0.001005**	0.000853	**-0.000826**	0.003578	**0.000604**
Std. Dev.	0.048059	0.013286	0.035912	0.015755	0.037974	0.016252
Skewness	-0.843526	0.194549	-1.286033	-0.898038	-1.499048	-0.357624
Kurtosis	6.141381	4.192009	7.489664	5.524600	11.05822	6.041953
J-B	122.3763	15.13321	257.6864	92.39512	711.5136	93.98870
Probability	0.000000	0.000517	0.000000	0.000000	0.000000	0.000000

Panel B: Panel Data for full period.

	LSP	LCPI	R	INF
Mean	4.574954	4.618735	0.002973	-0.000409
Std. Dev.	0.282254	0.143063	0.040963	0.015148
Skewness	-0.452100	-0.305948	-1.112308	-0.421325
Kurtosis	2.879325	2.795291	8.006419	5.622134
J-B	24.13208	12.07335	866.6297	219.0355
Probability	0.000006	0.002389	0.000000	0.000000

Empirical Results and Discussions

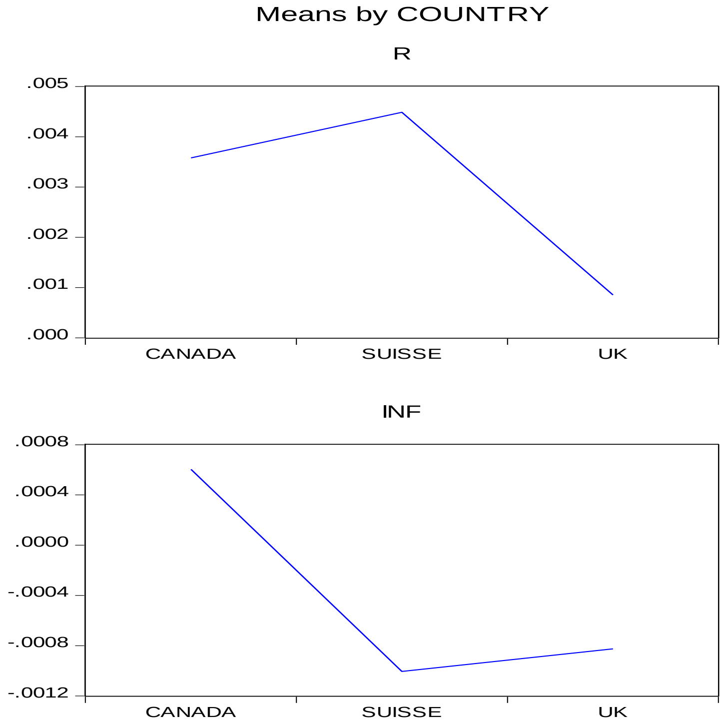
Table 1 (panel A), for each time series return (R = Δlog(SP) and inflation (INF = Δlog(CPI)), presents the descriptive statistics in average, we conclude that (see also Figure 1):

$$\bar{R}_{\text{UK}} < \bar{R}_{\text{Canada}} < \bar{R}_{\text{Suisse}},$$

$$\overline{INF}_{\text{Suisse}} < \overline{INF}_{\text{UK}} < \overline{INF}_{\text{Canada}},$$

where $\bar{R}$ and $\overline{INF}$ denote respectively the mean for R and for INF.

Figure 1. Average point estimate of R and INF by country.



For Panel Data (see Table 1 (panel B)), the same descriptive statistics are presented for prices in log (stock price LSP and consumer price index LCPI) and in first differences (return R and inflation rate INF). All skewness parameters are negative. Coefficient of kurtosis are greater than 3 for both variables, R (almost equal to 8) and INF (almost equal to 5). J-B test statistics reject the normality assumption. All considered variables have not Gaussian distribution (we reject null hypothesis that the sample is Normally distributed at 5% significance level).

The second step in our analysis is to test whether the variables in levels [stock price in log (LSP) and Consumer price index in log (LCPI)] are stationary or not. To this end, we employ a battery of unit-root tests. As shown in Table A1 (see Annex) all considered unit root tests (LLC, Breitung, IPS, ADF-F, and PP-F) indicate that stock price in log (LSP) and Consumer price index in log (LCPI) are non-stationary. However, opposite results were obtained for variables in 1st differences; the stock return (R) and the inflation rate (INF). So that, variables in level are integrated of order one, i.e. I(1) or Difference-Stationary.

Our study will relate the monthly return on the three stock market to the monthly rate of inflation for the three countries over the period from 1999M01 to 2018M04 (TN = 693). We consider then the dynamic equation (2) and we report results of the PMG, MG, and DF methodology within panel ARDL framework.[5] Table 2 shows the long run effects of inflation rate on stock return in four scenarios: for Full data set [1999M01—2018M04], for Pre the GFC period [1999M01—2007M12], during the GFC period [2008M01—2009M12], and for Post the GFC period [2010M01—2018M04].

When estimating panel ARDL equation (2), we use the maximum likelihood approach.[6] We did not report the short-run coefficients because only long-run parameters have importance in the generalized Fisher hypothesis. The long-run results obtained from the PMG and MG and Dynamic Fixed Effects (DFE) estimator are given at Table 2.[7]

As shown in Table 2, the Hausman test provides evidence favorable to the PMG (DFE) estimator for Pre (Post) GFC period. During crisis period as well as for full period of study, Hausman test provides evidence favorable to the DFE estimator. Then, we can say that it is the GFC period result which drives the results for full sample case.

According to the results of PMG estimator and at Pre GFC period (Table 2), the inflation rate is not significant even at the 10% significance level, and we cannot reject the null hypothesis of $\beta = 0$. Then, results do not support long-run causality at Pre GFC period. But, short run causality test results indicate significant causality only for UK (at 5% revel) and Suisse (at 10% level) stock market from inflation rate to stock return (we reject the null hypothesis of $\beta_{ij} = 0$). These results are not reported at Table 2 (but are available upon request). So, no strong causality can be deduced.

For the full period of study, the coefficient of inflation rate β is significant but is lower than unity ($\hat{\beta} = 0.39301$), while for the Post (Pre) GFC period, the coefficient β is not significant and is very lower than unity [$\hat{\beta} = -0.02227$ (0.0526)]. Thus, the results for full period do support a partial Fisher effect (and then long run causality from inflation to stock return is evident), while the Post GFC relation can be connoted by a worse hedge situation since $\hat{\beta} < 0$. This negative relationship post GFC can be due to the Money Illusion, Tax Effect, Proxy Effect, and/or Reverse Causality Hypotheses, and it may have important economic and policy implications. For instance, it would mean that investors would be better off in reducing their stock market investments in times of high inflation rates (Antonakakis, *et al.* 2017). However, during GFC period [2008M01-2009M12], a complete (or strong) Fisher effect does hold ($\hat{\beta} = 1.1683$), because the null hypothesis of $\beta = 1$ is not rejected at conventional significance levels (5%).

Additionally, the negative and significant error correction term estimator ($\hat{\varphi}$) indicates that there is a joint causality relationship between stock return (R) and Inflation rate in all considered cases. Precisely, $\hat{\varphi}$ indicates a causality from inflation rate to stock return that implying that inflation rate drives stock Return toward long-run stable equilibrium. This *unidirectional causality* from inflation to stock returns hints an *inefficiency* of these stock market which suggests that information on past

[5] We used Akaike Information Criteria (AIC) to select lag length for each individual country regression.

[6] This is done by STATA 15.

[7] DFE estimates the dynamic fixed effects model where all parameters, except intercepts, are constrained to be equal across panels.

values of inflation could *provide opportunities* for abnormal *gains* from the return R particularly in GFC period.

Table 2. Panel ARDL model results; PMG, MG, and DFE estimates from equation (3), (2) and FE model respectively

		PMG	MG	DFE	Hausman 1	Hausman 2
Full	$\hat{\beta}$	.272045	.3980174	.39302**	0.10	8.95
		(.16612)	(.42983)	(.170969)	(0.7507)	(0.0028)
	$\hat{\varphi}$	-.7327**	-.76304**	-.75712**	PMG	**DFE**
		(.05294)	(.04677)	(.037441)		
t-Statistic	(H_0: β = 1)	19.20***	1.96	12.60***		
Hedge ?				Yes		
Pre GFC	$\hat{\beta}$	.0526565	-.0057294	.1118374	0.33	0.15
		(.285425)	(.302975)	(.323041)	(0.5656)	(0.6957)
	$\hat{\varphi}$	-.73093**	-.74223**	-.71922**	PMG	**PMG**
		(.063249)	(.065605)	(.054234)		
t-Statistic	(H_0: β = 1)	11.02***	11.02***	7.56***		
Hedge ?		No				
Crisis period	$\hat{\beta}$	2.0566**	1.337227	**1.1683****	0.73	14.84
		(.489742)	(.973091)	(.541321)	(0.3923)	(0.0001)
	$\hat{\varphi}$	-.7621**	-.90580**	**-.77628****	PMG	**DFE**
		(.12959)	(.147341)	(.13066)		
t-Statistic	(H_0: β = 1)	4.65**	0.12	**0.10**		
Hedge ?				**Yes**		
Post GFC	$\hat{\beta}$	-.099922	.0403785	-.0222702	0.15	12.91
		(.1765201)	(.401032)	(.177838)	(0.6968)	(0.0000)
	$\hat{\varphi}$	-.85803**	-.9102**	-.90405**	PMG	**DFE**
		(.082799)	(.067086)	(.058561)		
t-Statistic	(H_0: β = 1)	38.83***	5.73**	33.04***		
Hedge ?				No		

Notes: (1) PMG estimates the pooled mean-group model where the long-run effects, β, are constrained to be equal across all panels. The short-run coefficients are allowed to differ across panels. MG estimates the mean-group model where the coefficients of the model are calculated from the unweighted average of the unconstrained, fully heterogeneous model. DFE estimates the dynamic fixed effects model where all parameters, except intercepts, are constrained to be equal across panels.

(2) The maximum number of lags for each variable is set at 1 and 0, and optimal lag lengths are selected by the AIC. Numbers in parenthesis are the standard errors. Probability value is reported for the Hausman test in parenthesis. Conclusion is given under p-value. ***, ** indicates 1% and 5% level of significance. Hausman 1 is to compare MG and PMG estimator. Hausman 2 is used to compare PMG and DF estimators. φ ≡ Speed of adjustment. (3) Three period are considered: Pre GFC from t = 1999M01 to 2007M12 (TN = 324), crisis period from 2008M01 to 2009M012 (TN = 72), and Post GFC period from 2010M01 to 2018M04 (TN = 300). Null hypothesis of no cointegration for cross-sectional unit *i* can be implemented as a test of H_0: $\varphi_i = \varphi = 0$ vs H_1: $\varphi_i < 0$. Source: Authors' calculations. Detailed results of the panel ARDL estimation are available upon request from the authors.

In conclusion, from the panel data analysis, evidence in favor of stock returns acting as an inflation hedge is partially existent for the full period, completely or strongly existent during the GFC period, and not existent pre and post the GFC. The results confirm then that the relationship between the two variables (stock return and inflation) has evolved heterogeneously overtime (Pre, during, and Post Global financial crisis (GFC)).

Conclusion

As mentioned earlier, there is no general consensus among empirical research on the validation of GFH (Antonakakis, *et al.* 2017). In addition, all the studies in the literature are based on time series data, and few papers, to the best of our knowledge, use panel data. This paper intends to bridge this gap and make some contributions to the empirical literature on the Generalized Fisher Hypothesis (GFH) and the inflation-hedging ability of countries commons stocks market. To this end, we consider a panel data from three democratic countries, including Canada, UK, and Suisse from 1999M01 to 2018M04 covering the 2008 GFC period.

Besides empirical studies based on time series data (details are not reported, only a sum up is given in Table 3 hereafter),[8] we demonstrate that the results can be more informative with panel data. As well, it is of great importance to see if the long run relationship between stock return and inflation can evolve heterogeneously overtime.

Findings confirm that GFH tests give different conclusions over considered sub-periods with either univariate time series or panel data. Results are sensitive to the decomposition of data for pre- and post-GFC periods, indicating that asset-inflation hedging relationship for the considered sample is time-varying. Table 3 gives a sum up of all the previous results. Looking at Table 3, panel data reveal unambiguous unstable relationship between return and inflation that is driven by Suisse stock market case.

Table 3: Results for inflation hedging in the full period, pre- during and post- GFC.

Data	Suisse	UK	Canada	Panel
Full period	Yes	No	Yes	Yes
Pre GFC	No	**No**	**Yes**	No
GFC	**Yes**	**No**	**Yes**	**Yes**
Post GFC	No	**No**	**Yes**	No

Note: This is a sum up of Table 2. Details of univariate ARDL results are not reported here but are available upon request. The results of first, second, and third column are the sum up of univariate time series models.

Based on the panel data analysis, results demonstrated that hedging property against inflation is true only during GFC crisis. And then, the major implication from eventual ability of financial assets to hedging against inflation is to encouraging investment and saving decisions in the three considered economics during crisis period as the GFC case (here deflation period). Indeed, since Suisse and Canadian stock return has a positive relationship with inflation, then including the UK in a portfolio allows investors to limit losses caused by inflation in UK stock market alone. Then, being simultaneously on the three considered market, investor will have some abnormal gain only during crisis period (here period of deflation).

[8] Based on univariate time series data, we conclude that Canadian (UK) stock return is (not) a hedge against inflation for the three sub-periods, while Suisse market return is a hedge against inflation only during GFC crisis. During crisis both Suisse and Canadian stock returns are superior hedge against inflation. Post crisis, the Canadian stock market is unique to be full hedge against inflation (this result is in accordance with (Richard and Ran, 2021)). No significant relationship is found in the UK context during crisis period (period of deflation). In addition, post and Pre crisis, UK stock market is found to be worse hedge against inflation.

ANNEX

Table A 1: Panel unit root tests at level and first difference (full period).

	LSP		LCPI	
Method	Statistic	Prob.**	Statistic	Prob.**
Null: Unit root (assumes common unit root process)				
LLC t*	0.39027	0.6518	0.70864	0.7607
Breitung t-stat	-1.83138	0.0335	-0.69113	0.2447
Null: Unit root (assumes individual unit root process)				
IPS W-stat	-0.49646	0.3098	0.45263	0.6746
ADF - Fisher χ^2	6.48521	0.3711	3.92995	0.6862
PP - Fisher χ^2	6.75911	0.3437	3.94596	0.6840
Conclusion		**I(1)**		**I(1)**
	R		INF	
Method	Statistic	Prob.**	Statistic	Prob.**
Null: Unit root (assumes common unit root process)				
LLC t*	-28.4817	0.0000	-30.8294	0.0000
Breitung t-stat	-14.3854	0.0000	-13.0169	0.0000
Null: Unit root (assumes individual unit root process)				
IPS W-stat	-21.5516	0.0000	-23.2148	0.0000
ADF - Fisher χ^2	263.928	0.0000	287.948	0.0000
PP - Fisher χ^2	267.942	0.0000	287.721	0.0000
Conclusion		**Stationary**		**Stationary**

Note: LLC ≡ Levin, Lin & Chu, IPS ≡ Im, Pesaran and Shin.

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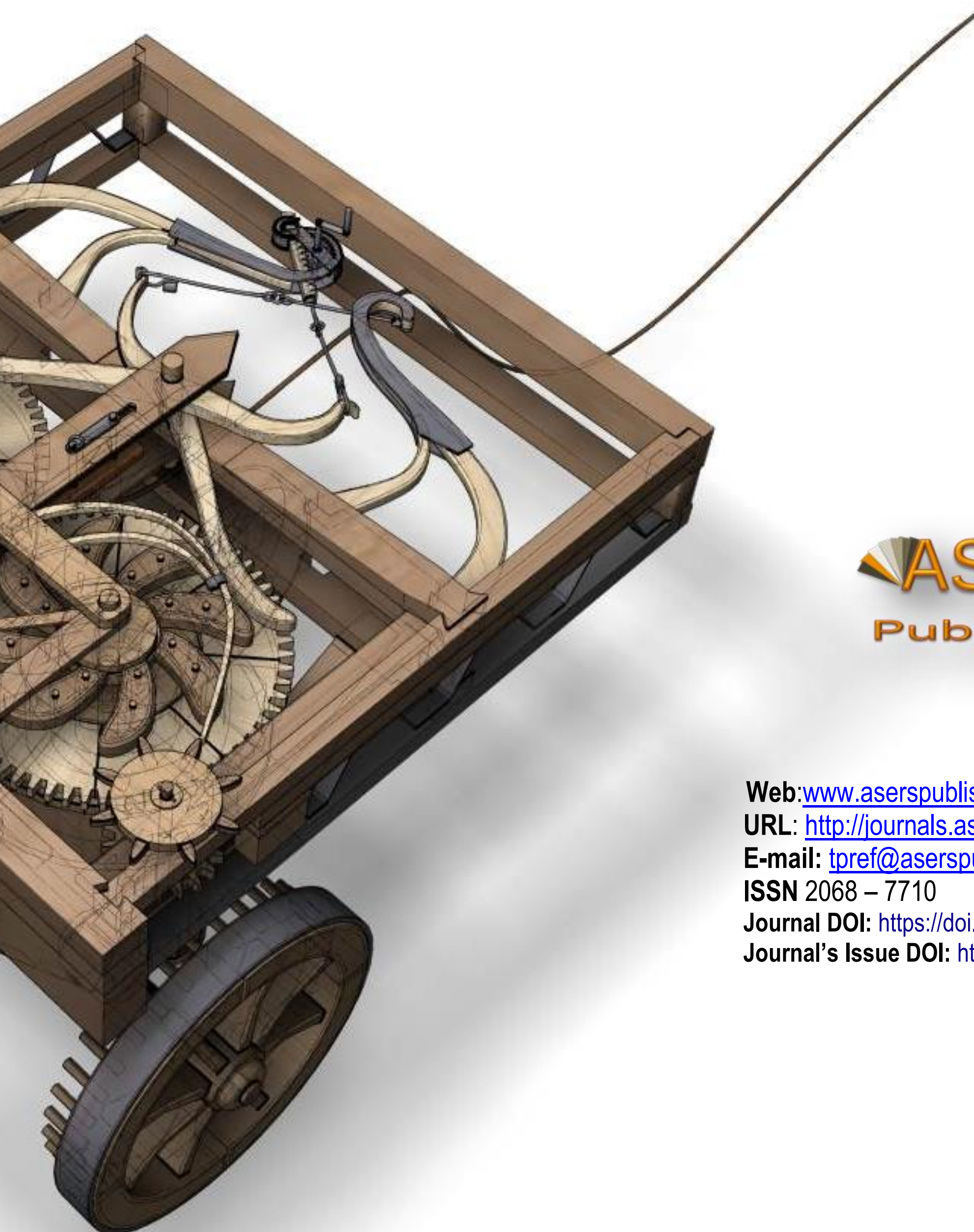

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