# The relationship between inflation and small-cap premiums and evaluating small-cap stocks as a hedge to inflation for G7 countries

Alison Cabana-Wong

A Thesis

in

The Department

of Finance

Presented in Partial Fulfillment of the Requirements

for the Degree of Master of Science (Finance) at

Concordia University

Montreal, Quebec, Canada

November 2022

**CONCORDIA UNIVERSITY**

**School of Graduate Studies**

This is to certify that the thesis prepared

By: Alison Cabana-Wong

Entitled: The relationship between inflation and small-cap premiums and evaluating small-cap stocks as a hedge to inflation for G7 countries

and submitted in partial fulfillment of the requirements for the degree of

**Master of Science (Finance)**

complies with the regulations of the University and meets the accepted standards with respect to originality and quality.

Signed by the final examining committee:

________________________________________ Chair
Dr. Rahul Ravi

________________________________________ Examiner
Dr. Alan Hochstein

________________________________________ Thesis Supervisor
Dr. Lorne N. Switzer

Approved by __________________________________________________
Dr. Nilanjan Basu

________ 2022 __________________________________________________

Dr. Anne-Marie Croteau

**Abstract**

The relationship between inflation and small-cap premiums and evaluating small-cap stocks as a hedge to inflation for G7 countries

Alison Cabana-Wong

Inflation erodes personal wealth, and portfolio managers must seek inflation hedging strategies to mitigate this risk. Small-caps stocks have been shown to react more negatively to macroeconomic shocks than large caps and a better understanding of this effect on small-cap returns can have important implications for portfolio management decisions. This paper studies the effect of inflation shocks on G7 small-cap premiums as well as the regime-dependent relationship of premiums with inflation and default risk. This paper also studies the long-run hedging properties of small-cap and large-cap stocks against inflation risk in G7 countries from 1994 to 2022 by estimating their Fisher Elasticities through univariate and panel analysis. Markov Switching regression shows that the variation in the small-cap premium can be explained by inflation and default risk factor regimes under weak and strong economic regimes. Cross-sectionally dependent panel cointegration analysis shows that G7 country large-cap and small-cap stocks are effective hedges against inflation in the long run in the post-GFC period. These results have important implications for building effective hedging strategies based on the market capitalization of publicly traded companies in the G7.

**Acknowledgements**

First and foremost, I would like to thank my research supervisor Dr. Lorne N. Switzer. It was my great privilege to have had the opportunity to work with him, and I immensely grateful for his expertise and guidance throughout the process of writing my thesis.

I have many people to thank for their encouragement and support; I am eternally grateful to my friends and family, my professors at the John Molson School of Business, and my cohort of the MSc Finance program. Writing this paper was the culmination of many years of long hours, sacrifice, and dedication, I owe so much of my accomplishment in completing the MSc to my partner Justinian La Rosa who supported me through every part of it.

## Table of Contents


| | |
|---|---|
| List of Tables | vi |
| List of Figures | ix |
| Body of Thesis | 1 |
| References | 82 |

## List of Tables

Table 1 – Sources of the stock price indices used in the analysis, the date range for each sample, and the source of data, for small cap and large cap indices of each G7 country

Table 2 – Descriptive statistics for monthly returns on small cap indices, by country, 1994-2022

Table 3- 12 month change in CPI by country, full date ranges by country, monthly data.

Table 4 – Threshold 3rd quintile of 12-month percentage change in CPI 12 by country

Table 5 – Markov Switching Model Regression Output for US 2003M06-2022M08

Table 6 – Constant Markov Transition probabilities and expected duration for the US from 2003M06-2022M08

Table 7 – Markov Switching Mode l Regression Output for Canada 2003M06-2022M08

Table 8 – Constant Markov Transition probabilities and expected duration for Canada from 2003M06-2022M08

Table 9 – Markov Switching Model Regression Output for Germany 2003M06-2022M08

Table 10 – Constant Markov Transition probabilities and expected duration for Germany from 2003M06- 2022M08

Table 11 – Markov Switching Model Regression Output for Italy 2003M06-2022M08

Table 12 – Constant Markov Transition probabilities and expected duration for Italy from 2003M06- 2022M08

Table 13 – Markov Switching Model Regression Output for Japan 2003M06-2022M08

Table 14 – Constant Markov transition probabilities and expected duration for Japan from 2003M06-2022M08

Table 15 – Markov Switching Model Regression Output for the UK 2003M06-2022M08

Table 16 – Constant Markov Transition probabilities and expected duration for the UK from 2003M06- 2022M08

Table 17 – Markov Switching Model Regression Output for the France 2003M06-2022M08

Table 18 – Constant Markov Transition probabilities and expected duration for France from 2003M06-2022M08

Table 19 – Results from the Granger Causality tests performed for variable pairs of 12-month change in CPI and small cap premiums for each country in the G7 from 1994M01-2022M08
Table 20 – Correlation table for 12 Month Change in CPI for G7 Countries 1994M01-2022M08

Table 21 – OLS statistics where the first difference of monthly price indices are the dependent variable and CPI is the independent variable for each G7 country, 1994M01-2022M08

Table 22 – Univariate Unit Root Tests performed on price levels for each G7 country goods prices, large cap stock prices, and small cap stock prices

Table 23 - Univariate Unit Root Tests performed on first differences of each G7 country goods prices, large cap stock prices, and small cap stock prices

Table 24 – Summary of results of panel unit root tests from Levin Lin and Chu (2002) and Pesaran (2007) for all panel groups, "x" denotes not I(1) stationary

Table 25 – Levin Lin and Chu (2002) panel unit root test with AIC lags; full sample 1994m01-2022m08

Table 26 – Levin Lin and Chu (2002) panel unit root test with AIC lags; pre and post-GFC for all panel groups

Table 27 – Pesaran (2007) panel unit root test for e ach panel group, for price level series, and for first difference series

Table 28 - Pesaran (2007) panel unit root test for e ach pane l group, for price level series, and for first difference series for pre and post-GFC subsamples

Table 29 – Trace statistic results for Johansen Cointegration test for first order stationary series I(1)

Table 30 – Test statistics for the Pedroni Panel Cointegration test for all panel groups in date sample range 1994M01-2022M08

Table 31 – Test statistics for the Pedroni Panel Cointegration test for all panel groups in date sample range 1994m01-2009M06

Table 32 – Test statistics for the Pedroni Panel Cointegration test for all panel groups in date sample range 2009M07 - 2022M08

Table 33 – Kao Engle-Granger Panel Cointegration test results for subsample date ranges 1994M01-2022M08, 1994M01-2009M06, 2009M07-2022M08

Table 34 – Vector Error Correction Model Outputs for stock prices variables with cointegrating relationships with goods prices

Table 35 – Model estimates for Fully Modified Ordinary Least Squares (FMOLS) cointegration panel regression

## List of Figures

Figure 1 – Monthly prices for North America Small Cap and Large Cap Indices 1970M01 – 2022M08

Figure 2 – Monthly prices for Ex-North America G7 Small Cap Indices and Large Cap Indices 1994-2022M07

Figure 3 – Monthly prices for Untied States S&P 500 Index and Small Cap Index 1928M02-2022M07. Shaded Area = 12 Month Change in CPI > 3.2023%

Figure 4 – Monthly prices for Canada small cap and large cap stock indices 1970M01-2022M08, Shaded area = 12 month change in CPI >12.890173%

Figure 5 – Monthly Prices for United Kingdom Large Cap and Small Cap Stock Indices 1994M01-2022M08, Shaded Area = 12 month change in CPI > 4.4352%

Figure 6 – Monthly Prices for France Large Cap and Small Cap Indices 1994M01-2022M08 Shaded Area = 12 month change in CPI>3.3998%

Figure 7 – Monthly Prices for Japan Large Cap and Small Cap Stock Indices 1994M01-2022M08 Shaded Area = 12 month change in CPI>1.7297%

Figure 8 – Monthly Prices for Germany Large Cap and Small Cap Stock Indices 1994M - 2022M08 Shaded Area = 12 month change in CPI > 2.0333%

Figure 9 – Monthly Prices Italy Large Cap and Small Cap Stock Indices 1994M01-2022M08 Shaded Area = 12 Month Change in CPI>4.6661%

Figure 10 – Markov Switching Smoothed Regime Probabilities for Canada from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2.There are two state switches which occur during the period of the GFC, and the COVID-19 crash.

Figure 11 – Markov Switching Smoothed Regime Probabilities for Germany from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show that the two states for Germany's small-cap premium are highly unstable.

Figure 12 – Markov Switching Smoothed Regime Probabilities for Germany from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show that the two states for Germany's small-cap premium are highly unstable.

Figure 13 – Markov Switching Smoothed Regime Probabilities for Italy from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes during the period corresponding with the GFC and the COVID-19 crash.

Figure 14 – Markov Switching Smoothed Regime Probabilities for Japan from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes during the period corresponding with the GFC and the COVID-19 crash.

Figure 15 – Markov Switching Smoothed Regime Probabilities for the UK from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes are highly unstable throughout the sample period.

Figure 16 - Markov Switching Smoothed Regime Probabilities for France from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes are highly unstable throughout the sample period.

Figure 17 – Response of US Large Cap Index to US CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 18 – Response of US Small Cap Index to US CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 19 – Response of Canada Large Cap to Canada CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 20 – Response of Canada Small Cap Index to Canada CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 21 – Response of Japan Large Cap Index to Japan CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 22 – Response of Japan Small Cap Index to Japan CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 23 – Response of UK Large Cap Index to UK CPI Cholesky One S.D. (d.f. adjusted) innovation ± 2 analytic asymptotic S.E.s

Figure 24 – Response of UK Small Cap Index to UK CPI Cholesky One S.D. (d.f. adjusted) innovation ± 2 analytic asymptotic S.E.s

Figure 25 – Response of Germany Large Cap Index to Germany CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 26 – Response of Germany Small Cap Index to Germany CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 27 – Response of France Large Cap Index to France CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 28 – Response of France Large Cap Index to France CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 29 – Response of Italy Large Cap Index to Italy CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

Figure 30– Response of Italy Small Cap Index to Italy CPI Cholesky One S.D. (d.f. adjusted) Innovation ± 2 analytic asymptotic S.E.s

# Introduction

Inflation is back in headlines after the COVID-19 market crash in 2020 and the invasion of Ukraine by the Russian Federation disrupted global supply chains and sent inflation skyrocketing past central bank targets. Inflation is measured by the annual change in the Consumer Price Index (CPI), which is an index of the monthly price of a basket of goods for a certain country. Inflation erodes personal wealth, and it is thus imperative that a portfolio manager seek inflation-hedging investment strategies to optimize real returns. From 1928-2022 the 3rd Quintile 12-month change in CPI value in the United States was 3.23%, but in August 2022 it has rapidly increased beyond this level attaining 8.25%, with the other countries in the G7 experiencing similar rapid increases in CPI values (USBLS, 2022).

Studies that examine stocks as a hedge against inflation often root their analysis in the Fisher Hypothesis (1930) which asserts that the real interest rate is the difference between the nominal interest rate and inflation. It can be extrapolated that stocks represent claims on nominal assets whose prices are independent of goods prices and therefore should have a unit positive relationship with nominal rates. For stocks to be a perfect hedge against inflation and for the Fisher Equilibrium to be true, the Fisher Elasticity between stocks and goods prices must be equal to 1. Despite a large number of publications on this topic, the literature is still divided on whether stock prices can be a material hedge against inflation under the Fisher Hypothesis. Anari and Kolari (2001) postulated that returns-based data fails to capture the long-run information on the relationship between inflation and stock returns and proposed a methodology to study the long-run effects that are preserved in price levels. Studies following their paper have explored the long-run relationship between inflation and stock prices to attempt to deliver improvements on portfolio management strategies under inflation risk and still yield contradictory results. For instance, the stationarity or presence of a unit root in inflation for industrialized countries is still being debated. Thus, there is no definitive consensus on whether inflation shocks are persistent or transitory and can be hedged at all. Madadpour and Asgari (2019) summarized the literature on this subject spanning 1930 to 2018 and found the literature to be divided on whether there is a positive relationship between inflation and stock returns; however, they found that choice of statistical method has a significant impact on the power and outcome of the study. They find that studies employing linear regressions in their analysis have the least robust results. While these often suggest that there exists a negative relationship. In contrast, while nonlinear studies find positive or no relationship at all, there remains ground to cover on improving the robustness of statistical methods on this topic.

There has been much less coverage in the literature when it comes to small-cap stocks as an inflation hedge. Small-cap stocks have been known to react more negatively to economic shocks than large-caps, though have been shown to outperform large-caps during recoveries in Canada and the United States (US) (Switzer, 2010). Still, the question as to whether they may provide a long-run inflation hedge in a way that would be significantly distinct from large-caps across G7 countries has not been answered. Greater exposure to consumption risk, interest rate risk, and greater rates of investment set small-cap stocks apart from large-caps and offer a compelling reason to further investigate their exposure to risks at a macro-economic level.

In this paper, the small-cap stock prices and inflation prices, represented by the monthly Consumer Price Index (CPI) will be analyzed for the potential inflation hedging properties by stocks in G7 countries; the United States (US), Canada, Germany, Japan, Italy, France and the United Kingdom (UK). The proposed by Anari and Kolari (2001) will be applied, with the addition of higher power statistical techniques as seen in Gregoriou and Kontonikas (2010; 2013), and Omay et al. (2010a, 2010b,2015).

Additionally, this paper seeks to provide insight into whether there are differences in the influence of inflation on small-cap premiums under high and low inflation rate regimes. While Fama and Schwert

(1977) find sub-sectors of the economy react differently to economic shocks, they also find a gradual convergence of the behavior of the market in the long run as the market gradually reallocates supply and demand. Despite the wide acclaim of the Fama French Three Factor Model (1992) showing small-cap outperformance of large-caps, abnormal returns attributed to small-caps continues to be debated upon in the literature. It can be stated that small-cap stock premiums reflect a compensation for their increased exposure to some macro-economic risk as they tend to grow faster than large-cap companies during expansionary economic environment and tend to have greater negative returns during contractions due to higher bankruptcy risk (Switzer and Picard 2020). Shocks to inflation can come from monetary policy designed to impose a target inflationary rate or from world events affecting goods prices, thus the impact on individual companies may vary with their exposure to such macroeconomic factors. If small-cap stocks have different sector mixes to large-cap stocks, it follows that studying how they may react differently to macroeconomic factors depending on which inflation regime they are in is worth investigating.

The paper is organized as follows; first is the hypothesis being tested, followed by the review of literature up to this point, the empirical methodology, a description of the data used for analysis, the results of the analysis are presented, and finally conclude with discussion of the results and limitations of the study.

## Hypothesis Development

This paper contributes to the literature in two ways. First, is by testing the long-run hedging properties of small-cap equities for inflation risk from G7 countries from 1994-2022, to determine whether small-cap equities provide an inflation hedge materially different from investing in the larger market. Second, by determining whether there is evidence of regime-dependent macroeconomic factors for the small-cap premiums with respect to inflation amongst G7 countries.

The first part of the empirical analysis will focus on estimating the Fisher Elasticities for long run analysis of the relationship between small-cap and large-cap stocks with goods prices to test the following hypothesis:

*H1: There is a positive long-run relationship between small-cap stock prices and goods prices.*

The second part of the empirical analysis will focus on inflation regime-dependent differences in inflation risk on small-cap stocks to test the following hypothesis:

*H2: Inflation has a causal relationship with small-cap premiums where the coefficients of inflation in strong economic states and weak economic states are not equal.*

## Literature Review

*Stocks as a hedge against inflation*

The relationship between stock returns and the inflation rate has been shown to be negative in the literature at least in the short-run, in apparent disagreement with the Fisher hypothesis. Of note; Fama & Schwert (1977), Bodie (1976), Fama (1981). Famously, the Fama (1981) money demand hypothesis shows that there exists an inverse relationship between real output and inflation, and thus the relationship between stock returns and rates of inflation is said to be negative. However, as pointed out by Anari and Kolari (2001), the study of stock returns and inflation rates does not capture long-run information or persistence from shocks and structural breaks. Anari and Kolari (2001), who developed a methodology to study the long-run relationship by analyzing price levels rather than rates, found that from 1953-1998 the 6 industrialized countries studied all have greater-than-unit Fisher elasticities and thus conclude that stocks are a good long-run hedge to inflation. Subsequent papers attempt to reveal insights on long-run relationships between goods and stock prices stemming from this methodology. Madapour and Asgari (2019) provide a comprehensive review of the literature on the Fisher Hypothesis for long run relationships between stock and inflation and find that, while the literature is divided on the sign of the relationship, the majority reject a significant Fisher Effect, concluding that the outcomes are highly dependent on statistical method and context. We therefore focus our review on later papers and their statistical methods.

Boamah (2017) find a positive long-run Fisher relationship from 1994-2014 for G7 countries except for Italy, which did not meet stationarity conditions for cointegration analysis. Gregoriou and Kontonikas (2010) find that there is an overall positive relationship between goods prices and stock prices supporting the Fisher hypothesis using a panel cointegration test. Omay et al. (2015) study 52 countries from 1997-2007 and examine whether there is cross-sectional dependence for shocks in the panel data and find significant cross-sectional dependence and cointegrating relationships between stock and goods prices. They also found significant positive Fisher coefficients ranging from 0.68 to 1.27. Ciner (2015) finds a positive relationship between stocks and goods prices in the US from 1985-2013 for commodities and technology firms at the industry level, suggesting that individual stocks can be a good hedge when faced with transitory inflation shocks. Bampinas and Pangiotidis (2016) find similar results to Ciner (2015). Parikh et al. (2019) breaks down CPI into its components to find optimal hedges from listed companies in the US equity market and conclude that using individual stocks is the only way to reliably hedge but were unable to find any sector-wide hedges. Tiwari et al. (2020) use wavelet methods to determine causality between stock returns and inflation and find no significant short-run relationship but did find a significant long-run relationship for both nominal and real inflation rates and stock returns. Sami (2021) uses Pesaran's (2001) cointegration test for US and Canada from 1960-2019 and find a significantly positive long-run economic relationship, with above-unit elasticities.

Other studies have found non-positive relationships in the long-run. Liu and Serletis (2022) use GARCH and Bivariate VARMA to study relationship between equity returns and inflation across developed and emerging markets and find the effect of inflation is negative for Canada, Japan, USA, but positive for the rest of the G7 countries. Chan et al. (2003) find few cointegrating relationships between stocks and goods prices in 18 developed economies, however, the number of cointegration relationships increases after the 1987 market crash possibly signaling a contagion effect. Eldomiaty (2020) find stock prices from 1999-2016 are negatively associated with inflation rates but positively associated with real rates for the DJIA30, and the NASDAQ 100. Bampinas and Pangiotidis (2016) find that the overall hedging ability of stocks has decreased since the Global Financial Crisis in their study spanning 1993-2012, with fewer firms providing a positive hedge to inflation during economic crises. Hoesli et al. (2007) provide evidence that UK large-cap and small-cap stocks from 1977-2003 have a negative relationship with unexpected inflation (shocks), but that the US equities (including small-caps) had a positive relationship with long-run inflation over the same period.

First order non-stationarity of the variables is a requirement for testing for long-run cointegration. The widely used Augmented Dickey-Fowler (ADF) tests a null hypothesis of a unit root by approximating an autoregressive moving-average model for the time series (Dickey and Fuller, 1979; Said and Dickey, 1984). Publications following the Anari and Kolari (2001) paper have shown that the univariate ADF test can be improved upon in terms of statistical power in the context of testing the stationarity of macroeconomic factors. The ADF test has been shown to have a bias towards over accepting a unit root, especially in the presence of structural breaks (Perron 1997).

Chen (2022) shows that the ADF test which incorporates structural breaks generally provides greater power in testing for stationarity, though they note that when the structural break is small the ADF test is not likely to fail. The modified-ADF test, otherwise known as the ADF-GLS test, first proposed by Elliot, Rosenberg, and Stock (1992), has been shown to improve the power of unit root testing in near-stationary series such as inflation (Narayan and Narayan, 2008, Cheung and Lai, 1995). The ADF-GLS test involves estimating a GARCH (1,1) process and jointly estimating a unit root test equation by detrending the series via Generalized Least Squares (GLS) before performing a traditional ADF test, which has been shown to be superior in statistical power to an ADF test or a GARCH (1,1) specification alone (Cook 2008). Narayan and Narayan (2008) test stationarity of inflation across OECD countries using both the ADF and ADF-GLS test, and ADF test with breaks and found that the ADF-GLS test and ADF test with breaks had similar power which were both superior to the ADF test. Harvey et al. (2013) shows that a modified ADF test with multiple breaks determined by taking the infimum of the sequence of GLS detrended statistics has significant and superior unit root test results in for both power and bias. Tsong et al. (2012) employ a panel unit-root test which allows for cross-sectional dependency to increase the power of analysis in the presence of interdependence in the data and find 15 of 19 countries show mean-reversion for their sample period of 1999 – 2010. However, Cook (2009) performs an analysis using the ADF-GLS model on 13 OECD countries from 1957-1994, and find that, excluding Japan and Luxemburg, all country CPI series are I(0) stationary. This result is supported by the conclusions of Culver and Papell (1997) and Tsong (2012) who used panel unit root tests on inflation and stock data in developed countries (Cook, 2009). Narayan and Narayan (2008) find inflation rates for 17 OECD countries to be stationary by KPSS multivariate testing, in contrast with their results of non-stationarity with univariate tests. Gregoriou and Kontonikas (2006) show that countries with inflation targeting policies have stationary CPI with respect to their target, which indicates that they are successful in implementation, the speed of reversion increasing with the deviation from the target. Using a Markov-Switching unit root regression model, Chen (2010) finds non-linearity in the presence of unit roots across 11 OECD countries to be interest rate regime dependent, finding shocks to have greater impact in a high persistence interest rate regime. Omay et al. (2015) re-test stock prices and goods prices for a Fisher equilibrium relationship, extending the methodology from Gregoriou and Kontonikas (2010) . They take the panel analysis further by testing for cross-sectional dependence using the framework provided by Pesaran (2007). They find that ignoring cross-sectional dependence has a material impact on the results suggesting this method may produce more robust results (Omay 2015). Tsong et al. (2012) test OECD countries for unit roots and find more robust results when sub-dividing the panel into a group of stationary and non-stationary series. Liu and Serletis (2022) get mixed results for the effect of Inflation on stocks, observing that it has a negative effect on US stocks, no effect on UK stocks, positive effect on Canada, France, Italy and Japan from 1982-2020.

*Macroeconomic determinants of small-cap premiums under inflation regimes*

The literature around macroeconomic factors which influence the small-cap premium find that small-caps have distinctive behaviors to economic shocks, exhibiting more volatility than large-cap stocks, indicating that they may be exposed differently to macro risks. Switzer and Picard (2020) study the macroeconomic factors relating to small-cap premiums in the US and Canada using an LSTAR regime model and find that US default risk has a significant impact on both markets large and small-cap stocks, and a positive relationship with the spread in returns. They find Canadian small-cap stocks to be significantly influenced by US macroeconomic factors suggesting a spillover effect.

Small-cap stocks have been observed to react asymmetrically to poor economic conditions, reacting more negatively than large-caps. Kang et al. (2011) find that greater under-performance of small-cap and value stocks due to a greater exposure to consumption risk than large or growth stocks during economic contractions. They show that the macroeconomic factors that size and value type companies are most significantly exposed to are the term spread, default spread, short rate, and aggregate yield.

The literature suggests there may be regime-dependent differences in the relationship between inflation and other macroeconomic factors. Switzer and Picard (2020) found that the US dividend yield relates to the US small-cap premium during expansionary periods while inflation and term structure risk are only significant during contractions. Amenc et al. (2019) find size factors are significantly related to surprises in the effective spread and systemic volatility due to their link with returns and fast reaction to events. In the paper by Kremer et al. (2013) it is shown that amongst 124 countries in their study, the industrialized countries show growth-contraction threshold inflation rates at around 2.5% which indicates that central bank policies for rate targets at 2% is material as rates above this are shown to have a negative correlation to economic growth, while below has a positive correlation. Omay and Hasanov (2010b) use smooth transition models and show non-linear significant state-dependent effects of inflation on real interest rates, being more positive during periods of high inflation. Amenc et al. (2019) finds evidence that common macroeconomic factors amongst variables diminish diversification benefits as shocks can make them highly correlated when they may otherwise have low correlation. Liu and Serletis (2022) have found mixed results in studying the effect of inflation on stock prices in G7 and EM7 and suggest that this could be explained by the variance between countries in monetary policy, policy regimes, and economic structure. Zhang et al. (2009) provide evidence for the US and the UK that unexpected inflation is negatively correlated with small-cap premiums. They show that small-caps perform better when interest rates are low and when the term spread is high, likely due to their high reinvestment rate (ergo short-term investment and low dividend payouts). Connolly et al. (2022) show that the effect of inflation shocks on equity risk premiums in the US is state dependent, being significantly positive during weak economic times (economic contractions) occurring in 1997-2017. They find that this relationship is much weaker, and negative, during strong economic times, suggesting that stocks may have a stronger potential for hedging inflation during weak economic times with the presence of inflation shocks.

## Methodology

The analysis is carried out in two parts. First, we test macro-economic factors influencing small-cap and large-cap premiums for G7 countries under inflation regimes using the Markov-Switching (MS) Regression model. Second, we test stock prices as hedges to goods prices for both market index prices and for small-cap prices in G7 countries. Statistical analysis is conducted using statistical software packages in EViews version 12.0 and in SAS version 9.4.

*Inflation Under Economic -Regimes*

The variables used for each country are small-cap premiums, inflation, term risk, default risk, the risk-free rate, and stock market dividend yield. Small-cap premiums are the difference in returns for small-caps over large-caps (the market index), inflation is the 12 month change in goods prices, term risk is the difference between 10-year government bond yields and 3-month government bond yields, default risk is the difference between 10-year corporate bond yields and 3-month government bond yields or where this data is not available the 10 year swap spread index is used, the risk free rate is the overnight rate or the 30 day yield on government treasury bills, and the dividend yield from the stock market index represents the monthly dividend paid out by the large-cap index. Each series is in monthly periods.

The Markov-Switching Regression (Hamilton, 1989) is a two-state parametric model which provides probabilities of switching between two states governed by a Markovian state variable. The model assumes that regime changes occur instantaneously across variables through the threshold parameter. The MS model assumes that there are two linear regressions each associated with a regime expressed as an “m” state in the model. In other words, the conditional mean of our independent variable in state *m* is expressed as the linear specification of our independent variables in state *m*. Residuals are assumed to be normally distributed, and the sample variances are assumed to be regime dependent. The one-step ahead probability of regime change is calculated through an iterative maximization of the log-likelihood function. The dependent variable in the model is the small-cap premium, the regime-dependent independent variables are the 12-month change in CPI and default risk, corresponding to the results of a significant default risk factor for small-cap premiums from Switzer (2010) the non-switching regressor variables are term risk, dividend yield, and the risk-free rate as tested in Switzer and Picard (2020).

The autocorrelation is confirmed for the 12-month CPI for each country time series, serial correlation of the residuals and heteroskedasticity is verified.

*Long-run and Short-Run relationship between stock prices and goods prices*

We begin our analysis by taking the natural log of each time series of level prices to improve model goodness-of-fit. The dependent variable is the small-cap or large-cap price, while the regressing variables are the CPI prices (goods prices) for that respective country. We analyze the data with linear statistical methods commonly used in stock return analysis such as Pearson's correlation and Ordinary Least Squares (OLS) regression to evaluate the linear relationship. The methodology from Anari and Kolari (2001) provides the preliminary framework for our analysis, the methodology used in Gregoriou and Kontonikas (2010) and Omay (2015) is added to improve the robustness of the statistical analysis.

*Pearson's Correlation*

The Pearson's correlation is estimated for CPI, small-cap stock prices, and large-cap stock prices, across the countries under study. The potential diversification benefit for US inflation as a position for a US portfolio manager is considered.

*OLS Regression*

We run an ordinary least squares regression on the first difference of stock prices and of the CPI for each country to test the linear relationship as done in Gregoriou and Kontonikas (2010). First differences of stock prices as the independent variable and goods prices as the dependent for the full sample.

*Unit Root Tests*

For a long-run relationship to be studied, the stationarity of the series must be verified and all variables must be stationary of the same order otherwise there is a risk of a spurious regression result. If level prices are found to have a unit root, or in other words found to be non-stationary, but that their first differences have no unit root and are stationary, then we can conclude that this series is stationary of the first order I(1). Anari and Kolari (2001) employ a simple Augmented Dickey-Fuller test to establish whether the series are stationary and thus suitable for long-run cointegration analysis. However, the use of a linear univariate root test has been found to be insufficient in power for determining the stationarity of inflation over long time horizons. Each series is tested for first order stationarity I(1) through multiple methods and compared for statistical power, starting with the univariate tests; ADF test, modified ADF or ADF-GLS test, then the ADF test with breaks and finally panel root tests are conducted (first and second generation). The break point date is estimated as the date which minimizes the t-statistic in the sample. Lags are selected from the Akaike Information Criteria produced by the vector autoregressive model (VAR) for each time series.

Panel groups are formed; Full Sample all G7 countries, North American Countries (US, Canada), Ex-North American Countries (Germany, Japan, Italy, UK), Countries with first-order stationary goods prices I(1), countries with non-first-order stationary goods prices. Creating panel groups of stationary I(1) and nonstationary at I(1) in Tsong et al. (2012) provided new revelations about the time series which had non-I(1) stationary results from the univariate tests.

The first-generation panel unit root test that will be carried out is the Levin, Lin and Chu (2002) test which assumes independent time series sample effects with the presence of a common root, with the same AR(1) coefficient, and a time trend. The null hypothesises is the presence of a unit root and the estimate is the t-statistic. The second panel unit root test is the second generation Pesaran (2007) Cross-sectionally augmented IPS test (CIPS) which is takes the mean of individual ADF statistics from each panel member (Im et al. 2003), controls for cross-dependence with the inclusion of the CADF statistic, and assumes the panel is heterogeneous. The CADF statistic is the cross-sectional average of lagged level and first-differences of series in the panel. The null hypothesis of the CADF test is that there is no cross-sectional dependence, the null hypothesis of the CIPS tests is that all series in the panel contain a unit root. The Modified Akaike Information Criterion is used for determining the appropriate lag length according to each variable. We test each variable with the assumption of an intercept and each for no linear deterministic trend, and then again for the presence of a linear deterministic trend. Intuitively, panel models with cross-sectional dependence makes economic sense as country economies are likely to be interconnected through globalization.

*Cointegration tests*

The linear combination of two non-stationary variables is likely to also be non-stationary, however if variables are cointegrated then their linear combination results in a stochastic trend stationary series. Once the series have been identified as first-order stationary I(1), they can be tested for cointegration using Johansen's (1991) cointegration test with CPI as the first variable. A group of I(1) series is said to be cointegrated if there is at least one combination of the variables that is stationary that exists, the implication being that there is evidence of a long-run relationship between the variables. The series are tested under the assumption of a constant, but no deterministic linear trend, and then again with a constant and a linear deterministic trend. The null hypothesis is of no cointegrating vectors, or at most 1 cointegrating vectors as the total number of equations that are being tested is two (stock prices and goods prices). If the test statistic is greater than the critical value at 5% significance, then we reject the null hypothesis. Lag intervals are set according to the Akaike Information Criterion estimated from the VAR model. Panel Johansen cointegration tests are then conducted on the panel groups which are determined to be I(1) stationary. The Trace Test statistic will be used in analysis as the sample size is sufficiently large, and the power is greater than that of the max eigenvalue test.

*Vector Error Correction Model (VECM), and Fully Modified Ordinary Least Squares Regression (FMOLS)*

Anari and Kolari (2001) estimate the coefficients for the regression of cointegrating series using a VECM with the stock prices as dependent variables and goods prices as independent variables. The VECM is a system of vectors for two or more endogenous variables where a cointegrating relationship is assumed. Coefficients obtained from the VECM can be interpreted as the long-run Fisher Elasticities between goods prices and stock prices. If a group I(1) stationary series are comoving together, there should be an error correction term observed otherwise known as the cointegrating term. The speed of adjustment statistic provides a sign and magnitude of the first-order reversion back to the long-term trend, if it is negative, it signals those deviations from the trend are transitory.

A Fully Modified Ordinary Least Squares (FMOLS) regression is estimated for the panel groups as done in Gregoriou and Kontonikas (2010) and Omay (2015) who both find that the models estimated in their analysis to be of high goodness-of-fit. The FMOLS regression finds the optimal coefficients for cointegrating regressions while accounting for serial correlation. The FMOLS model controls for endogeneity by accounting for both the leads and the lags in the model. The dependent variable is the stock price panel, and the regressor variable is the goods prices panel. This coefficient represents the Fisher Elasticity in our model.

*Short-Run relationship*

The Impulse-Response functions (IRF) are estimated from each time series pair of goods prices and stock prices through a VAR model using Cholesky method variance decomposition specified with 12 lags, pertaining to the monthly periods in the data, and a horizon of 240 periods. Goods price series are the impulse function, and stock price series are the response function. The IRF will show the path of the time series variable after a unit shock of goods prices on stock prices.

## Data Sources

*Small-cap premiums and macroeconomic factor variables*

Data for the macroeconomic factors used were retrieved from Factset, CRSP, Bloomberg Terminal, and from the St. Louis Federal Reserve (St. Louis Fed.). ETFs are used to represent stock market prices and returns as they present a liquid, low cost, and investable vehicle for a portfolio manager to build an investment strategy. The broad market indices represent the larger cap market capitalization for comparison with the small-cap indices. The date range for analysis is limited by the availability of data for each G7 country and will span 2003M06-2022M08. Stock index sources found in Table 1.

For the US variables: Term risk; 10-Year US Government Bond yield at Constant Maturity Minus 3-Month US Treasury yield at constant maturity from the St. Louis Fed. Default risk; Market Yield on US Treasury Securities at 10-Year Constant Maturity minus Yield on US Corporate Paper at 10-year10-year constant maturity from the St. Louis Fed.. Dividend yield is the percent yield from the monthly dividend payment on the S&P 500 Index returns from Factset. Risk free rate; the 1-month US Treasury bill at constant maturity. Small-Cap premium; the difference between the $10^{th}$ percentile and $90^{th}$ percentile market cap returns from the Kenneth R. French database. The date range for the sample is 1927M11 to 2022M08.

For Canadian variables: Term risk; 10-year Canadian Government bond yield at constant Maturity minus 3-month Canadian government bond yield at constant maturity from the St. Louis Fed.. Default risk; the 10-Year Swap Spread Index in CAD from Bloomberg. Dividend yield; the percent yield from the monthly dividend payment on the S&P/TSX Composite Index returns from Bloomberg. Risk-free rate; the 1 month Canadian Government bill yield at constant maturity from the St. Louis Fed.. Small-cap premium; the difference between the BMO Nesbitt Burns Canada

Small-cap Index (Morningstar) combination with MSCI Canada Small-cap (Factset) returns and the S&P/TSX Composite Index returns. The date range for the sample is from 1927M11 to 2022M08.

For Germany variables: Term risk; the 10-year German Government bond yield at constant maturity minus the 3-month German Government bond yield at constant maturity from the St. Louis Fed.. Default risk; Germany 10-Year Swap Spread Index, local currency, from Bloomberg. Dividend yield; the percent yield from the monthly dividend payment on DAX Index returns from Bloomberg. Risk-free rate; German interbank overnight rate yield from the St. Louis Fed.. Small-cap premium; The difference between SDAX Index returns and DAX Index returns. The date range for the sample is 1994M01 -2022M08.

For Japan variables: Term risk; Japan 10 year government bond, constant maturity, minus the Japan 3 month government bond at constant maturity from the St. Louis Fed.. Default risk; no dataset found was sufficient to represent default risk and this variable will be omitted from analysis. Dividend yield; the percent yield from monthly dividend payment on the Nikkei 225 Index returns, taken from Bloomberg. Risk-free rate; Japan interbank rate from the St. Louis Fed.. Small-cap premium; the difference between Japan S&P Small-cap Index returns and the Nikkei 225 Index returns. The date range for the sample is 2002M04 – 2022M08.

For Italy variables: Term risk; The 10-year Italy Government Bond Yield at constant maturity, minus the 3-month Italy Government Bond Yield at constant maturity taken from the St. Louis Fed.. Default risk; The 10-year Swap Spread Index from Bloomberg. Dividend yield; the percent yield from the monthly dividend payment on the FTSE Italy Large-cap Index returns, from Bloomberg. Risk-free rate; Italy Interbank Rate from the St. Louis Fed.. Small-cap premium; The difference between the S&P Italy Small-cap Index returns and the FTSE Italy Large-cap Index returns. The date range for the sample is 2003M06-2022M08.

For United Kingdom (UK) variables: Term risk; the 10-year government bond from the Bank of England minus the 3-month government bond from the Bank of England from the St. Louis Fed.. Default risk; the 10-year Swap Spread Index, in local currency, from Bloomberg. Dividend yield; the percent yield from monthly dividend payment on the FTSE 100 Index returns from Bloomberg. Risk-free rate; the overnight interbank rate from the Bank of England, from the St. Louis Fed.. Small-cap premium; The difference between the S&P United Kingdom Small-cap Index returns and the FTSE 100 Index returns. The date range for the sample is 1999M12-2022M08.

For France variables: Term risk; The 10-year France government bond yield at constant maturity less the 3-month France government bond yield at constant maturity from the St. Louis Fed.. Default risk; France 10-year Swap Spread Index from Bloomberg. Dividend yield; the percent yield from monthly dividend payment on the FTSE CAC 40 index. Risk-free rate; the France overnight bank rate from the St. Louis Fed.. Small-cap premium; the difference between the MSCI France Small-cap Index returns and the FTSE CAC 40 Index returns. The date range for the sample is 1999M05-2022M08.

*Stock prices and goods prices variables*

Monthly market price data were retrieved from Morningstar, CRSP, and Factset. G7 Country data complete datasets spanned January 1994 to August 2022, while the datasets for the US spanned from February 1928 to August 2022, and the Canadian data from January 1970 to August 2022. For the purpose of analysis, the sample for analysis will be limited to January 1994 to August 2022. All data are in local currencies and economic data is not adjusted for seasonality. CPI data for the United States and Canada were retrieved from the St. Louis Fed.. CPI data for Japan, France, Germany, Italy and the United Kingdom were retrieved from Factset and the St. Louis Fed.

US stock price data were retrieved via the Kenneth R. French Data Library which compiles CRSP Index and U.S. market price data, as well as Factset. Canadian small-cap stock prices were retrieved from Morningstar and used the BMO Nesbitt Burns Canada Small-cap Index up until January 2015, and then the MSCI Canada Small-cap Index from January 2015 – August 2022. Small-cap Index data for Japan, Italy, and the United Kingdom were retrieved from Morningstar, for France and Germany these data were retrieved from Factset. Large-cap Index data for Italy were retrieved from Morningstar, while Canada, Germany, France, and Japan data were retrieved from Factset.

**Table 1 Sources of the stock price indices used in the analysis, the date range for each sample, and the source of the data, for small caps and large cap indices of each G7 country.**

*Panel 1 shows the description of each small cap index used in the analysis, Panel 2 shows the description of each large cap and market index used in analysis, including date range and source, for all G7 countries.*

*Panel 1*

| Country | Small cap Index | Date Range | Description |
|---|---|---|---|
| **United States** | CRSP US 1st Decile Index (Kenneth R. French Data Library) | 1928M02-2022M08 | US Stock Market Prices for the 1st decile of market capitalization of publicly traded companies. |
| **Canada** | BMO Nesbitt Burns Canada Small cap Index | 1970M01-2022M08 | The BMO Nesbitt Burns Canada Small cap Index consists of 400 companies with a market |
| **Japan** | Japan S&P Small cap Index (Morningstar) | 1994M01-2022M08 | Stocks included in the S&P Japan 500 but not in the S&P/TOPIX 150 or S&P Japan Mid Cap 100, contains 250 companies. |
| **France** | MSCI France Small cap Index (Factset) | 1994M01-2022M08 | The index contains 84 companies, represents the 14th percentile of market capitalizations of free-float adjusted market capitalization in the French equity market. |
| **Germany** | SDAX Index (Factset) | 1994M12-2022M08 | Index of 70 small-medium sized companies traded on the DAX, smaller in capitalization than stocks from the MDAX (EUR). |
| **Italy** | S&P Italy Small cap Index (Morningstar) | 1994M01-2022M08 | S&P Small cap Index for Italy, in local currency (EUR) downloaded from Morningstar, contains the bottom 15% of market capitalization companies that are publicly traded. |
| **United Kingdom** | S&P United Kingdom Small TR GBP (Morningstar) | 1994M01-2022M07 | This index contains the bottom 15% of market capitalization companies that are publicly traded. |

*Panel 2*

| Country | Large cap Index | Date Range | Description |
|---|---|---|---|
| **United States** | S&P 500 Index (Factset) | 1928M02-2022M08 | The S&P 500 includes 500 companies from the top performing industries of the U.S. market. |
| **Canada** | MSCI Canada Index (Factset) | 1970M01-2022M08 | Tracks the performance of large and mid-sized capitalization companies in Canada. |
| **Japan** | Nikkei 225 (Factset) | 1949M05-2022M08 | Price-weighted equity index of 225 stocks from the PMTSE (JPY) |
| **France** | FTSE CAC 40 (Factset) | 1994M01-2022M08 | The index tracks the top 100 market capitalisation and the most active stocks listed on Euronext Paris and contains 40 companies. |
| **Germany** | DAX Index (Factset) | 1994M01-2022M08 | The Dax Index measures the performance of the Prime Standard's 30 largest German companies in terms of order book volume and market capitalization and contains 40 companies. |
| **Italy** | FTSE Italy Large cap USD (Morningstar) | 1994M01-2022M08 | FTSE Italy Large cap Equity Index in local currency (EUR) downloaded from Factset |
| **United Kingdom** | FTSE 100 Index (Factset) | 1994M01-2022M08 | Market-cap weighted index of the 100 largest companies listed on the LSE. |

## Description of the Data

The descriptive statistics for monthly returns on small-cap indices for each country of the G7 is for the sample period from January 1994 to August 2022 (Table 2). The US & Canada are summarized as the North American group, while France, Germany, Italy, Japan and the United Kingdom are summarized as the Ex-North American Group. The North American group has greater mean and maximum returns, with 0.097% and 20.27% for the first, and 0.09% and 11.11% for the second group. Standard deviations are greater for the North American group with 5.4% versus 3.4%. The holding period return for North American small-cap stocks is 249.18% while for ex-North America it is 218.56%. The country with the greatest holding period return is from the United States with 339.08% and the country with the lowest is Japan with 111.00%. The country with the greatest recorded monthly return for small-cap stocks is Germany with 34.79% and the lowest minimum recorded is for Canada and France at -27.83% and -27.81% respectively. The country with the lowest standard deviation on small-cap returns is Japan with 4.94% and the highest is the United States and Italy with 6.31% and 6.25% respectively.

For large-cap stock returns, the North America group has a higher mean with 0.065% versus 0.0405% for ex-North America. The ex-North American group has a higher standard deviation of 3.59% versus 3.04%. The country with the greatest mean market returns is the United States with 0.0711% and the lowest mean return is for Japan with 0.002910%. The country greatest standard deviation in stock price returns is Italy with 8.36% and the lowest is the United Kingdom with 3.9%. The country with the greatest holding period return is the United States with 243.91% and the lowest is Italy with 104.97%.

**Table 2 Descriptive Statistics for monthly returns on small cap indices, by country, 1994-2022.**

Panel A shows the descriptive statistics for the small cap price indices for the G7 countries in the analysis while Panel B shows the descriptive statistics for the large cap price indices. Statistics are limited to the date range of 1994-2022. North American group represents Canada and the US, the Ex-North America group represents Germany, France, Italy, Japan, and the UK.

*Panel A*

| | N. America Small cap Index | Ex-N. America Small cap Index | United States Small cap Index | Canada Small cap Index | France Small cap Index | Germany Small cap Index | Italy Small cap Index | Japan Small cap Index | UK Small cap Index |
|---|---|---|---|---|---|---|---|---|---|
| **Mean** | 0.0073 | 0.0064 | 0.0099 | 0.0044 | 0.0064 | 0.0065 | 0.0076 | 0.0032 | 0.0083 |
| **Median** | 0.0097 | 0.009 | 0.0129 | 0.0106 | 0.0069 | 0.0092 | 0.0064 | 0.0068 | 0.0122 |
| **Maximum** | 0.2028 | 0.111 | 0.295 | 0.2477 | 0.2244 | 0.3479 | 0.2927 | 0.1537 | 0.1768 |
| **Minimum** | -0.243 | -0.1658 | -0.2217 | -0.2783 | -0.2781 | -0.2079 | -0.2509 | -0.1774 | -0.2474 |
| **Std. Dev.** | 0.0549 | 0.037 | 0.0631 | 0.0546 | 0.0609 | 0.0585 | 0.0625 | 0.0494 | 0.0471 |
| **Skewness** | -0.5581 | -0.8419 | -0.0076 | -0.8846 | -0.6426 | 0.3312 | 0.3391 | -0.2122 | -0.8401 |
| **Kurtosis** | 6.1512 | 5.9808 | 5.3271 | 7.6242 | 5.7885 | 9.0413 | 5.7517 | 3.2355 | 6.6848 |
| **Sum** | 2.4918 | 2.1856 | 3.3908 | 1.5228 | 2.1868 | 2.2239 | 2.6127 | 1.11 | 2.8501 |
| **Sum Sq. Dev.** | 1.0301 | 0.4665 | 1.3598 | 1.0235 | 1.2668 | 1.171 | 1.3377 | 0.8332 | 0.758 |
| **Observations** | 343 | 342 | 342 | 344 | 343 | 343 | 343 | 342 | 343 |

*Panel B*

| | N. America Large cap Index | Ex-N. America Large cap Index | US Large cap Index | Canada Large cap Index | France Large cap Index | Germany Large cap Index | Italy Large cap Index | Japan Large cap Index | UK Large cap Index |
|---|---|---|---|---|---|---|---|---|---|
| **Mean** | 0.0065 | 0.004 | 0.0071 | 0.0058 | 0.0045 | 0.007 | 0.0031 | 0.0029 | 0.003 |
| **Median** | 0.01 | 0.0055 | 0.0122 | 0.0101 | 0.0094 | 0.0096 | 0.0043 | 0.0073 | 0.0076 |
| **Maximum** | 0.0876 | 0.134 | 0.1268 | 0.127 | 0.2012 | 0.2138 | 0.255 | 0.1614 | 0.1235 |
| **Minimum** | -0.1551 | -0.1219 | -0.1694 | -0.1896 | -0.1749 | -0.2542 | -0.2369 | -0.2383 | -0.1381 |
| **Std. Dev.** | 0.0304 | 0.036 | 0.0434 | 0.0425 | 0.053 | 0.0597 | 0.0638 | 0.0553 | 0.0391 |
| **Skewness** | -1.0828 | -0.4877 | -0.6003 | -0.824 | -0.27 | -0.478 | 0.1439 | -0.319 | -0.5714 |
| **Kurtosis** | 6.5103 | 4.3628 | 4.0316 | 5.8298 | 3.7893 | 4.8695 | 4.4173 | 3.6026 | 3.9209 |
| **Sum** | 2.2341 | 1.3807 | 2.4391 | 2.0072 | 1.5319 | 2.4054 | 1.0497 | 0.9981 | 1.0118 |
| **Sum Sq. Dev.** | 0.3172 | 0.4399 | 0.6454 | 0.6194 | 0.9617 | 1.2201 | 1.3829 | 1.0458 | 0.5226 |
| **Observations** | 343 | 341 | 343 | 344 | 343 | 343 | 342 | 343 | 343 |

*CPI Data and Inflation Regime Threshold*

The descriptive statistics of the 12 months change of CPI for each country shows mean values of 1.68% at the least for France, and 5.38% at the highest for Italy. The highest maximum value reported in the sample is 25% for Japan, and the lowest is France with 6.08%. Italy has the greatest standard deviation values and France has the lowest. France does have the most rangebound values for CPI with a [min, max] of [-0.007255, 0.07912], the highest is the United States with [-0.1074, 0.1967], the inflation figures for the United States go back to 1928 while the ex-North American sample is from 1994. Full descriptive statistics for the datasets are found in Table 3.

For our analysis of high inflation and low inflation regimes, we choose the 3rd quartile 12 month in CPI from the complete CPI dataset range by country as the threshold between the two regimes, representing approximate target rates for central banks. The value for each country threshold is found in Table 4. The range of 3rd quartile CPI threshold values is from 1.7297% for Japan to 4.67% for Italy.

**Table 3 12-Month Change in CPI By Country, full date ranges by country, monthly data.**

| | US CPI | Canada CPI | Japan CPI | France CPI | Germany CPI | Italy CPI | UK CPI |
|---|---|---|---|---|---|---|---|
| **Mean** | 0.0312 | 0.0305 | 0.024 | 0.0169 | 0.0185 | 0.0538 | 0.0271 |
| **Median** | 0.0269 | 0.0228 | 0.0085 | 0.0169 | 0.0161 | 0.0343 | 0.0237 |
| **Maximum** | 0.1967 | 0.1753 | 0.25 | 0.0608 | 0.0791 | 0.2524 | 0.101 |
| **Minimum** | -0.1074 | -0.1196 | -0.0256 | -0.0073 | -0.0054 | -0.0201 | -0.0012 |
| **Std. Dev.** | 0.0398 | 0.0385 | 0.0432 | 0.0103 | 0.0135 | 0.0545 | 0.0184 |
| **Skewness** | 0.1676 | 0.1262 | 2.7991 | 0.5077 | 1.7363 | 1.5491 | 1.3684 |
| **Kurtosis** | 6.3855 | 5.6611 | 12.5052 | 3.9958 | 7.0089 | 4.6536 | 4.9592 |
| **Observations** | 1120 | 1136 | 619 | 416 | 416 | 812 | 440 |

**Table 4 - Threshold 3rd Quintile of 12-month percentage change in CPI 12 by country**

| Country | 3rd Quintile CPI | Sample Date Range |
|---|---|---|
| **United States** | 3.20230 | 1928M06 - 2022M08 |
| **Canada** | 2.89017 | 1927M11 - 2022M06 |
| **Japan** | 1.72970 | 1970M01 - 2022M07 |
| **France** | 0.03400 | 1955M01 - 2022-M07 |
| **Germany** | 0.02033 | 1948M07 - 2022-M07 |
| **Italy** | 0.04666 | 1954M01 - 2022M08 |
| **United Kingdom** | 0.04435 | 1960M01 - 2022M07 |

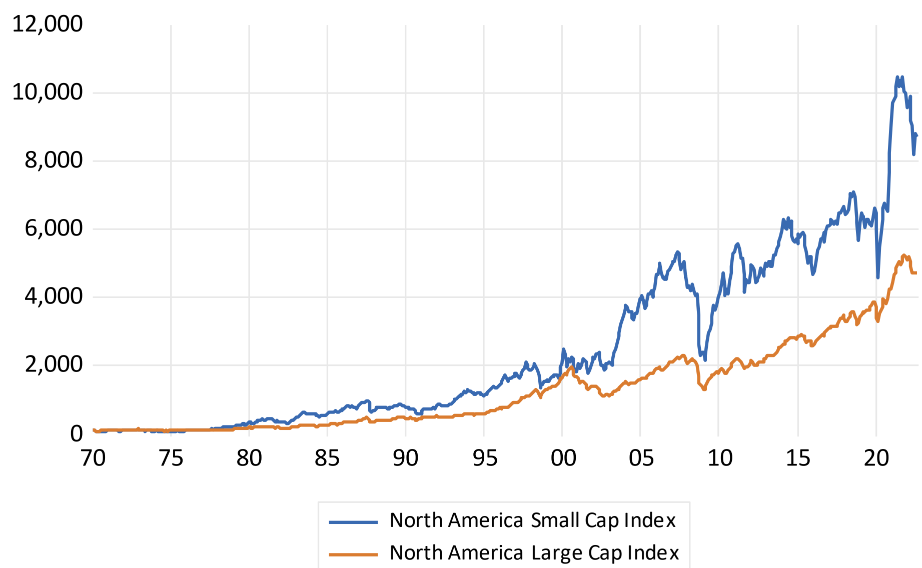
**Table 4 - Threshold 3rd Quintile of 12-month percentage change in CPI 12 by country**

| Country | 3rd Quintile CPI | Sample Date Range |
|---|---|---|
| **United States** | 3.20230 | 1928M06 - 2022M08 |
| **Canada** | 2.89017 | 1927M11 - 2022M06 |
| **Japan** | 1.72970 | 1970M01 - 2022M07 |
| **France** | 3.39980 | 1955M01 - 2022-M07 |
| **Germany** | 2.03330 | 1948M07 - 2022-M07 |
| **Italy** | 4.66610 | 1954M01 - 2022M08 |
| **United Kingdom** | 4.43520 | 1960M01 - 2022M07 |

## Growth in Small-cap and Large-cap Indices with Inflation Regimes (Graphs)

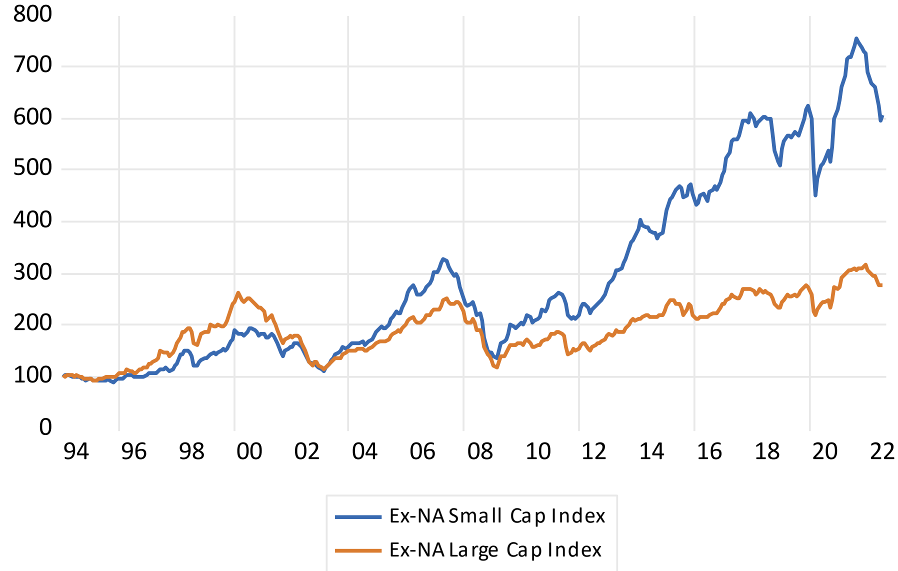
The large-cap and small-cap index monthly prices for each country are plotted in the line graphs Figure 1 to 9, with the high inflation periods highlighted. The high inflation periods are defined as the 12-month change in CPI beyond the threshold 3rd Quartile level from Table 5. The graphs show that G7 countries outside of North America have a greater divergence between their small-cap and large-cap indices after the Global Financial Crisis (GFC). The Canada & US Indices for small-cap prices are shown to be more volatile. For most countries, the price levels for small-caps converge onto the large-caps during the GFC. The German indices appear to track one another and not show as much divergence between large-caps and small-cap prices as the rest of the G7 countries.

Figure 1 Monthly Prices for North America Small Cap and Large Cap Indices 1970M01-2022M08



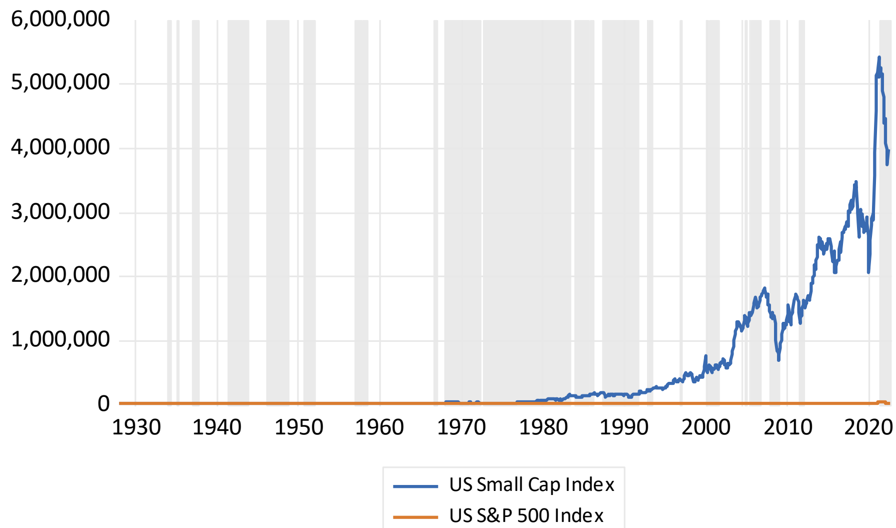

*Monthly large-cap and small-cap stock index prices for North America ex-Mexico including Canada and the United States of America in local currency, equally weighted in local currency. Small-cap stocks track large-caps until around 2002 and then outperform more significantly after 2008.*

Figure 2  Monthly Prices for Ex-North America G7 Small Cap Indices and Large Cap Indices 1994M01-2022M07



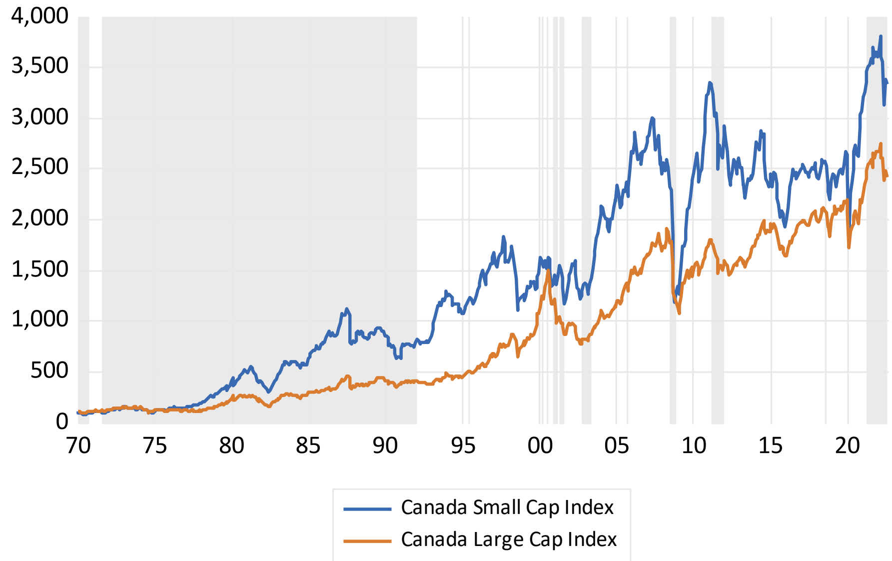

*Monthly prices for ex-North American countries in the G7 include Italy, Germany, the United Kingdom, Japan. Equally weighted index of goods prices in local currencies. Large-cap Equities outperform Small-caps until after 2002, but Small-cap prices experience greater growth after 2008.*

Figure 3  Monthly Prices for United States S&P 500 Index and Small Cap Index 1928M02-2022M07
Shaded Area = 12 Month Change in CPI > 3.2023%



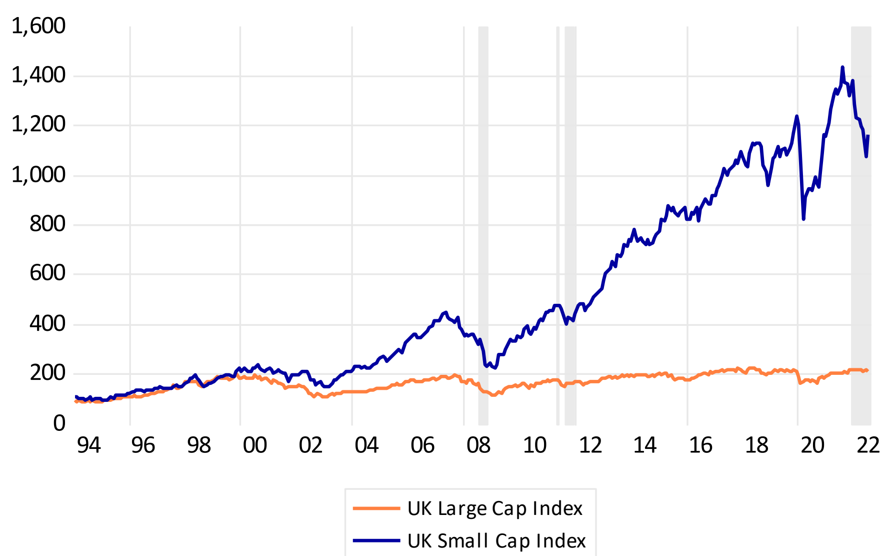

*Monthly Prices for US Large-cap S&P 500 Index and US DFA Small-cap Index and CRSP Small-cap Index, followed by log prices of each index from 1956M01-2022M08 in USD, high inflation shaded in grey. Small-cap stocks outperform large-caps throughout the sample period, it is observed in the log prices that outside of the high inflation decades of the 1970-1980s,high inflation has a downward pressure on prices.*

Figure 4 Monthly Prices for Canada Small Cap and Large Cap Stock Indices 1970M01 - 2022M08
Shaded Area = 12 Month Change in CPI > 2.890173%



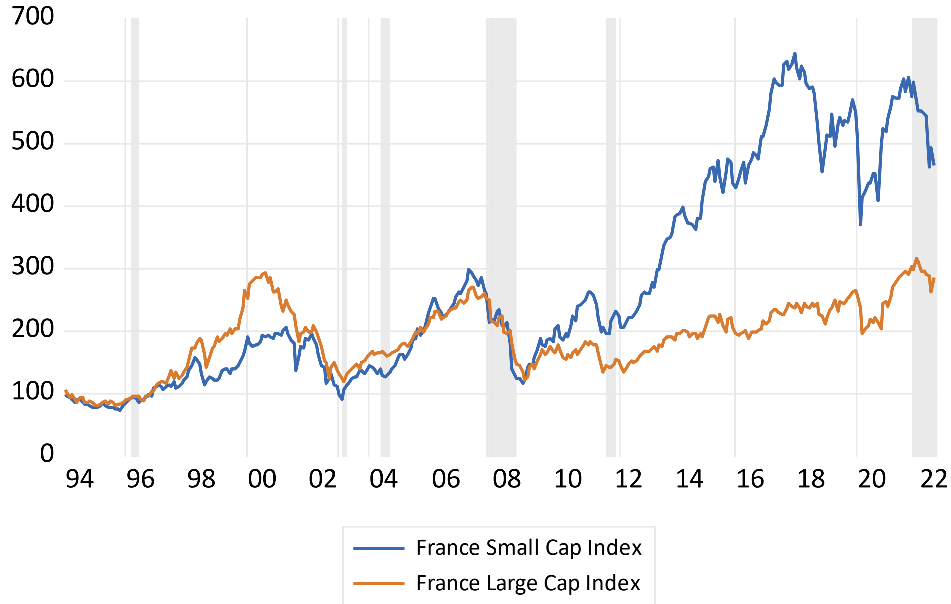

*Monthly Prices for Canada Small-cap Index and S&P/TSX Composite Index 1970M01-2022M08 in CAD, periods of high inflation shaded in grey. Small-cap prices outperform large-cap prices but converge during the 2000, 2008, and 2020 market crashes. Periods of high inflation have downward pressure on prices for both indices after the 1990s.*

Figure 5 Monthly Prices for United Kingdom Large Cap and Small Cap Stock Indices 1994M01-2022M08
Shaded Area = 12 month change in CPI > 4.4352%



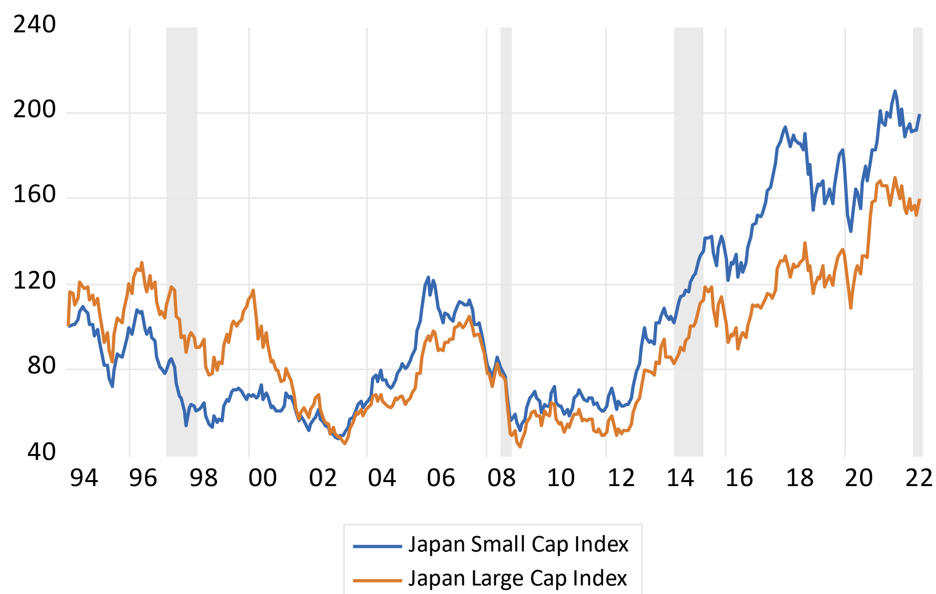

*Monthly price stock Indices for the United Kingdom 1994M01-2022M08 S&P United Kingdom Small-cap Index and the S&P United Kingdom Large-cap Index in local currency, high inflation shaded in grey. Small-cap prices outpace Large-cap prices after 1998 and appear under pressure during high inflation.*

Figure 6  Monthly Prices for France Large Cap and Small Cap Indices 1994M01-2022M08
Shaded Area = 12 month change in CPI>3.3998%



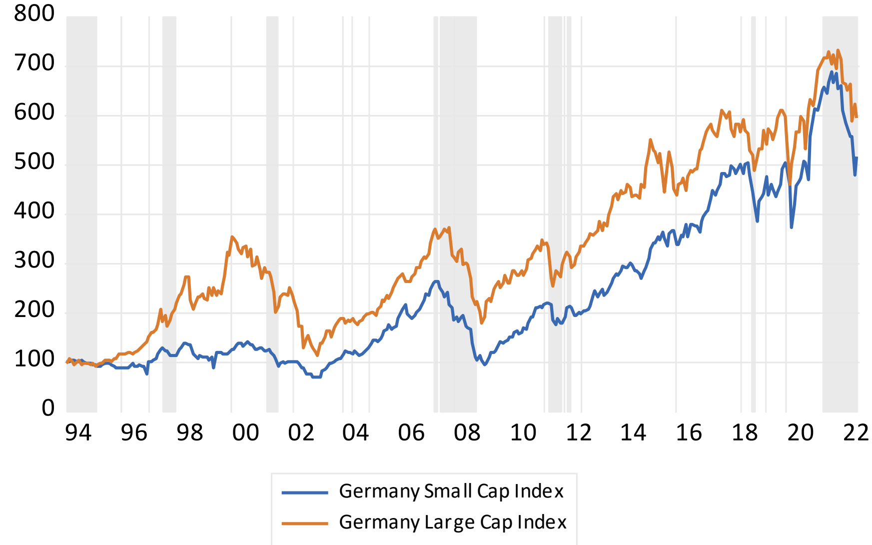

*Monthly price stock indices for France 1994M01-2022M08 MSCI France Small-cap Index, FTSE France CAC Index in local currency, high inflation shaded in grey. Small-caps underperform Large-cap prices until 2008 and then the price widens. Both indices see downward pressure during high inflation periods.*

Figure 7  Monthly Prices for Japan Large Cap and Small Cap Stock Indices 1994M01-2022M08
Shaded Area = 12 month change in CPI>1.7297%



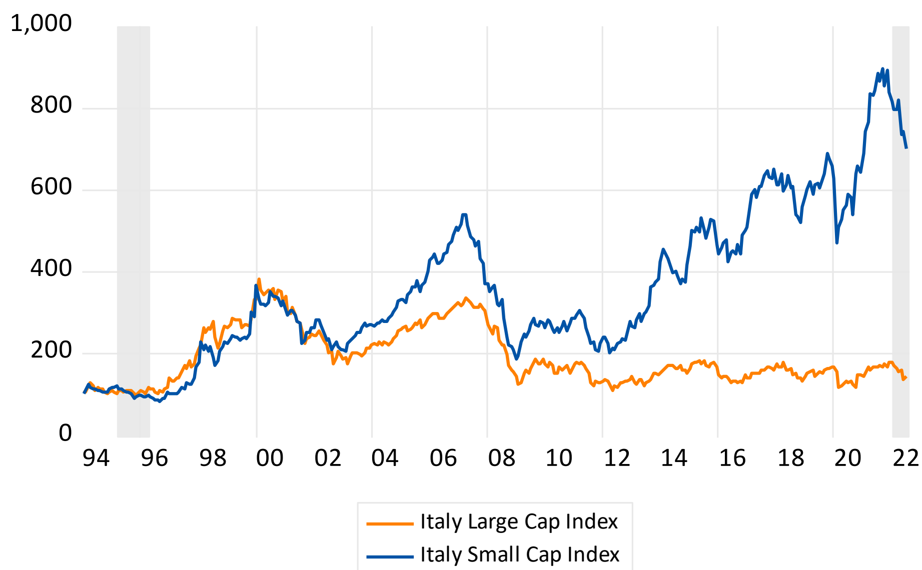

*Monthly price stock indices for Japan 1994M01-2022M08 Nikkei 225 Index, the S&P Japan Small-cap Index in local currencies, high inflation shaded in grey. Small-cap prices underperform from 1994-2000 and then track Large-cap prices until 2008 where Small-cap prices outperform consistently. High inflation periods have downward pressure on both indices except for in 2014.*

Figure 8  Monthly Prices for Germany Large Cap and Small Cap Stock Indices 1994M - 2022M08
Shaded Area = 12 month change in CPI > 2.0333%




*Monthly price stock indices 1994M01-2022M08 FTSE DAX Large-cap Index and FTSE SDAX Small-cap Index in local currency, high inflation shaded in grey. Large-cap prices outperform Small-cap prices throughout the sample period. High inflation has downward pressure on both price indices.*

Figure 9  Monthly Prices Italy Large Cap and Small Cap Stock Indices 1994M01-2022M08
Shaded Area = 12 Month Change in CPI>4.6661%




*Monthly price stock indices 1994M01-2022M08 FTSE Italy Large-cap Index and MSCI  Italy Small-cap Index in local currency, periods of  high inflation shaded in grey. Large-cap and small-cap prices track each other from 1994 to 2002 and Small-caps subsequently outperform. High inflation periods appear to put downward pressure on both indices*

## Empirical Results

*Markov Switching Regression Model*

The Markov Switching Model regression is estimated with the small-cap premium as the dependent variable, the 12-month change in CPI and the default risk being the switching independent variables, while dividend yield, term risk and the risk-free rate are non-switching regressors.

For the US, the results show that the coefficients for the US CPI and the US default risk are both significant at the 5% level in Regime state 2, both having negative signs. In regime state 1 there are no significant coefficients. The non-regressing variables all have significant coefficients, with the US dividend yield having a positive coefficient of 0.3109, and Term risk having a slightly negative coefficient of -0.0678 (Table 5). Regime state 1 for the small-cap premium has an expected duration of 32.36 month and has greater stability than state 2 when already in that state, 0.9691 versus 0.2859 (Table 6). From these results we can infer that regime 1 represents good economic times, as this fits with the actual economic state having extended stock market growth in the sample period. The graphs seen in Figure 10 show that events triggering a switch to the second regime from the first fit with economic events such as the GFC, and the market pullback in 2018. The period corresponding to the COVID-19 crash does not show a regime state change in this model. These results suggest that small-cap premiums are more affected by CPI and Default risk during weaker economic states than in strong ones, where there is no significant effect at all. The risk-free rate has a strongly negative relationship with small-cap premiums, which fits with the consensus that small-caps are strongly affected by interest rates. The weakly positive relationship with the US dividend yield suggests that ~~when the overall market is doing well, and~~

~~can~~ a market that can return cashflows to investors has a positive impact on the small-cap premium.~~, the small-cap premium is likely to benefit positively.~~

**Table 5 Markov Switching Model Regression Output for US 2003M06-2022M08**

This table shows the output from the Markov Switching model regression estimation for the US, the small cap premium is the dependent variable, the percent 12 month change in CPI, and default risk are switching regressors, while the dividend yield, term risk, and risk free rate are non-switching regressors.

*Regime 1*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **US CPI % change** | -0.1091 | 0.0941 | -1.1594 | 0.2463 |
| **US Default Risk** | 0.0305 | 0.2616 | 0.1168 | 0.9070 |
| **C** | 0.0005 | 0.0052 | 0.0892 | 0.9290 |
| **LOG(SIGMA)** | -3.7090 | 0.0490 | -75.6980 | 0.0000 |

*Regime 2*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **US CPI % change** | -0.1320 | 0.0142 | -9.2915 | 0.0000 |
| **US Default Risk** | -0.2120 | 0.0128 | -16.6206 | 0.0000 |
| **C** | 0.0063 | 0.0004 | 17.3052 | 0.0000 |
| **LOG(SIGMA)** | -8.9746 | 0.2476 | -36.2512 | 0.0000 |

*Common Variables*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **US Dividend Yield** | 0.3109 | 0.0160 | 19.4804 | 0.0000 |
| **US Term Risk** | -0.0678 | 0.0108 | -6.2567 | 0.0000 |
| **Risk Free Rate** | -2.1004 | 0.0650 | -32.3180 | 0.0000 |

**Table 6 Constant Markov Transition probabilities and expected duration for the US from 2003M06-2022M08**

This table shows the probability of staying in regime state 1 if already in state 1, and same for state 2, as well as the expected duration of each state.

*Transition probabilities*

| | 1 | 2 |
|---|---|---|
| **1** | 0.9691 | 0.0309 |
| **2** | 0.7141 | 0.2859 |

*Expected durations*

| | 1 | 2 |
|---|---|---|
| | 32.3632 | 1.4004 |

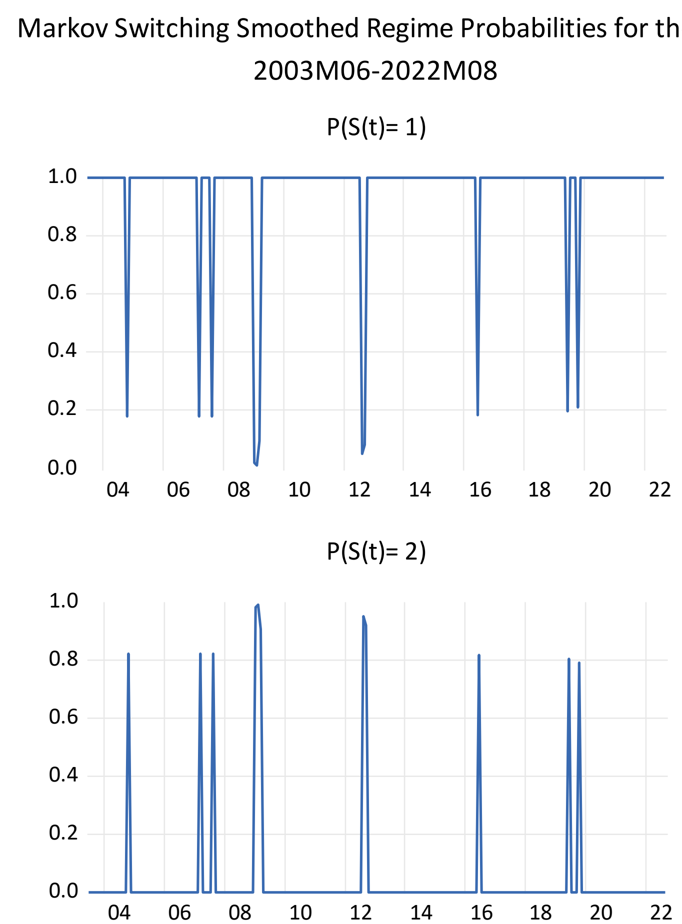



*Figure 10 – Markov Switching Smoothed Regime Probabilities for the US from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2.The multiple regime state changes are indicative of the somewhat unstable second state probabilities.*

The regression outputs~~results~~ for ~~for~~ Canada can be found in Table 7.~~.~~ The regime states are more stable than those for the US, shown in Table 8, regime 1 seems to correspond to a weak economic state, with a duration of 5 months ~~versus~~compared with regime 2 which has a duration of 82 months. There is a significant coefficient for Canada CPI of -10.0836 indicating a strongly negative relationship with the small-cap premium in regime 1. In regime 2, the coefficient for CPI is negative and significant but at -0.7455 indicates that it has a much weaker relationship. Default risk is not significant factor in either regime. Only the dividend yield is significant from the non-switching regressors, which is strongly negative at -9.4795 indicating that when the market is returning cash to investors, the small-cap premium is severely negatively impacted. ~~The regime states are more stable than those for the US, shown in Table 8, regime 1 seems to correspond to a weak economic state, with a duration of 5 months versus regime 2 which has a duration of 82 months.~~ The regime probabilities graph in Figure 10 shows that the two regime switches occur during the GFC, and COVID-19 crash indicating that the small-cap regime state changes occurred only during economic crises.

**Table 7 Markov Switching Model Regression Output for Canada 2003M06-2022M08**

This table shows the output from the Markov Switching model regression estimation for Canada, the small cap premium is the dependent variable, the percent 12 month change in CPI, and default risk are switching regressors, while the dividend yield, term risk, and risk free rate are non-switching regressors.

*Regime 1*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Canada CPI % change** | -10.0836 | 3.8624 | -2.6107 | 0.0090 |
| **Canada Default Risk** | 0.0150 | 0.0279 | 0.5384 | 0.5903 |
| **C** | 0.1471 | 0.0565 | 2.6038 | 0.0092 |
| **LOG(SIGMA)** | -2.3006 | 0.2141 | -10.7446 | 0.0000 |

*Regime 2*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Canada CPI % change** | -0.7455 | 0.2455 | -3.0368 | 0.0024 |
| **Canada Default Risk** | 0.0054 | 0.0108 | 0.5000 | 0.6170 |
| **C** | 0.0413 | 0.0163 | 2.5341 | 0.0113 |
| **LOG(SIGMA)** | -3.1901 | 0.0511 | -62.4737 | 0.0000 |

*Common Variables*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Canada Risk Free Rate** | -0.1599 | 0.2712 | -0.5897 | 0.5554 |
| **Canada Term Risk** | 0.3463 | 0.3962 | 0.8741 | 0.3821 |
| **Canada Dividend Yield** | -9.4795 | 4.4382 | -2.1359 | 0.0327 |

**Table 8 Constant Markov Transition probabilities and expected duration for Canada from 2003M06-2022M08**

This table shows the probability of staying in regime state 1 if already in state 1, and same for state 2, as well as the expected duration of each state.

*Transition probabilities*

| | **1** | **2** |
|---|---|---|
| **1** | 0.8037 | 0.1963 |
| **2** | 0.0122 | 0.9878 |

*Expected durations*

| **1** | **2** |
|---|---|
| 5.0946 | 82.0087 |

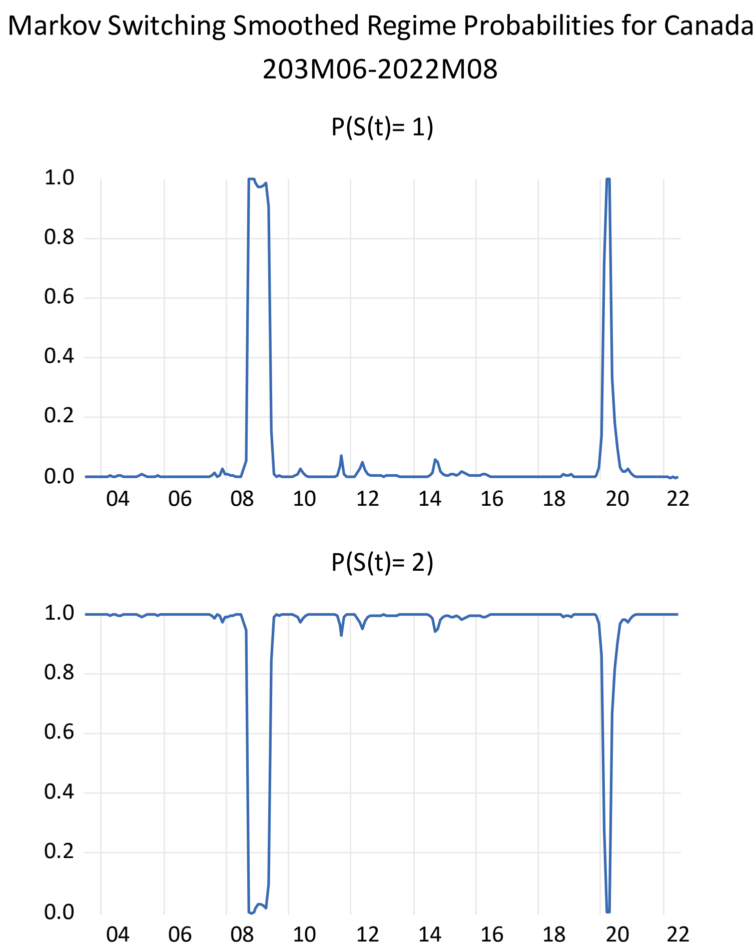



*Figure 11 – Markov Switching Smoothed Regime Probabilities for Canada from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2.There are two state switches which occur during the period of the GFC, and the COVID-19 crash.*

The output for Germany results in a significant and weakly negative coefficient for default risk in regime 1 and regime 2, with a slightly more negative relationship in regime 1. The coefficient for CPI in regime 2 is not significant at 5% but shows a weakly negative relationship with Germany's small-cap premium. The two states are shown to be unstable in Table 10, and in Figure 12, the probability of staying in state 1 is 0.9090 when in state 1, but the probability of staying in state 2 when in state 2 is much lower at 0.48. Regime 1 seems to correspond to stronger economic times, and the state expected duration is 10.98 months versus 1.92 months for state 2, fitting with the current economic reality. The only significant factor of the non-switching regressors is the dividend yield on the DAX index, which has a strongly negative coefficient of -4.892. Removing insignificant variables from the common variables did not improve the stability of the model.

**Table 9 Markov Switching Model Regression Output for Germany 2003M06-2022M08**

This table shows the output from the Markov Switching model regression estimation for Germany, the small cap premium is the dependent variable, the percent 12 month change in CPI, and default risk are switching regressors, while the dividend yield, term risk, and risk free rate are non-switching regressors.

*Regime 1*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Germany CPI % change** | -0.3162 | 0.4329 | -0.7304 | 0.4652 |
| **Germany Default Risk** | -0.1032 | 0.0312 | -3.3076 | 0.0009 |
| **C** | 0.0021 | 0.0096 | 0.2160 | 0.8290 |
| **LOG(SIGMA)** | -2.5671 | 0.0559 | -45.8929 | 0.0000 |

*Regime 2*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Germany CPI % change** | -0.2349 | 0.1648 | -1.4249 | 0.1542 |
| **Germany Default Risk** | -0.0511 | 0.0098 | -5.2246 | 0.0000 |
| **C** | -0.0025 | 0.0047 | -0.5419 | 0.5879 |
| **LOG(SIGMA)** | -4.8492 | 0.2198 | -22.0624 | 0.0000 |

*Common Variables*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Germany Risk Free Rate** | -0.0322 | 0.2765 | -0.1164 | 0.9073 |
| **Germany Term Risk** | 0.0008 | 0.0014 | 0.5852 | 0.5584 |
| **Germany Dividend Yield** | 3.0293 | 0.5459 | 5.5496 | 0.0000 |

**Table 10 Constant Markov Transition probabilities and expected duration for Germany from 2003M06-2022M08**

This table shows the probability of staying in regime state 1 if already in state 1, and same for state 2, as well as the expected duration of each state.

*Transition probabilities*

| | 1 | 2 |
|---|---|---|
| **1** | 0.9090 | 0.0910 |
| **2** | 0.5191 | 0.4809 |

*Expected durations*

| 1 | 2 |
|---|---|
| 10.9886 | 1.9265 |

Markov Switching Smoothed Regime Probabilities for Germany
2003M06-2022M08

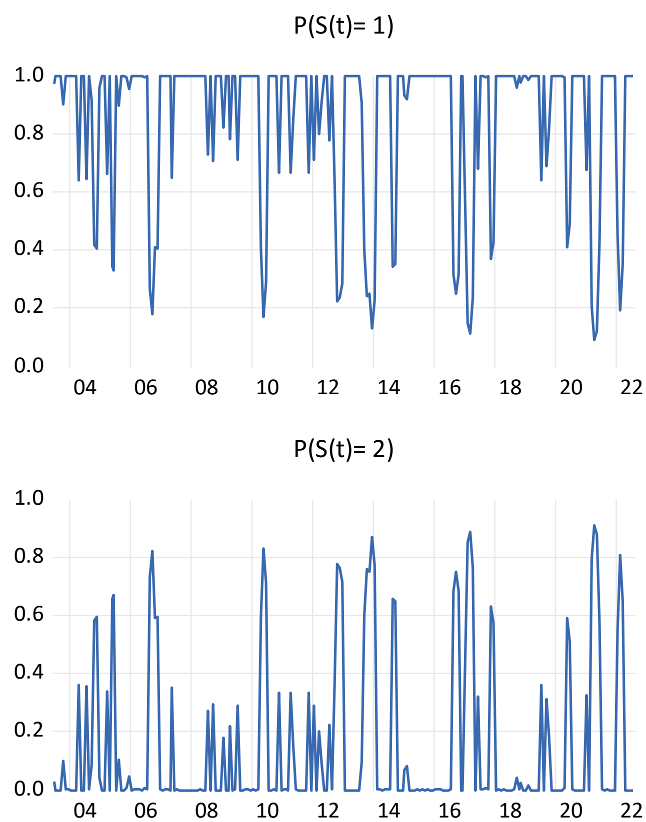



*Figure 12 – Markov Switching Smoothed Regime Probabilities for Germany from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show that the two states for Germany's small-cap premium are highly unstable.*

Removing the term risk and risk free rate variables for the model for Italy greatly increased the stability. The results output in Table 11 for Italy show that both CPI and default risk variables are significant for regime 1, though the default risk coefficient is near zero, CPI does show a positive relationship with the small-cap premium. For regime 2, neither coefficient is significant but are both weakly negative. Despite this significant result in regime 1, in Table 12, the regime state probabilities show that regime 1 is highly unstable at near zero, while the stability of regime 2 is very high at 0.9707 which indicates the effect of CPI and default risk on the small-cap premium in regime 1 is likely to be short lived. Regime 1 seems to be associated with a weak economic state, while regime 2 is likely to correspond to stronger economic states, with an expected duration of 1 and 34 months respectively. The state switches shown in Figure 12 appear to correspond with the GFC and COVID-19, with additional state changes occurring pre-GFC which could perhaps be attributed to post Euro instability. Dividend yield has a strong negative relationship with small-cap premiums regardless of state.

**Table 11 Markov Switching Model Regression Output for Italy 2003M06-2022M08**

This table shows the output from the Markov Switching model regression estimation for Italy, the small cap premium is the dependent variable, the percent 12 month change in CPI, and default risk are switching regressors, while the dividend yield, term risk, and risk free rate are non-switching regressors.

*Regime 1*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Italy CPI % change** | 0.9785 | 0.0099 | 99.2308 | 0.0000 |
| **Italy Default Risk** | 0.0014 | 0.0001 | 16.8095 | 0.0000 |
| **C** | 0.0264 | 0.0002 | 173.5101 | 0.0000 |
| **LOG(SIGMA)** | -9.0689 | 0.3199 | -28.3493 | 0.0000 |

*Regime 2*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Italy CPI % change** | -0.0550 | 0.3933 | -0.1398 | 0.8888 |
| **Italy Default Risk** | -0.0033 | 0.0040 | -0.8145 | 0.4154 |
| **C** | 0.0179 | 0.0084 | 2.1235 | 0.0337 |
| **LOG(SIGMA)** | -2.5072 | 0.0492 | -50.9793 | 0.0000 |

*Common Variables*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Italy Dividend Yield** | -3.8062 | 0.0134 | -283.7228 | 0.0000 |

**Table 12 Constant Markov Transition probabilities and expected duration for Italy from 2003M06-2022M08**

This table shows the probability of staying in regime state 1 if already in state 1, and same for state 2, as well as the expected duration of each state.

*Transition probabilities*

| | **1** | **2** |
|---|---|---|
| **1** | 0.0001 | 0.9999 |
| **2** | 0.0293 | 0.9707 |

*Expected durations*

| **1** | **2** |
|---|---|
| 1.0001 | 34.0843 |

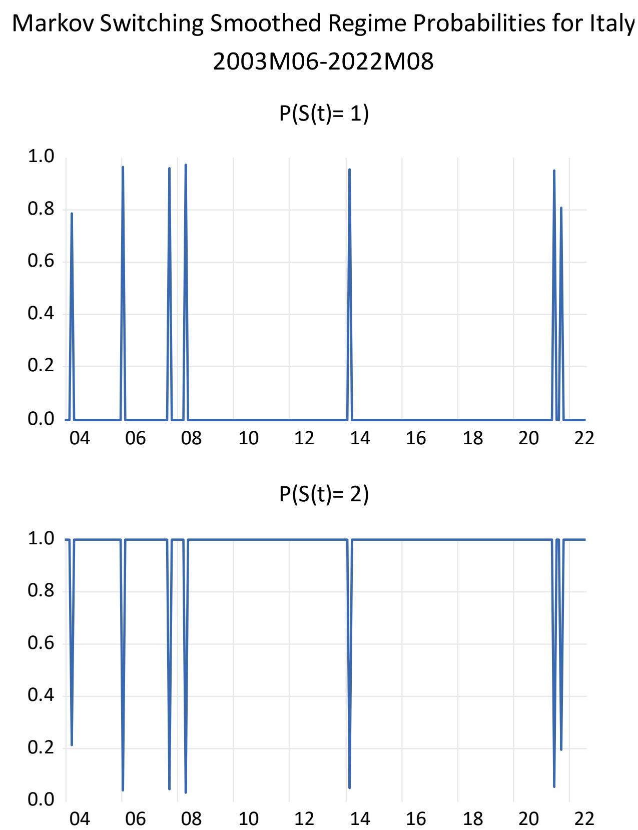



*Figure 13 – Markov Switching Smoothed Regime Probabilities for Italy from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes during the period corresponding with the GFC and the COVID-19 crash.*

The default risk variable was not included in the model for Japan as the data available for the time period were not suitable for this analysis. Also, upon removing the risk free rate from the non-switching regressors, the stability of the model was greatly improved and so it was omitted. Results in Table 13 show that the CPI coefficient is positive and significant in regime 2, at 0.3542, but is not significant and less positive in regime 1. The dividend yield and term risk variables were both significant, though the term risk factor has a near zero coefficient, the dividend yield has a strongly positive coefficient of 3.1709. These results must be considered in the light that regime 2 is highly unstable. Table 14 shows that the probability of staying in regime 1 is 0.9782 while for regime 2 it is nearly 0. The expected duration of 45 months for regime 1 relative to 1 month for regime 2 confirms that regime 1 is likely associated with a stronger economic state. The state switch probabilities shown in Figure 14 confirm this. The strongly positive relationship between the dividend yield and small-cap premium indicate that a market returning cashflow to investors is strongly positive signal for small-cap relative performance.

**Table 13 Markov Switching Model Regression Output for Japan 2003M06-2022M08**

This table shows the output from the Markov Switching model regression estimation for Japan, the small cap premium is the dependent variable, the percent 12 month change in CPI, and default risk are switching regressors, while the dividend yield, term risk, and risk free rate are non-switching regressors.

*Regime 1*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Japan CPI % change** | 0.1370 | 0.1568 | 0.8739 | 0.3822 |
| **C** | -0.0079 | 0.0017 | -4.6924 | 0.0000 |
| **LOG(SIGMA)** | -3.7364 | 0.0482 | -77.5219 | 0.0000 |

*Regime 2*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Japan CPI % change** | 0.3642 | 0.0001 | 2485.6570 | 0.0000 |
| **C** | 0.0001 | 0.0000 | 63.7175 | 0.0000 |
| **LOG(SIGMA)** | -13.1563 | 0.3480 | -37.8028 | 0.0000 |

*Common Variables*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **Japan Dividend Yield** | 3.1709 | 0.0061 | 520.9576 | 0.0000 |
| **Japan Term Risk** | 0.6169 | 0.0002 | 3555.9870 | 0.0000 |

**Table 14 Constant Markov Transition probabilities and expected duration for Japan from 2003M06-2022M08**

This table shows the probability of staying in regime state 1 if already in state 1, and same for state 2, as well as the expected duration of each state.

*Transition probabilities*

| | 1 | 2 |
|---|---|---|
| **1** | 0.9782 | 0.0218 |
| **2** | 0.9996 | 0.0004 |

*Expected durations*

| 1 | 2 |
|---|---|
| 45.7686 | 1.0004 |

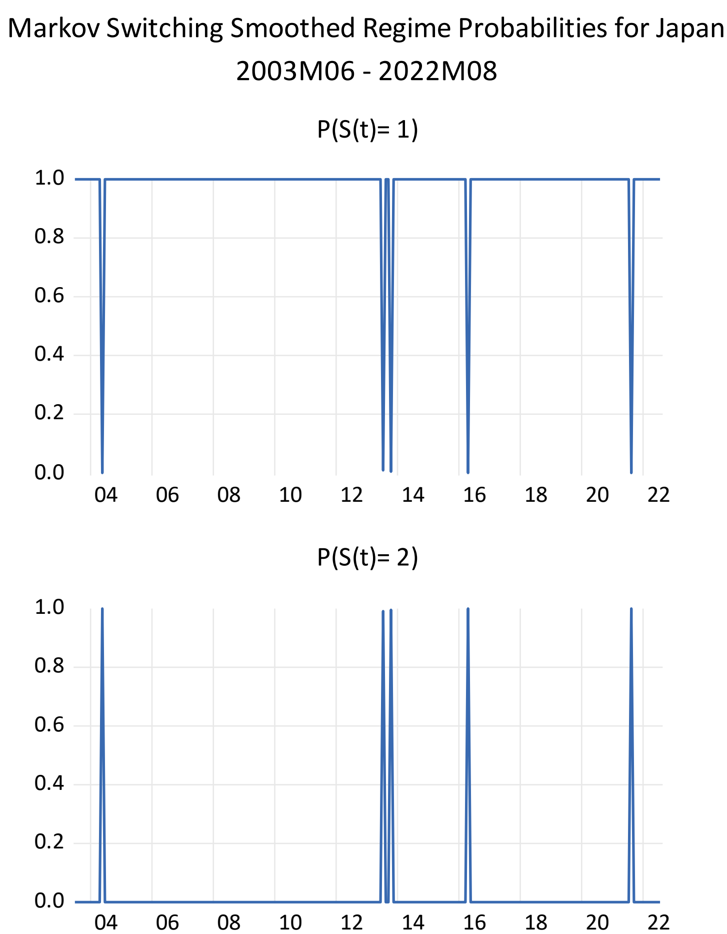



*Figure 14 – Markov Switching Smoothed Regime Probabilities for Japan from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes during the period corresponding with the GFC and the COVID-19 crash.*

Table 16 shows that the UK regimes may not be meaningful as regime 1 has an expected duration of 1.5410 and 1.1361 for regime 2 and the probability of switching states is higher than the probability of staying in the same state for both regimes. The model did not improve in stability by removing variables. In Table 15, results show that the coefficient for CPI in Regime 1 is significant and negative, while in regime 2 it is not significant but positive. No other coefficient is significant, but it may be interesting to note that default risk is positively related to small-cap premiums in both regimes.

**Table 15 Markov Switching Model Regression Output for the UK 2003M06-2022M08**

This table shows the output from the Markov Switching model regression estimation for the UK, the small cap premium is the dependent variable, the percent 12 month change in CPI, and default risk are switching regressors, while the dividend yield, term risk, and risk free rate are non-switching regressors.

*Regime 1*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **UK CPI % change** | -0.9235 | 0.3592 | -2.5713 | 0.0101 |
| **UK Default Risk** | 0.3193 | 0.5348 | 0.5970 | 0.5505 |
| **C** | 0.0036 | 0.0173 | 0.2105 | 0.8333 |
| **LOG(SIGMA)** | -2.9180 | 0.0793 | -36.8197 | 0.0000 |

*Regime 2*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **UK CPI % change** | 0.4954 | 0.3252 | 1.5233 | 0.1277 |
| **UK Default Risk** | 0.7138 | 0.5668 | 1.2593 | 0.2079 |
| **C** | 0.0230 | 0.0160 | 1.4309 | 0.1525 |
| **LOG(SIGMA)** | -3.3078 | 0.1123 | -29.4421 | 0.0000 |

*Common Variables*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **UK Dividend Yield** | -1.7787 | 1.5873 | -1.1206 | 0.2625 |
| **UK Risk Free Rate** | -0.4530 | 0.5537 | -0.8181 | 0.4133 |
| **UK Term Risk** | -0.0039 | 0.5867 | -0.0066 | 0.9947 |

**Table 16 Constant Markov Transition probabilities and expected duration for the UK from 2003M06-2022M08**

This table shows the probability of staying in regime state 1 if already in state 1, and same for state 2, as well as the expected duration of each state.

| *Transition probabilities* | | |
|---|---|---|
| | **1** | **2** |
| **1** | 0.3511 | 0.6489 |
| **2** | 0.8802 | 0.1198 |
| *Expected durations* | | |
| | **1** | **2** |
| | 1.5410 | 1.1361 |

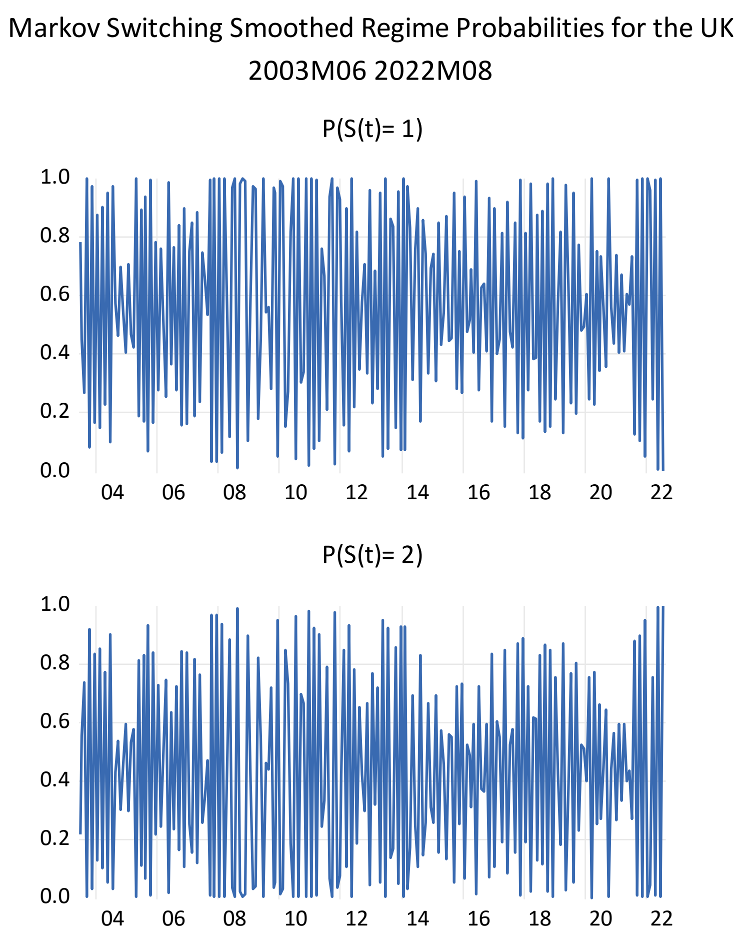



*Figure 15 – Markov Switching Smoothed Regime Probabilities for the UK from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes are highly unstable throughout the sample period.*

Table 17 shows the model estimation outputs for France. Default risk in regime 2 is significant but near zero, and CPI is significant at the 10% significance level in regime 1 and shows a negative relationship with small-cap premiums. Determining which regime corresponds to a strong or weak economic state is difficult as shown in Table 18 and Figure 16, while regime same state probabilities are somewhat high, the durations of regimes 1 and 2 are 4.0110 and 3.6844 respectively, and the switching probability graph shows little in terms of economic significance. Of the non-switching regressors, dividend yield has a negative relationship with small-cap premiums at 10% significance. Removing variables did not improve the model stability.

**Table 17 Markov Switching Model Regression Output for the France 2003M06-2022M08**

This table shows the output from the Markov Switching model regression estimation for France UK, the small cap premium is the dependent variable, the percent 12 month change in CPI, and default risk are switching regressors, while the dividend yield, term risk, and risk free rate are non-switching regressors.

*Regime 1*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **France CPI % change** | -1.4304 | 0.8019 | -1.7838 | 0.0745 |
| **France Default Risk** | 0.0002 | 0.0006 | 0.3560 | 0.7219 |
| **C** | 0.0258 | 0.0183 | 1.4117 | 0.1580 |
| **LOG(SIGMA)** | -2.4181 | 0.0890 | -27.1571 | 0.0000 |

*Regime 2*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **France CPI % change** | 0.1650 | 0.6425 | 0.2568 | 0.7973 |
| **France Default Risk** | 0.0039 | 0.0010 | 3.8317 | 0.0001 |
| **C** | -0.0064 | 0.0106 | -0.6075 | 0.5435 |
| **LOG(SIGMA)** | -3.4314 | 0.1518 | -22.6015 | 0.0000 |

*Common Variables*

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| **France Dividend Yield** | -1.4332 | 0.7936 | -1.8058 | 0.0709 |
| **France Risk Free Rate** | 0.2363 | 0.2450 | 0.9648 | 0.3346 |
| **France Term Risk** | 0.3349 | 0.4771 | 0.7020 | 0.4827 |

**Table 18 Constant Markov Transition probabilities and expected duration for France from 2003M06-2022M08**

This table shows the probability of staying in regime state 1 if already in state 1, and same for state 2, as well as the expected duration of each state.

*Transition probabilities*

| | 1 | 2 |
|---|---|---|
| **1** | 0.7507 | 0.2493 |
| **2** | 0.2714 | 0.7286 |

*Expected durations*

| 1 | 2 |
|---|---|
| 4.0110 | 3.6844 |

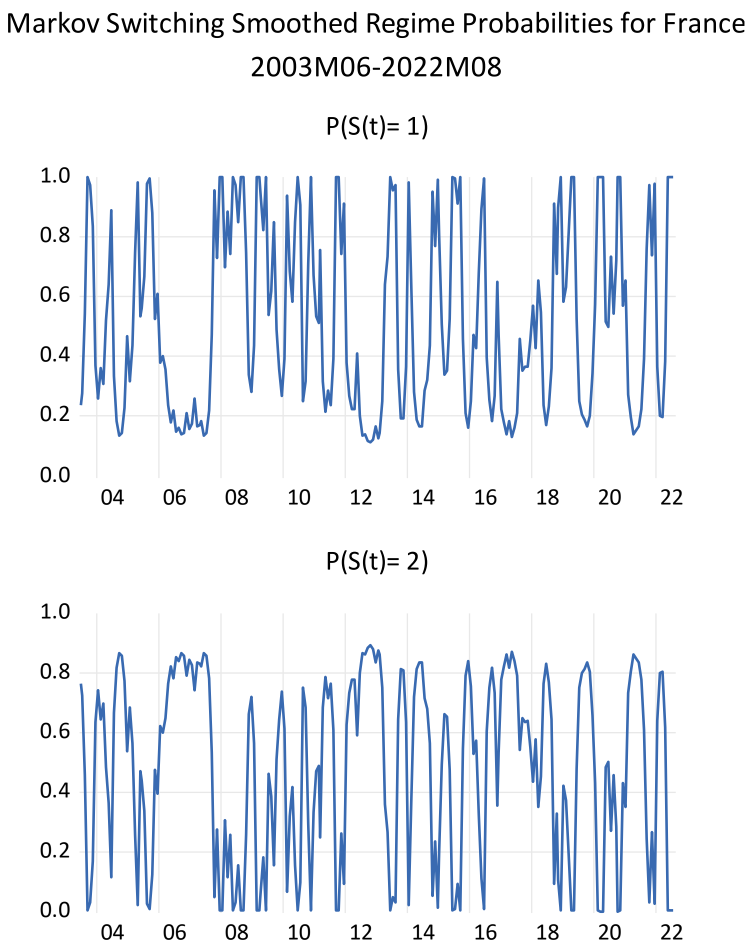



*Figure 16 – Markov Switching Smoothed Regime Probabilities for France from 2003M06-2022M08. The top graph shows the probability of staying in state 1 if already in state 1, while the bottom shows the probability of staying in state 2 when already in state 2. The graphs show regime changes are highly unstable throughout the sample period.*

*Granger Causality*

The two-way Granger Causality test is conducted using lags of 12 as the data in the series are of monthly returns and results are shown in Table 19. The null hypothesis for the Granger Causality test is that the first time series variable does not Granger-Cause the second time series variable, if the p-value of the F-Test is below 5% we can reject the null hypothesis. The series tested are the small-cap premium of each country and the respective 12-month change in CPI. The results show that at 5% significance we only find that the France 12-month change in CPI granger causes the France small-cap premium. At 10% significance, we find that the Canada 12-month change in CPI Granger causes the Canada small-cap premium.

**Table 19 Results from the Granger Causality tests performed for variable pairs of 12-month change in CPI and small cap premiums for each country in the G7 from 1994M01-2022M08**

The output in the table below shows each Granger pair tested, by country, with the corresponding F-statistic and p-value. A p-value less than 0.05 indicates that we can reject the null hypothesis of no Granger Causal relationship directionally from the first variable to the second.

| Country Small cap Premium | Null H0 | F-Statistic (p-value) | P-value |
|---|---|---|---|
| **United States** | The US Small Cap Premium does not Granger Cause US CPI | 0.6484 | 0.7999 |
| | US CPI does not granger cause the US Small Cap Premium | 0.9208 | 0.5263 |
| **Canada** | The Canada Small Cap Premium does not Granger Cause Canada CPI | 1.4771 | 0.1314 |
| | **Canada CPI Does not Granger Cause the Canada Small Cap Premium** | **1.6537** | **0.0763** |
| **Japan** | Japan Small Cap Premium does not Granger Cause Japan CPI | 0.7682 | 0.6831 |
| | Japan CPI does not Granger Cause the Japan Small Cap Premium | 0.6300 | 0.8163 |
| **France** | The France Small Cap Premium does not Granger Cause France CPI | 0.5530 | 0.8786 |
| | **France CPI does not Granger Cause the France Small Cap Premium** | **2.3608** | **0.0065** |
| **Germany** | The Germany Small Cap Premium does not Granger Cause Germany CPI | 0.9293 | 0.5178 |
| | Germany CPI does not Granger Causethe Germany Small Cap Premium | 1.4108 | 0.1594 |
| **Italy** | The Italy Small Cap Premium does not Granger Cause Italy CPI | 0.8738 | 0.5742 |
| | Italy CPI does not Granger Cause the Italy Small Cap Premium | 1.0125 | 0.4372 |
| **UK** | The UK Small Cap Premium does not Granger Cause UK CPI | 1.4369 | 0.1479 |
| | UK CPI does not Granger Cause the UK Small Cap Premium | 0.8408 | 0.6082 |

*Pearson's correlation*

Table 20 shows the Pearson's correlation for the 12-month change in inflation amongst all the G7 countries for the sample period. The results show low correlation throughout; Canada, France and Germany are the highest with Japan having the lowest correlation at 0.0895 implying a diversification benefit with respect to inflation from the G7 countries.

**Table 20 Correlation table for 12 Month Change in CPI for G7 Countries 1994M01-2022M08**

This table shows the Pearson's correlation between each country pair for the sample period.

| | US | Canada | UK | Italy | Japan | France | Germany |
|---|---|---|---|---|---|---|---|
| **US** | 1.0000 | | | | | | |
| **Canada** | 0.7483 | 1.0000 | | | | | |
| **UK** | 0.5353 | 0.4907 | 1.0000 | | | | |
| **Italy** | 0.5693 | 0.3961 | 0.5294 | 1.0000 | | | |
| **Japan** | 0.0895 | 0.1015 | 0.1619 | -0.0386 | 1.0000 | | |
| **France** | 0.7360 | 0.6527 | 0.6287 | 0.7583 | 0.0540 | 1.0000 | |
| **Germany** | 0.7899 | 0.6188 | 0.6889 | 0.6242 | 0.2460 | 0.7738 | 1.0000 |

*OLS for First Differences Last Prices and CPI*

Results of Ordinary Least Squares regression are found in Table 21 below with the country CPI price levels as the independent variable and the first difference stock price index as the dependent variable. The significant **β**x coefficient values for the countries tested at 5% significance are for the United States Small-cap Index, United States Large-cap index, United Kingdom Small-cap Index, and the Italy Small-cap Index. The United States CPI coefficient is near zero for both indices. The Italy CPI coefficient, we see an estimated **β**x of 4.203683 showing a strongly positive vector in the linear relationship implying that Italian CPI prices have a strong positive relation with small-cap returns. The coefficient for the United Kingdom CPI prices shows a weakly positive relation with small-cap prices.

**Table 21 OLS statistics where the first difference of monthly price indices are the dependent variable and CPI is the independent variable for each G7 country, 1994M01-2022M08**

The estimation output shows the regression results from Ordinary Least Squares regression, where the stock price returns time series is the dependent variable and the goods prices are the independent variable for each country.

| Dependent Variable | x | P-value | Sample range | N. obs |
|---|---|---|---|---|
| **US Small cap** | 0.0004 | 0.0069 | 1994M01 – 2022M05 | 342 |
| **US Large cap** | 0.0000 | 0.0032 | 1994M01 - 2022M08 | 344 |
| **Canada Small cap** | 0.0069 | 0.4943 | 1994M01-2022M06 | 342 |
| **Canada Large cap** | 0.0142 | 0.2928 | 1994M01-2022M06 | 342 |
| **Japan Small cap** | 0.0028 | 0.3228 | 1994M01-2022M07 | 342 |
| **Japan Large cap** | 0.0017 | 0.5480 | 1994M01-2022M06 | 341 |
| **France Small cap** | 0.0115 | 0.2766 | 1994M02 -2022M07 | 342 |
| **France Large cap** | 0.0057 | 0.3262 | 1994M01 – 2022M07 | 342 |
| **Germany Small cap** | 0.0137 | 0.1185 | 1994M02 – 2022M07 | 342 |
| **Germany Large cap** | 0.0155 | 0.1808 | 1994M01-2022M07 | 342 |
| **Italy Small cap** | 4.2037 | 0.0000 | 1994M01 – 2022M08 | 342 |
| **Italy Large cap** | -0.0002 | 0.9805 | 1994M01 -2022M08 | 343 |
| **UK Small cap** | 0.0366 | 0.0420 | 1994M02 – 2022M07 | 342 |
| **UK Large cap** | 0.0036 | 0.3493 | 1994M02 -2022M07 | 342 |

*Univariate and Panel Cointegration Tests*

The results of the three univariate unit root tests show that the price level data for small-cap and large-cap indices of the G7 countries are all non-stationary at 5% significance with the exception of Germany when considering each test output for it's statistical power (Table 22). The US small-cap Index, and the Canada Large-cap Index test statistic reject the null hypothesis of a unit root at 10% significance with the ADF test, but do not reject the null with the Modified ADF test and the ADF test with structural breaks. The ADF test with structural breaks has a higher power than the ADF test for both country large-cap and small-cap indices. At 10% significance, the Germany Small-cap Index rejects the null hypothesis with an R-squared value of

0.99 but does not reject the null hypothesis for the ADF or the modified-ADF test which have R-squared values of 0.02 and 0.01 respectively.

The unit root tests on the first differenced data are less congruent (Table 23). The ADF test and the ADF test with breaks show the resulting test statistic for each first differenced price series for the US reject the null hypothesis, however the modified-ADF test does not reject the null hypothesis for the large-cap and goods prices. The R-squared values for the ADF test and modified-ADF test are similar for these variables meaning neither test dominates the other. For Canada and Italy first differenced goods prices there is a similar outcome; the modified-ADF test does not reject the null hypothesis but the ADF test and ADF test with breaks do, while the ADF and modified-ADF test have similar R-squared values meaning neither dominates. The France and Germany first differenced goods prices test statistics for both the ADF and modified-ADF test fail to reject the null hypothesis, and while the ADF test with breaks does reject the null hypothesis, it has a much lower R-squared value than the other tests. For the UK first differenced goods prices, the ADF test and modified-ADF test do not reject the null hypothesis, and each have R-squared values of 0.75, while the ADF test with breaks does reject the null hypothesis but with an R-squared of only 0.02. The United Kingdom first differenced large-cap stock prices test statistic rejects the null hypothesis with the ADF test and ADF test with breaks but not with the modified-ADF test, the ADF test and modified-ADF test have similar R-squared values and the result of rejection of the null hypothesis is accepted in this case. For the UK first differenced goods prices, the ADF test and modified-ADF test are in agreement for not rejecting the null hypothesis of the presence of a unit root with a much higher R-squared value than the ADF test with breaks. For Japan first differenced goods prices, the test statistic estimated by the ADF test does not reject the null hypothesis while the modified-ADF test does with the same R-squared value of 0.54, while the ADF test with breaks rejects the null hypothesis but with a much lower R-squared value of 0.08. From these results we can conclude that goods prices for France, Germany, and the UK are not likely to be first order stationary I(1) due to not being able to reject the H0 with the ADF or modified-ADF test and the low R-squared values for the ADF test with breaks. For the United States, Canada, and Japan goods prices test statistics, as well as the United States and the United Kingdom Large-cap Index and we accept the results of being weakly first order stationary I(1), due to the ADF and modified-ADF test not being in agreement.

**Table 22 Univariate Unit Root Tests performed on price levels for each G7 country goods prices, large cap stock prices, and small cap stock prices**

*Table containing the results from the ADF test, modified ADF or ADF-GLS test, and the ADF test with breakpoints. T-statistics, p-values and R-squared estimates are shown. The null hypothesis for each test is the presence of a unit root. Critical values for the ADF test are -3.9851 for 1%, -3.4230 for 5%, -3.1344 for 10%, critical values for modified ADF test are -3.4742 for 1%, -2.9016 for 5%, -2.5903 for 10%, critical values for the ADF test with breaks is -5.3476 for 1%, -4.8598 for 5%, and -4.60732 for 10%.*

| | | ADF | | | Modified ADF (DF-GLS) | | | ADF w/ breaks | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | LC | SC | CPI | LC | SC | CPI | LC | SC | CPI |
| **USA** | **T-Statistic** | -1.7614 | -3.1817 | -1.8943 | -1.5354 | -3.0332 | -1.4912 | -3.2111 | -4.1551 | -3.4239 |
| | **p-value** | 0.7214 | 0.0898 | 0.6553 | 0.1256 | 0.0026 | 0.1369 | 0.8409 | 0.2838 | 0.7367 |
| | **R^2** | 0.01 | 0.07 | 0.3 | 0.01 | 0.07 | 0.29 | 0.99 | 0.99 | 1 |
| **Canada** | **T-Statistic** | -3.1558 | -2.5023 | -2.3605 | -2.7791 | -2.0994 | -2.3459 | -3.8864 | -4.0208 | -2.8913 |
| | **p-value** | 0.0953 | 0.3269 | 0.3997 | 0.0058 | 0.0365 | 0.0196 | 0.4447 | 0.3592 | 0.9396 |
| | **R^2** | 0.04 | 0.02 | 0.07 | 0.04 | 0.01 | 0.07 | 0.99 | 0.93 | 0.99 |
| **France** | **T-Statistic** | -1.9808 | -2.3983 | -2.9858 | -1.8406 | -2.3172 | -3.2074 | -3.9082 | -3.8708 | -3.7179 |
| | **p-value** | 0.6093 | 0.3798 | 0.1378 | 0.0665 | 0.0211 | 0.0015 | 0.4304 | 0.455 | 0.5535 |
| | **R^2** | 0.01 | 0.02 | 0.4 | 0.01 | 0.02 | 0.4 | 0.97 | 0.99 | 0.99 |
| **Germany** | **T-Statistic** | -2.4374 | -2.3967 | -2.7395 | -2.0609 | -1.8572 | -3.3175 | -4.2664 | -4.6218 | -3.8797 |
| | **p-value** | 0.3595 | 0.3806 | 0.2214 | 0.0401 | 0.0641 | 0.001 | 0.2215 | 0.0958 | 0.4494 |
| | **R^2** | 0.02 | 0.02 | 0.37 | 0.01 | 0.01 | 0.37 | 0.99 | 0.99 | 0.99 |
| **Italy** | **T-Statistic** | -2.2528 | -2.0044 | -3.0916 | -1.2706 | -1.7546 | -1.0869 | -4.1723 | -4.1662 | -3.5832 |
| | **p-value** | 0.4582 | 0.5964 | 0.11 | 0.2047 | 0.0802 | 0.2778 | 0.2739 | 0.2774 | 0.6403 |
| | **R^2** | 0.02 | 0.01 | 0.31 | 0 | 0.01 | 0.18 | 0.97 | 0.99 | 0.99 |
| **UK** | **T-Statistic** | -2.4134 | -2.5218 | -1.6289 | -2.102 | -2.4749 | -1.8874 | -4.1436 | -4.1882 | -2.9352 |
| | **p-value** | 0.3719 | 0.3174 | 0.7797 | 0.0363 | 0.0138 | 0.06 | 0.2903 | 0.2655 | 0.93 |
| | **R^2** | 0.02 | 0.02 | 0.53 | 0.01 | 0.02 | 0.52 | 0.97 | 0.99 | 0.99 |
| **Japan** | **T-Statistic** | -1.5372 | -2.1275 | -1.4007 | -1.0007 | -1.1303 | -1.6835 | -3.2038 | -3.5193 | -2.906 |
| | **p-value** | 0.815 | 0.5281 | 0.8593 | 0.3177 | 0.2592 | 0.0933 | 0.8437 | 0.6813 | 0.9364 |
| | **R^2** | 0.01 | 0.04 | 0.26 | 0 | 0.02 | 0.26 | 0.97 | 0.99 | 0.98 |

**Table 23 Univariate Unit Root Tests performed on first differences of each G7 country goods prices, large cap stock prices, and small cap stock prices**

*Table containing the results from the ADF test, modified ADF or ADF-GLS test, and the ADF test with breakpoints. T-statistics, p-values and R-squared estimates are shown. The null hypothesis for each test is the presence of a unit root. Critical values for the ADF test are -3.4493 for 1%, -2.8978 for 5%, -2.5712 for 10%, critical values for modified ADF test are -2.5718 for 1%, -1.94177 for 5%, -1.6161 for 10%, critical values for the ADF test with breaks is -4.9491 for 1%, -4.4437 for 5%, and -4.1936 for 10%.*

| | | ADF | | | Modified ADF (DF-GLS) | | | ADF w/ breaks | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | LC | SC | CPI | LC | SC | CPI | LC | SC | CPI |
| USA | T-Statistic | -18.0221 | -15.0178 | -11.2974 | -2.8462 | -14.4763 | -0.506 | -18.5998 | -15.0059 | -11.3215 |
| | p-value | 0 | 0 | 0 | 0.0047 | 0 | 0.6132 | <0.01 | <0.01 | <0.01 |
| | R^2 | 0.49 | 0.4 | 0.28 | 0.46 | 0.38 | 0.32 | 0.05 | 0.04 | 0.29 |
| Canada | T-Statistic | -16.2552 | -17.1365 | -14.6495 | -6.1171 | -16.8631 | 0.1387 | -17.1 | -18.1204 | -14.5606 |
| | p-value | 0 | 0 | 0 | 0 | 0 | 0.8898 | <0.01 | <0.01 | <0.01 |
| | R^2 | 0.44 | 0.46 | 0.39 | 0.35 | 0.45 | 0.42 | 0.08 | 0.06 | 0.06 |
| France | T-Statistic | -16.6851 | -17.3496 | -1.5716 | -3.9271 | -7.2672 | -1.0858 | -17.6668 | -17.9128 | -18.3928 |
| | p-value | 0 | 0 | 0.496 | 0.0001 | 0 | 0.2784 | <0.01 | <0.01 | <0.01 |
| | R^2 | 0.45 | 0.47 | 0.7 | 0.41 | 0.41 | 0.69 | 0.09 | 0.04 | 0 |
| Germany | T-Statistic | -17.9742 | -17.7199 | -0.4639 | -17.4021 | -17.2377 | -0.8745 | -18.6056 | -17.6832 | -10.7673 |
| | p-value | 0 | 0 | 0.8948 | 0 | 0 | 0.3825 | <0.01 | <0.01 | <0.01 |
| | R^2 | 0.49 | 0.48 | 0.69 | 0.47 | 0.47 | 0.69 | 0.72 | 0 | 0.05 |
| Italy | T-Statistic | -18.7624 | -17.1776 | -5.9774 | -6.0319 | -16.4546 | -0.856 | -19.4361 | -17.9206 | -7.9509 |
| | p-value | 0 | 0 | 0 | 0 | 0 | 0.3927 | <0.01 | <0.01 | <0.01 |
| | R^2 | 0.51 | 0.46 | 0.44 | 0.48 | 0.44 | 0.51 | 0.05 | 0.07 | 0.16 |
| UK | T-Statistic | -18.3262 | -16.5653 | -0.576 | -2.5726 | -5.0047 | -0.9694 | -18.607 | -16.9219 | -18.0464 |
| | p-value | 0 | 0 | 0.8725 | 0.0105 | 0 | 0.3331 | <0.01 | <0.01 | <0.01 |
| | R^2 | 0.5 | 0.45 | 0.75 | 0.43 | 0.39 | 0.75 | 0.02 | 0.04 | 0.02 |
| Japan | T-Statistic | -16.885 | -15.9824 | -3.3023 | -16.3727 | -15.8977 | -3.2932 | -18.1199 | -16.7842 | -14.387 |
| | p-value | 0 | 0 | 0.0156 | 0 | 0 | 0.0011 | <0.01 | <0.01 | <0.01 |
| | R^2 | 0.46 | 0.43 | 0.54 | 0.44 | 0.43 | 0.54 | 0.08 | 0.07 | 0.08 |

*Panel Unit Root Tests*

The panel groups being tested will be, the full sample of G7 countries, North American (US and Canada), and Ex-North American countries (UK, Germany, France, Japan, Italy). Furthermore, countries are also grouped according to the results of the univariate stationarity tests; countries which are first order stationary for goods prices are Canada, Italy, Japan, USA. The countries which are not first order stationary I(1) according to the univariate tests are and Italy, Germany, and the UK. As the literature shows the GFC marks a structural break for many developed economies, each series group is subdivided into date ranges before and after the GFC, 1994M01-2009M06 and 2009M07-2022M08.

For the Levin Lin and Chu (2002) panel test for the full date range 1994M01-2022M08, panel test-statistic results shows that each series group is first order stationary I(1), with the exception of the group of univariate not I(1) stationary large-cap prices which remains non first order stationary in 1$^{st}$ generation panel testing. Results are shown in Table 25.

Results of the panel unit root tests are summarized in Table 24. Results for the Pesaran (2007) CIPF panel unit root test for the full date range show that the large-cap prices for the univariate I(1) stationary group and Ex-North America group, and North America group goods prices and small-cap prices are not first order stationary I(1), but the remainder of the groups are first order stationary I(1) for both deterministic constant and deterministic constant and trend assumptions. The Pesaran (2007) test results are shown in Table 27.

Results of the Levin Lin and Chu (2002) test for the pre-GFC subsample show that all panel groups of goods price series are not first order I(1) stationary. In the post-GFC period, all panel groups with exception of the univariate I(1) stationary groups are not I(1) stationary for goods prices, as well as for the large-cap prices for the univariate I(1) stationary panel group and the Ex-North America group. These results are found in Table 26.

The results for the Pesaran (2007) test on the pre-GFC subsample shows that the good prices and large-cap prices are not I(1) stationary for the univariate I(1) stationary group as well as the large-cap prices for the ex-North America group, and the small-cap prices for the North America group.

In the post-GFC sub-sample, the panels shown to be non-stationary I(1) are the large-cap prices for the univariate I(1) stationary group, goods prices for the univariate not I(1) stationary group, the large-cap and small-cap prices for the ex-North America groups . The results are shown in Table 28.

**Table 24 Summary of results of panel unit root tests from Levin Lin and Chu (2002) and Pesaran (2007) for all panel groups, "x" denotes not I(1) stationary**

This table shows a summary of which panel groups are not first order I(1) stationary, indicated by an "x" , CPI represents goods prices panels, LC represents large cap price panels, and SC represents small cap price panels. Llc refers to the Levin Lin and Chu (2002) test, I(1) represents the panel group of univariate first order stationary countries I(1), not I(1) is for the countries which were not first order stationary in the univariate testing, ex-NA is the group of countries, the UK, Germany, Italy, France, and Japan, and NA represents the group of Canada and the US, full is the group containing all countries in thje G7

| | 1994-2022 | | | 1994-2009 | | | 2009-2022 | | |
|---|---|---|---|---|---|---|---|---|---|
| Levin Lin and Chu (2002) | | | | | | | | | |
| **llc** | **CPI** | **LC** | **SC** | **CPI** | **LC** | **SC** | **CPI** | **LC** | **SC** |
| full | | | | x | | | x | | |
| I(1) | | | | x | | | | x | |
| not I(1) | | x | | x | | | x | | |
| ex-NA | | | | x | | | x | x | |
| NA | | | | x | | | x | | |
| *Pesaran (2007)* | | | | | | | | | |
| | **CPI** | **LC** | **SC** | **CPI** | **LC** | **SC** | **CPI** | **LC** | **SC** |
| full | | | | | | | | | |
| I(1) | | x | | x | x | | | x | |
| not I(1) | | | | | | | x | | |
| ex-NA | | x | | | x | | | x | x |
| NA | x | | x | | | x | | | |

**Table 25 – Levin Lin and Chu (2002) panel unit root test with AIC lags; full sample 1994m01-2022m08**

Results output from the Levin Lin and Chu (2002) panel unit root test, with AIC lags determined by the VAR model. T-statistic is the test statistic for the null hypothesis of the presence of a unit root. CPI represents goods prices panels, Large cap is for the panels of large cap stock prices, and Small Cap is the panel for small cap stock prices. Test outputs for the level and first differences of each panel.

| | | CPI | | Large cap | | Small cap | |
|---|---|---|---|---|---|---|---|
| | | **level** | **1st diff** | **level** | **1st diff** | **level** | **1st diff** |
| **Full Sample** | **T-Statistic** | -1.7515 | -9.1495 | -0.0542 | -58.3236 | 0.4922 | -53.6150 |
| | **p-value** | 0.0399 | 0.0000 | 0.4784 | 0.0000 | 0.6887 | 0.0000 |
| **I(1) stationary** | **T-Statistic** | 0.2500 | -27.0222 | -3.0389 | -34.2895 | -0.1681 | -34.2895 |
| | **p-value** | 0.5987 | 0.0000 | 0.0012 | 0.0000 | 0.4333 | 0.0000 |
| **not I (1) stationary** | **T-Statistic** | -0.1020 | 20.2329 | 0.2109 | -39.3454 | 1.0899 | -34.4648 |
| | **p-value** | 0.4594 | 1.0000 | 0.5835 | 0.0000 | 0.8621 | 0.0000 |
| **Ex- North America** | **T-Statistic** | -0.1084 | -9.1163 | -2.7048 | -42.8997 | 0.5061 | -45.0951 |
| | **p-value** | 0.4569 | 0.0000 | 0.0034 | 0.0000 | 0.6936 | 0.0000 |
| **North America** | **T-Statistic** | 0.4152 | -18.6592 | -0.0244 | -29.2645 | 0.0898 | -29.0267 |
| | **p-value** | 0.6610 | 0.0000 | 0.4903 | 0.0000 | 0.5358 | 0.0000 |

**Table 26 Levin Lin and Chu (2002) panel unit root test with AIC lags; pre and post-GFC for all panel groups**

Results output from the Levin Lin and Chu (2002) panel unit root test, with AIC lags determined by the VAR model. T-statistic is the test statistic for the null hypothesis of the presence of a unit root. CPI represents goods prices panels, Large cap is for the panels of large cap stock prices, and Small Cap is the panel for small cap stock prices. Test outputs for the level and first differences of each panel.

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| *Subsample date range 1994M01 - 2009M06* | | | | | | | |
| | | **CPI** | | **Large cap** | | **Small cap** | |
| | | **level** | **1st diff** | **level** | **1st diff** | **level** | **1st diff** |
| **Full Sample** | **T-Statistic** | 0.3712 | 10.1400 | 1.2333 | -28.2286 | 2.1285 | -23.3300 |
| | **p-value** | 0.6448 | 1.0000 | 0.8913 | 0.0000 | 0.9834 | 0.0000 |
| **I(1) stationary** | **T-Statistic** | 1.7332 | 9.2683 | 0.9701 | -19.1427 | 1.9091 | -14.7028 |
| | **p-value** | 0.9585 | 1.0000 | 0.8340 | 0.0000 | 0.9719 | 0.0000 |
| **not I (1) stationary** | **T-Statistic** | 0.6670 | 10.9725 | 1.5996 | -19.1519 | 1.0572 | -18.0932 |
| | **p-value** | 0.7476 | 1.0000 | 0.9452 | 0.0000 | 0.8540 | 0.0000 |
| **Ex- North America** | **T-Statistic** | -0.1100 | 13.8956 | 1.8191 | -19.7105 | 1.2671 | -22.5356 |
| | **p-value** | 0.4562 | 1.0000 | 0.9655 | 0.0000 | 0.8974 | 0.0000 |
| **North America** | **T-Statistic** | 4.9145 | 5.3466 | 0.5916 | -18.7456 | 2.5044 | -6.4521 |
| | **p-value** | 1.0000 | 1.0000 | 0.7229 | 0.0000 | 0.9939 | 0.0000 |
| *Subsample date range 2009M07 - 2022M08* | | | | | | | |
| | | **CPI** | | **Large cap** | | **Small cap** | |
| | | **level** | **1st diff** | **level** | **1st diff** | **level** | **1st diff** |
| **Full Sample** | **T-Statistic** | 4.7793 | 15.2942 | -1.3196 | -36.5240 | 0.2982 | -35.8781 |
| | **p-value** | 1.0000 | 1.0000 | 0.0935 | 0.0000 | 0.6172 | 0.0000 |
| **I(1) stationary** | **T-Statistic** | 4.6545 | -3.5907 | -4.9476 | -25.4433 | -0.1586 | -25.7100 |
| | **p-value** | 1.0000 | 0.0000 | 0.0000 | 0.0000 | 0.4370 | 0.0000 |
| **not I (1) stationary** | **T-Statistic** | 3.1077 | 16.1735 | -1.2052 | -21.2331 | 0.6501 | -25.3017 |
| | **p-value** | 0.9990 | 1.0000 | 0.1141 | 0.0000 | 0.7422 | 0.0000 |
| **Ex- North America** | **T-Statistic** | 2.1356 | 2.7129 | -4.6045 | -24.0253 | 0.8274 | -32.1037 |
| | **p-value** | 0.9836 | 0.9967 | 0.0000 | 0.0000 | 0.7960 | 0.0000 |
| **North America** | **T-Statistic** | 7.5034 | 4.7610 | -0.9448 | -23.2245 | -0.9485 | -16.1046 |
| | **p-value** | 1.0000 | 1.0000 | 0.1724 | 0.0000 | 0.1714 | 0.0000 |

**Table 27 – Pesaran (2007) panel unit root test for each panel group, for price level series, and for first difference series**

The table shows the output from the Pesaran (2007) panel root tests for price level and first differenced series, under the assumption of a determinist constant only, and deterministic constant and linear trend. The null hypothesis is the presence of a unit root. The critical values for the price level series test statistics are -3.03 for 1%, -2.83 for 5%, -2.73 for 10%. For the first differenced (FD) series it is -2.53 for 1%, -2.32 for 5%, -2.21 for 10%.CPI represents goods prices panels, Large cap is for the panels of large cap stock prices, and Small Cap is the panel for small cap stock prices. Test outputs for the level and first differences of each panel.

*Assuming deterministic constant only.*

| | | **CPI** | | **Large cap** | | **Small cap** | |
|---|---|---|---|---|---|---|---|
| | **Statistic** | **Level** | **1st diff** | **Level** | **1st diff** | **Level** | **1st diff** |
| **Full Sample** | **t-stat** | -1.43758 | -8.31027 | -0.70799 | -12.8047 | -1.04002 | -15.447 |
| | **p-value** | >=0.10 | <0.01 | >=0.10 | <0.01 | >=0.10 | <0.01 |
| **I(1) stationary** | **t-stat** | 0.02539 | -13.0115 | -5.66151 | -18.3264 | -0.61062 | -13.353 |
| | **p-value** | >=0.10 | <0.01 | <0.01 | <0.01 | >=0.10 | <0.01 |
| **not I (1) stationary** | **t-stat** | -1.70119 | -13.7043 | -0.98644 | -13.9626 | -1.69 | -18.7332 |
| | **p-value** | >=0.10 | <0.01 | >=0.10 | <0.01 | >=0.10 | <0.01 |
| **Ex North America** | **t-stat** | 0.34767 | -14.4631 | -4.59926 | -15.6313 | -1.81 | -13.4613 |
| | **p-value** | >=0.10 | <0.01 | <0.01 | <0.01 | >=0.10 | <0.01 |
| **North America** | **t-stat** | -2.87969 | <0.01 | -0.61893 | >=0.10 | 0.33 | >=0.10 |
| | **p-value** | -24.834 | <0.01 | -18.5589 | <0.01 | -16.2467 | <0.01 |

*Assuming deterministic constant and linear trend.*

| | | **CPI** | | **Large cap** | | **Small cap** | |
|---|---|---|---|---|---|---|---|
| | **Statistic** | **Level** | **1st diff** | **Level** | **1st diff** | **Level** | **1st diff** |
| **Full Sample** | **t-stat** | -1.32082 | -12.6445 | -1.74169 | -12.6445 | -2.03 | -15.503 |
| | **p-value** | >=0.10 | <0.01 | >=0.10 | <0.01 | >=0.10 | <0.01 |
| **I(1) stationary** | **t-stat** | -1.29436 | -13.0987 | -5.31403 | -18.3962 | -2.35 | -13.4415 |
| | **p-value** | >=.10 | <0.01 | <0.01 | <0.01 | >=0.10 | <0.01 |
| **not I (1) stationary** | **t-stat** | -2.19959 | -13.8356 | -2.29079 | -13.9588 | -2.76 | -18.7554 |
| | **p-value** | >=0.10 | <0.01 | >=0.10 | <0.01 | <0.10 | <0.01 |
| **Ex North America** | **t-stat** | -0.70517 | -14.6019 | -4.7812 | -15.7272 | -2.19941 | -13.5867 |
| | **p-value** | >=0.10 | <0.01 | <0.01 | <0.01 | >=0.10 | <0.01 |
| **North America** | **t-stat** | -3.06214 | -24.7848 | -0.06439 | -18.5459 | -4.01 | -16.2268 |
| | **p-value** | <0.01 | <0.01 | >=0.10 | <0.01 | <0.01 | <0.01 |

**Table 28 – Pesaran (2007) panel unit root test for each panel group, for price level series, and for first difference series for pre and post-GFC subsamples**

The table shows the output from the Pesaran (2007) panel root tests for price level and first differenced series,  under the assumption of a determinist constant only, and deterministic constant and linear trend. The null hypothesis is the presence of a unit root. The critical values for the price level series test statistics are -3.03 for 1%, -2.83 for 5%, -2.73 for 10%. For the first differenced (FD) series it is -2.53 for 1%, -2.32 for 5%, -2.21 for 10%.CPI represents goods prices panels, Large cap is for the panels of large cap stock prices, and Small Cap is the panel for small cap stock prices. Test outputs for the level and first differences of each panel.

*Assuming deterministic constant and linear trend 1994M01 - 2009M06*

| | | **CPI** | | **Large cap** | | **Small cap** | |
|---|---|---|---|---|---|---|---|
| | **Statistic** | **Level** | **1st diff** | **Level** | **1st diff** | **Level** | **1st diff** |
| **Full Sample** | **t-stat** | -2.4331 | -4.8659 | -2.1641 | -10.792 | -1.822 | -12.182 |
| | **p-value** | >=0.10 | <0.01 | >0.10 | <0.01 | >=0.10 | <0.01 |
| **I(1) stationary** | **t-stat** | -3.0349 | -10.153 | -4.459 | -13.133 | -2.1994 | -10.347 |
| | **p-value** | <0.01 | <0.01 | <0.01 | <0.01 | >=0.10 | <0.01 |
| **not I (1) stationary** | **t-stat** | -2.336 | -10.343 | -1.5259 | -11.141 | -2.4212 | -16.186 |
| | **p-value** | <0.05 | <0.01 | >=0.10 | <0.01 | >=0.10 | <0.01 |
| **Ex North America** | **t-stat** | -2.1203 | -11.644 | -4.0851 | -11.351 | -1.9447 | -10.802 |
| | **p-value** | >=0.10 | <0.01 | <0.01 | <0.01 | >=0.10 | <0.01 |
| **North America** | **t-stat** | -2.42 | -19.904 | -1.7025 | -13.406 | -3.7982 | -12.28 |
| | **p-value** | >=0.10 | <0.01 | >=0.10 | <0.01 | <0.01 | <0.01 |

*Assuming deterministic constant and linear trend 2009M07-2022M08*

| | | **CPI** | | **Large cap** | | **Small cap** | |
|---|---|---|---|---|---|---|---|
| | **Statistic** | **Level** | **1st diff** | **Level** | **1st diff** | **Level** | **1st diff** |
| **Full Sample** | **t-stat** | -2.2211 | -6.5568 | -2.3831 | -11.636 | -1.879 | -11.552 |
| | **p-value** | >=0.10 | <0.01 | >=0.10 | <0.01 | >=0.10 | <0.01 |
| **I(1) stationary** | **t-stat** | -0.4595 | -8.5913 | -5.1659 | -13.649 | -2.143 | -12.672 |
| | **p-value** | >=0.10 | <0.01 | <0.01 | <0.01 | >=0.10 | <0.01 |
| **not I (1) stationary** | **t-stat** | -3.326 | -12.572 | -1.609 | -19.877 | -1.3773 | -16.653 |
| | **p-value** | <0.01 | <0.01 | >=0.10 | <0.01 | >=0.10 | <0.01 |
| **Ex North America** | **t-stat** | -0.591 | -6.9113 | -4.3796 | -12.281 | -2.7834 | -8.3985 |
| | **p-value** | >0.10 | <0.01 | <0.01 | <0.01 | <0.10 | <0.01 |
| **North America** | **t-stat** | -1.9453 | -13.014 | -1.1209 | -13.259 | -1.6113 | -11.445 |
| | **p-value** | >=0.10 | <0.01 | >=0.10 | <0.01 | >=0.10 | <0.01 |

*Johansen Cointegration Test and Pedroni Panel Cointegration Test*

The univariate time series and panels that are shown to be first order stationary I(1) meet the criteria to be tested for a cointegrating relationship. The results of the Johansen Cointegration test, with the assumption of an intercept and linear deterministic trend show that only the Japan Large-cap Index has a significant cointegrating vector between the stock and goods prices indices. When running the Johansen test with the assumption of an intercept but no linear determinist trend assumption, we find that the United States and Canada small-cap and large-cap prices each had 1 cointegrating relationship with goods prices, with the Canada large-cap prices having 2 cointegrating relationships, and the Italy large-cap prices also had 1 cointegrating relationship with goods prices. This can be interpreted as there being the presence of a long-run relationship when there is a deterministic intercept and no deterministic linear relationship for these country stock indices. Results are summarized in Table 29.

The results of the Pedroni panel cointegration Test for the full date range shows that there are only cointegrating relationships amongst the sample groups is the small-cap price panels for the North America group post-GFC. With the Kao method, the null hypothesis of no cointegrating relationships is rejected for the full panels post-GFC for both the large-cap and small-cap prices, for the univariate non I(1) stationary group small-cap prices panel from the full date range, and the large-cap prices panel from the pre-GFC sub-sample are found to reject the H0. In the Ex-North America group only the small-cap prices panel for the pre-GFC period is found to reject H0. For the North America panel, only the small-cap prices panel in the post-GFC sub-sample rejects the H0 of no cointegration. Results are summarized in Table 30 and Table 31, Table 32, Table 33, Table 34, and Table 35.

**Table 29 Trace statistic results for Johansen Cointegration test for first order stationary series I(1)**

*The table shows the results from the Johansen Cointegraiton test where the H0 is no cointegration. Lags from AIC from the VAR model, under two assumption scenarios; intercept and linear deterministic trend, and intercept but no linear deterministic trend.*

*Assumption of intercept and linear deterministic trend*

| Hypothesised Number of Cointegrating Equations at 5% significance | US Small cap Index | US Large cap Index | Canada Small cap Index | Canada Large cap Index | Japan Small cap Index | Japan Large cap Index | Italy Small cap Index | Italy Large cap Index | Critical value at 5% significance |
|---|---|---|---|---|---|---|---|---|---|
| 0 | 12.2287 (0.1463) | 6.8748 (0.5923) | 9.586 (0.3139) | 11.6188 (0.1762) | 13.773 (0.0894) | **18.0717 (0.0200)** | 7.5630 (0.5132) | 8.0863 (0.4565) | **15.4947** |
| At most 1 | 0.0228 (0.8789) | 0.1785 (0.6726) | 0.4574 (0.4988) | 0.1375 (0.7107) | 1.0355 (0.3089) | **3.841 (0.3660)** | 2.3451 (-0.1257) | 0.6828 (0.4086) | **3.8415** |
| Lags | 3 | 3 | 2 | 2 | 5 | 5 | 4 | 4 | |

*Assumption of intercept, no linear deterministic trend*

| Hypothesised Number of Cointegrating Equations at 5% significance | United States Small cap Index | United States Large cap Index | Canada Small cap Index | Canada Large cap Index | Japan Small cap Index | Japan Large cap Index | Italy Small cap Index | Italy Large cap Index | Critical value at 5% significance |
|---|---|---|---|---|---|---|---|---|---|
| 0 | **45.6255 (0.000)** | **41.6539 (0.000)** | **51.4818 (0.000)** | **48.3654 (0.000)** | 14.4514 (0.2595) | 18.1181 (0.0960) | 19.6974 (0.0596) | **25.5819 (0.0084)** | **20.2618** |
| At most 1 | 8.9978 (0.0538) | 4.3653 (0.3605) | 8.7752 (0.0592) | **10.8581 (0.0236)** | 0.69104 (0.9834) | 0.83412 (0.9706) | 5.3479 (0.2474) | 5.2032 (0.2619) | **9.1645** |
| Lags | 3 | 3 | 2 | 2 | 5 | 5 | 4 | 4 | |

**Table 30 Test statistics for the Pedroni Panel Cointegration test for all panel groups in date sample range 1994M01-2022M08**

P-value results for the the Pedroni Panel Cointegration outputs a test statistic for a v-statistic, rho-statistic, pp-statistic, adf-statistic, group rho-statistic, group pp-statistic, group-ADF statistic. The null hypothesis is no cointegration. Reject the null hypothesis at a p-value <0.05, majority of test statistics must be rejected for a cointegrating relationship to be confirmed.

| | | Full Sample | | | | I(1) stationary | | | | not I (1) stationary | | | | Ex North America | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. |
| Large cap | Panel v-Statistic | 7.4911 | 0.0000 | 5.9700 | 0.0000 | - | - | - | - | 6.6807 | 0.0000 | 6.6376 | 0.0000 | - | - | - | - |
| | Panel rho-Statistic | 1.1423 | 0.8733 | 0.7554 | 0.7750 | - | - | - | - | 1.8978 | 0.9711 | 1.5441 | 0.9387 | - | - | - | - |
| | Panel PP-Statistic | 1.5661 | 0.9413 | 1.3272 | 0.9078 | - | - | - | - | 3.0609 | 0.9989 | 2.5267 | 0.9942 | - | - | - | - |
| | Panel ADF-Statistic | 0.9831 | 0.8372 | 0.0948 | 0.5378 | - | - | - | - | 0.5429 | 0.7064 | 0.2885 | 0.6135 | - | - | - | - |
| | Group rho-Statistic | 1.6792 | 0.9534 | - | - | - | - | - | - | 2.0783 | 0.9812 | - | - | - | - | - | - |
| | Group PP-Statistic | 2.5346 | 0.9944 | - | - | - | - | - | - | 3.4703 | 0.9997 | - | - | - | - | - | - |
| | Group ADF-Statistic | 0.3457 | 0.6352 | - | - | - | - | - | - | 0.2950 | 0.6160 | - | - | - | - | - | - |
| Small cap | Panel v-Statistic | 0.6760 | 0.2495 | 0.8097 | 0.2091 | -0.8466 | 0.8014 | 1.2985 | 0.0971 | 7.1643 | 0.0000 | 7.3819 | 0.0000 | -0.9470 | 0.8282 | 1.8600 | 0.0314 |
| | Panel rho-Statistic | -0.9451 | 0.1723 | -0.9699 | 0.1661 | 0.4905 | 0.6881 | 0.2076 | 0.5822 | 1.5800 | 0.9429 | 1.0735 | 0.8585 | 0.6173 | 0.7315 | 0.6748 | 0.7501 |
| | Panel PP-Statistic | -1.0127 | 0.1556 | -1.0069 | 0.1570 | -0.3981 | 0.3453 | 0.1826 | 0.5724 | 2.6934 | 0.9965 | 2.0220 | 0.9784 | -0.3374 | 0.3679 | 0.8590 | 0.8048 |
| | Panel ADF-Statistic | -0.6551 | 0.2562 | -0.6510 | 0.2575 | -0.6827 | 0.2474 | -0.5315 | 0.2975 | 0.3702 | 0.6444 | -0.1648 | 0.4346 | -0.7534 | 0.2256 | -0.6201 | 0.2676 |
| | Group rho-Statistic | -0.1166 | 0.4536 | - | - | 0.5798 | 0.719 | - | - | 1.7438 | 0.9594 | - | - | 1.8808 | 0.9700 | | |
| | Group PP-Statistic | -0.4741 | 0.3177 | - | - | 0.7833 | 0.7833 | - | - | 3.0339 | 0.9988 | - | - | 2.6587 | 0.9961 | | |
| | Group ADF-Statistic | -0.3256 | 0.3724 | - | - | 0.0826 | 0.5329 | - | - | 0.0066 | 0.5026 | - | - | -0.0492 | 0.4804 | | |

**Table 31 Test statistics for the Pedroni Panel Cointegration test for all panel groups in date sample range 1994m01-2009M06**

P-value results for the the Pedroni Panel Cointegration outputs a test statistic for a v-statistic, rho-statistic, pp-statistic, adf-statistic, group rho-statistic, group pp-statistic, group-ADF statistic. The null hypothesis is no cointegration. Reject the null hypothesis at a p-value <0.05, majority of test statistics must be rejected for a cointegrating relationship to be confirmed.

| | | Full Sample | | | | not I (1) stationary | | | | Ex North America | | | | North America | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. |
| Large cap | Panel v-Statistic | 12.4968 | 0.0000 | 6.0642 | 0.0000 | 5.2533 | 0.0000 | 5.5696 | 0.0000 | - | - | - | - | 11.0744 | 0.0000 | 10.3498 | 0.0000 |
| | Panel rho-Statistic | -2.2854 | 0.0111 | -1.2261 | 0.1101 | 0.2889 | 0.6137 | 0.1164 | 0.5463 | - | - | - | - | -4.1445 | 0.0000 | -4.0392 | 0.0000 |
| | Panel PP-Statistic | -2.3631 | 0.0091 | -1.3745 | 0.0846 | 0.5459 | 0.7074 | 0.3494 | 0.6366 | - | - | - | - | -3.2282 | 0.0006 | -3.1662 | 0.0008 |
| | Panel ADF-Statistic | -0.5981 | 0.2749 | -0.2941 | 0.3843 | 0.0909 | 0.5362 | -0.1363 | 0.4458 | - | - | - | - | -0.3187 | 0.3750 | -0.0973 | 0.4612 |
| | Group rho-Statistic | -1.2159 | 0.1120 | | | 0.3371 | 0.6320 | | | - | - | - | - | -2.9891 | 0.0014 | | |
| | Group PP-Statistic | -1.7018 | 0.0444 | | | 0.5120 | 0.6957 | | | - | - | - | - | -3.0151 | 0.0013 | | |
| | Group ADF-Statistic | -0.7691 | 0.2209 | | | 0.1559 | 0.5620 | | | - | - | - | - | 0.0726 | 0.5289 | | |
| Small cap | Panel v-Statistic | 12.3375 | 0.0000 | 6.9607 | 0.0000 | 4.3724 | 0.0000 | 4.6389 | 0.0000 | -1.6560 | 0.9511 | 0.8926 | 0.1860 | - | - | - | - |
| | Panel rho-Statistic | -1.8199 | 0.0344 | -1.2201 | 0.1112 | 0.4587 | 0.6768 | 0.3335 | 0.6306 | 1.1186 | 0.8683 | 0.3083 | 0.6211 | - | - | - | - |
| | Panel PP-Statistic | -2.0808 | 0.0187 | -1.3459 | 0.0892 | 0.5805 | 0.7192 | 0.4357 | 0.6685 | 0.6210 | 0.7327 | 0.0673 | 0.5268 | - | - | - | - |
| | Panel ADF-Statistic | -1.0873 | 0.1384 | -0.3738 | 0.3543 | 0.2255 | 0.5892 | 0.0324 | 0.5129 | -0.6733 | 0.2504 | -0.1085 | 0.4568 | - | - | - | - |
| | Group rho-Statistic | -0.6781 | 0.2489 | | | 0.7640 | 0.7776 | | | 0.9642 | 0.8325 | | | - | - | - | - |
| | Group PP-Statistic | -1.3950 | 0.0815 | | | 0.7577 | 0.7757 | | | 0.7256 | 0.7660 | | | - | - | - | - |
| | Group ADF-Statistic | -0.8051 | 0.2104 | | | 0.3725 | 0.6452 | | | 0.5233 | 0.6996 | | | - | - | - | - |

**Table 32 Test statistics for the Pedroni Panel Cointegration test for all panel groups in date sample range 2009M07 - 2022M08**

P-value results for the the Pedroni Panel Cointegration outputs a test statistic for a v-statistic, rho-statistic, pp-statistic, adf-statistic, group rho-statistic, group pp-statistic, group-ADF statistic. The null hypothesis is no cointegration. Reject the null hypothesis at a p-value <0.05, majority of test statistics must be rejected for a cointegrating relationship to be confirmed.

| | | Full Sample | | | | I(1) stationary | | | | North America | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. | Statistic | Prob. | Weighted Statistic | Prob. |
| Large cap | Panel v-Statistic | 5.5906 | 0.0000 | 5.0462 | 0.0000 | - | - | - | - | 2.8556 | 0.0021 | 3.0797 | 0.0010 |
| | Panel rho-Statistic | 3.9707 | 1.0000 | 2.9470 | 0.9984 | - | - | - | - | 2.5324 | 0.9943 | 2.3954 | 0.9917 |
| | Panel PP-Statistic | 6.6530 | 1.0000 | 5.0145 | 1.0000 | - | - | - | - | 4.2680 | 1.0000 | 4.0275 | 1.0000 |
| | Panel ADF-Statistic | 1.2135 | 0.8875 | 0.0065 | 0.5026 | - | - | - | - | 0.4262 | 0.6650 | 0.3313 | 0.6298 |
| | Group rho-Statistic | 3.9027 | 1.0000 | | | - | - | - | - | 2.5273 | 0.9943 | | |
| | Group PP-Statistic | 6.9930 | 1.0000 | | | - | - | - | - | 4.7243 | 1.0000 | | |
| | Group ADF-Statistic | 1.4202 | 0.9222 | | | - | - | - | - | 0.5789 | 0.7187 | | |
| Small cap | Panel v-Statistic | 5.4137 | 0.0000 | 5.1234 | 0.0000 | 5.6832 | 0.0000 | 3.8929 | 0.0000 | 3.3377 | 0.0004 | 2.2987 | 0.0108 |
| | Panel rho-Statistic | 3.5955 | 0.9998 | 2.5684 | 0.9949 | -3.8229 | 0.0001 | 0.4788 | 0.6840 | -3.3171 | 0.0005 | -2.1690 | 0.0150 |
| | Panel PP-Statistic | 5.9820 | 1.0000 | 4.3688 | 1.0000 | -2.6680 | 0.0038 | 1.3842 | 0.9169 | -2.5366 | 0.0056 | -1.7366 | 0.0412 |
| | Panel ADF-Statistic | 0.3553 | 0.6388 | -0.3425 | 0.3660 | -3.3319 | 0.0004 | -1.5302 | 0.0630 | -2.5905 | 0.0048 | -2.7687 | 0.0028 |
| | Group rho-Statistic | 3.5180 | 0.9998 | | | 1.1045 | 0.8653 | | | -1.5101 | 0.0655 | | |
| | Group PP-Statistic | 6.2002 | 1.0000 | | | 2.9285 | 0.9983 | | | -1.5652 | 0.0588 | | |
| | Group ADF-Statistic | 0.8111 | 0.7913 | | | -0.0661 | 0.4736 | | | -2.5828 | 0.0049 | | |

**Table 33 Kao Engle-Granger Panel Cointegration test results for subsample date ranges 1994M01-2022M08, 1994M01-2009M06, 2009M07-2022M08**

The results show the test statistic and associated p-values for the Kao Engle-Granger Panel Cointegration test, the null hypothesis is of no cointegration in the residuals

| | | Full | p-value | I(1) stationary | p-value | not I(1) stationary | p-value | Ex-North America | p-value | North America | p-value |
|---|---|---|---|---|---|---|---|---|---|---|---|
| 1994M01-2022M08 | Large cap | 0.4475 | 0.3273 | - | - | 0.1028 | 0.4591 | - | - | - | - |
| | Small cap | 0.3535 | 0.3619 | -0.331 | 0.3704 | -2.551 | 0.0054 | -0.2903 | 0.3858 | - | - |
| 1994M01-2009M06 | Large cap | 0.3813 | 0.3515 | - | - | -2.3662 | 0.0009 | - | - | -0.5298 | 0.2981 |
| | Small cap | -1.262 | 0.1035 | - | - | -1.0834 | 0.1393 | -3.1098 | 0.0009 | -0.5743 | 0.2829 |
| 2009M07-2022M08 | Large cap | 2.2924 | 0.0109 | - | - | - | - | - | - | 0.4527 | 0.3254 |
| | Small cap | 2.7163 | 0.003 | -1.24 | 0.1077 | - | - | - | - | 2.2469 | 0.0123 |

*Long-run relationship between goods prices and stock prices*

The 14 cointegrating pairs of variables are now tested for a long-run relationship using the Vector Error Correction Model. Each model is checked for residual normality, serial correlation, and heteroskedasticity. We find that throughout the results, there is no model with a significant speed of adjustment statistic except for the Italy Small-cap Index (Table 36). The model for this variable shows weak endogeneity, and the speed of adjustment is near zero and positive, the model has a low R-squared of 0.05. This suggests that for all of the series tested, save for the exceptions mentioned, goods prices have long run relationship with stock prices at 5% significance.

The panel results from the FMOLS model are a stark difference from the individual series' models. We see that the beta coefficient is significant at 5% according to the corresponding t-statistic. The ranges of the Beta values are from 1.0797 to 1.2269. Given that all the coefficients are greater than unit, we can infer that each of these panel groups are a hedge to their respective inflation panels according to the Fisher elasticity. We see that the small-cap panel group of univariate non stationary I(1) variables, which is to have a positive and greater than unit beta coefficient over the full date range. The panel of these country small-cap indices, France, Germany, the United Kingdom, can be said to have a positive long-run relationship with the respective panel goods prices. The same group has a positive Fisher Elasticity for the pre-GFC period of 1.1656, which is lower than that for the full date range. The highest Fisher Elasticity is for the post-GFC full sample small-cap prices panel with 1.2264. Results are summarized in Table 37.

**Table 34 Vector Error Correction Model Outputs for stock prices variables with cointegrating relationships with goods prices**

Model specified with p-1 lags, assumptions according to the cointegration results, country stock prices as dependent variables and country goods prices for the independent variable. T-statistics are found in the square brackets, values less than threshold of 2 and with p-values less than 0.05 are significant.

| Dependent Variable | Coefficient | Speed of | (p-value) | R^2 |
|---|---|---|---|---|
| **Japan Large cap Index (intercept, linear trend)** | 12.6722 [-4.56] | -0.02024 | 0.1578 | 0.06 |
| **United States Small cap Index (intercept,** | 4.2231 [-6.66] | 0.00653 | 0.1642 | 0.04 |
| **United States Large cap Index (intercept,** | 2.4442 [-3.27] | 0.003852 | 0.1724 | 0 |
| **Canada Small cap Index (intercept, no** | -0.10803 [2.61] | 0.000199 | 0.9267 | 0.02 |
| **Canada Large cap Index (intercept, no** | -5.7041 [1.29] | 0.000325 | 0.8808 | 0.01 |
| **Japan Small cap (intercept, no trend)** | 13.7118 [-3.041] | 0.00261 | 0.7672 | 0.06 |
| **Japan Large cap (intercept, no trend)** | 12.6722[-4.57] | -0.0177 | 0.2046 | 0.05 |
| **Italy Large cap (Intercept, no trend)** | -4.1613 [1.91] | -0.004102 | 0.0698 | 0.03 |
| **Italy Small cap (Intercept, no trend)** | **-40.13456 [1.64]** | **0.00043** | **0.0224** | 0.05 |

**Table 35 Model estimates for Fully Modified Ordinary Least Squares (FMOLS) cointegration panel regression**

Model with no trend specification assumption, goods prices as the independent variable, stock prices as the dependent variable.

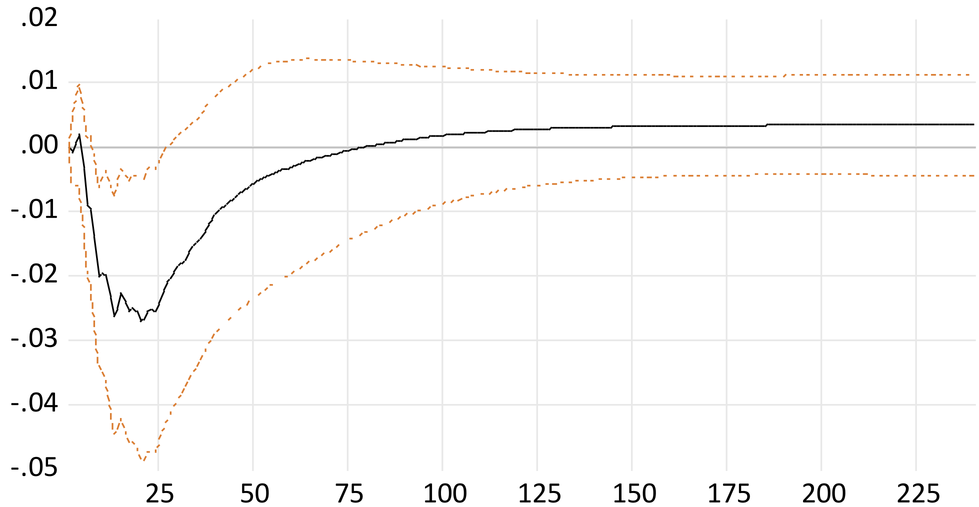
| Dependent Variable | Beta | t-statistic | p-value |
|---|---|---|---|
| **Panel Group Not I(1) stationary small cap 1994-2022** | 1.2269 | 139.5941 | 0 |
| **Panel Group Not I(1) stationary large cap 1994-2009** | 1.1656 | 182.1425 | 0 |
| **Panel Ex North America Group Small cap 1994-2009** | 1.0797 | 244.05 | 0 |
| **Panel Group Full Sample Large cap Index 2009-2022** | 1.1399 | 379.1311 | 0 |
| **Panel Group Full Sample Small cap Index 2009-2022** | 1.2264 | 225.0325 | 0 |
| **Panel North America Group Small cap 2009-2022** | 1.122 | 100.8309 | 0 |

*Short-Run Relationship*

The VAR model residuals are verified for normality, autocorrelation and heteroskedasticity, none were found to violate asymptotic assumptions. Each model was found to have high R-squared values ranging from 0.93-0.99 indicating a robust result from the estimation. The IRF are estimated using log level data, with lags of 12 and a horizon of 240. Each IRF follows a path that is consistent with the literature as initial shocks from goods prices results in an immediate drawdown in stock prices followed by a recovery, and stability. A few exceptions arise; Canada small-cap price series experiences an initial sharp drop in prices and see a trough in period 9 but never reaches full recovery to the initial price level but does within the bounds of one standard deviation which reaches par in period 22. The magnitude of the negative response to a unit shock for the small-cap price series is far deeper than for the large-cap price series. The Japan small-cap price series follows a similar pattern, hitting a trough in period 9 and never crossing par, however the first standard deviation does cross at period 10. The spread between the negative shocks to the Japan large-cap and small-cap price is less than that for Canada. The Italian large-cap stock price series mean response also does not reach par, but the first standard deviation crosses at period 39. In nearly all countries we see that the magnitude of negative shock to stock prices is greater for small-caps than for the respective large-cap prices. However, in more small-cap series than large-cap series, we see a higher-level period of stability. For the US, UK and France, the small-cap prices recover faster than the large-cap prices, reaching par in earlier periods despite having greater negative responses. Figures 18-31 show the graphs of IRF paths.

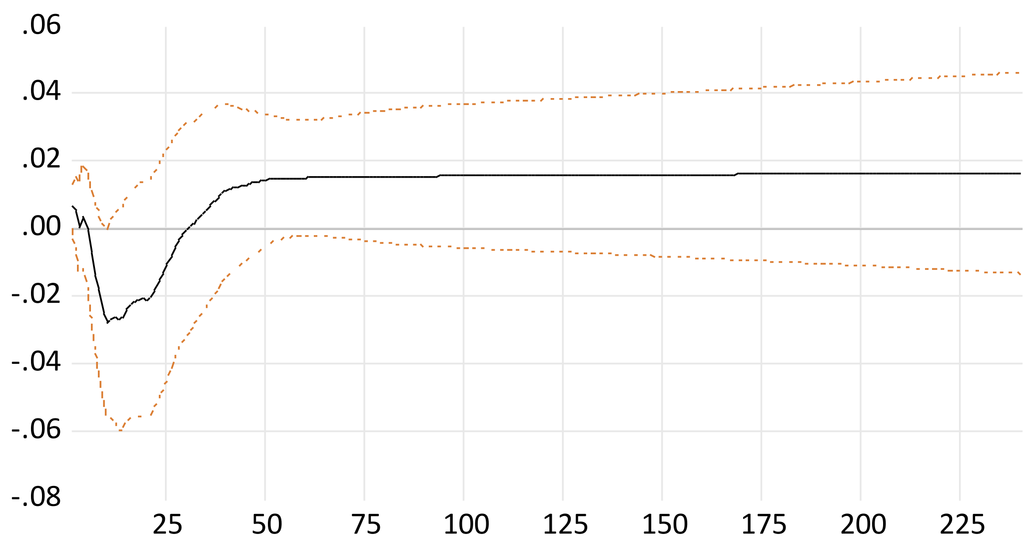
*Impulse Response Functions*

Figure 17 Response of US Large Cap Index to US CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



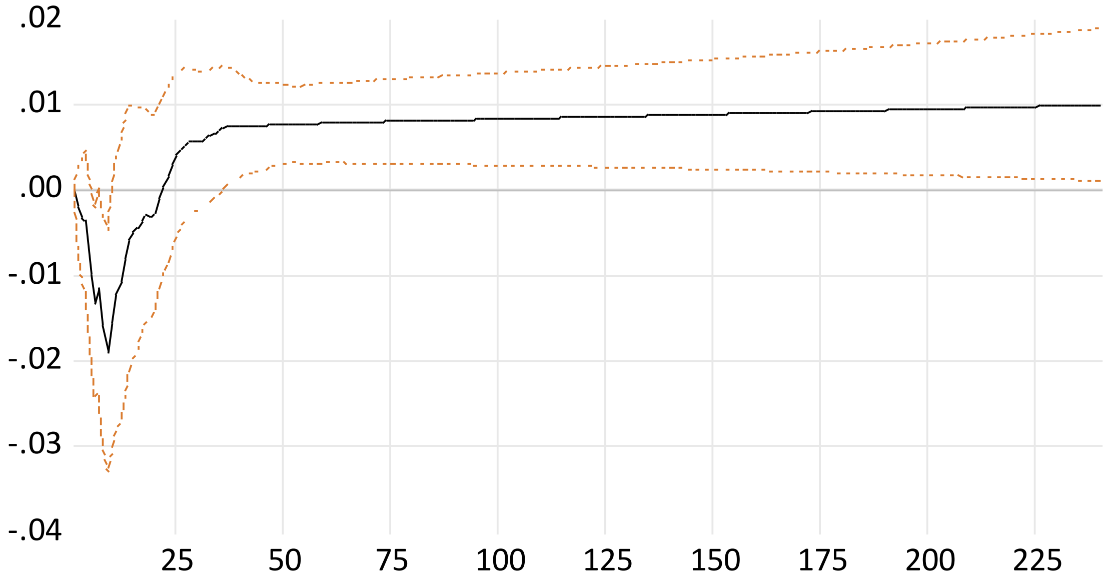

*In response to an inflation shock, stock prices reach a valley of -0.027 in period 21 and reaches par in period 77.*

Figure 18 Response of US Small Cap Index to US CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



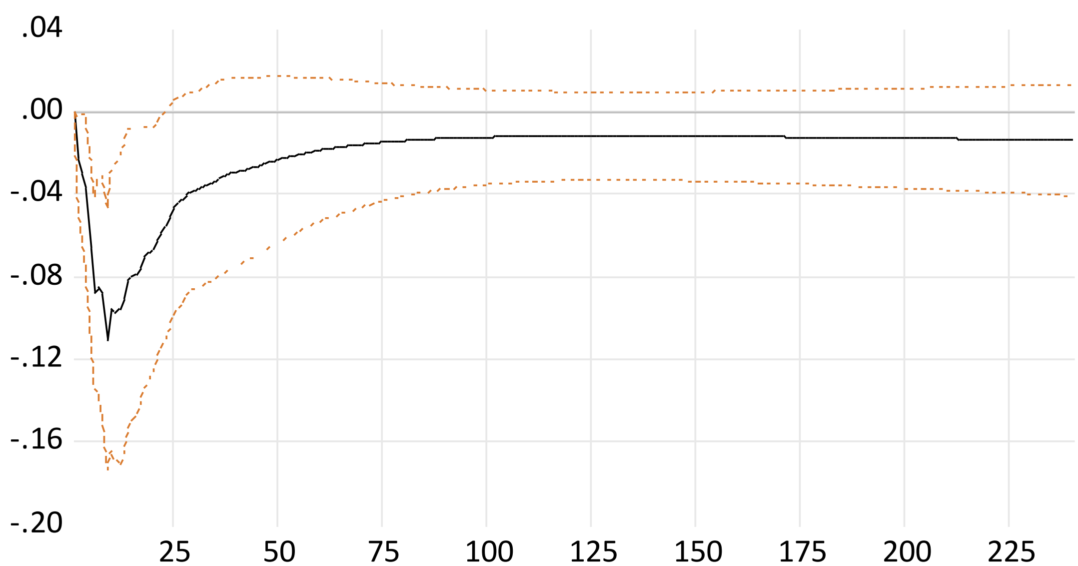

*In response to an inflation shock, stock prices reach a valley of -0.027 in period 10 and reaches par in period 30.*

Figure 19 Response of Canada Large Cap to Canada CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



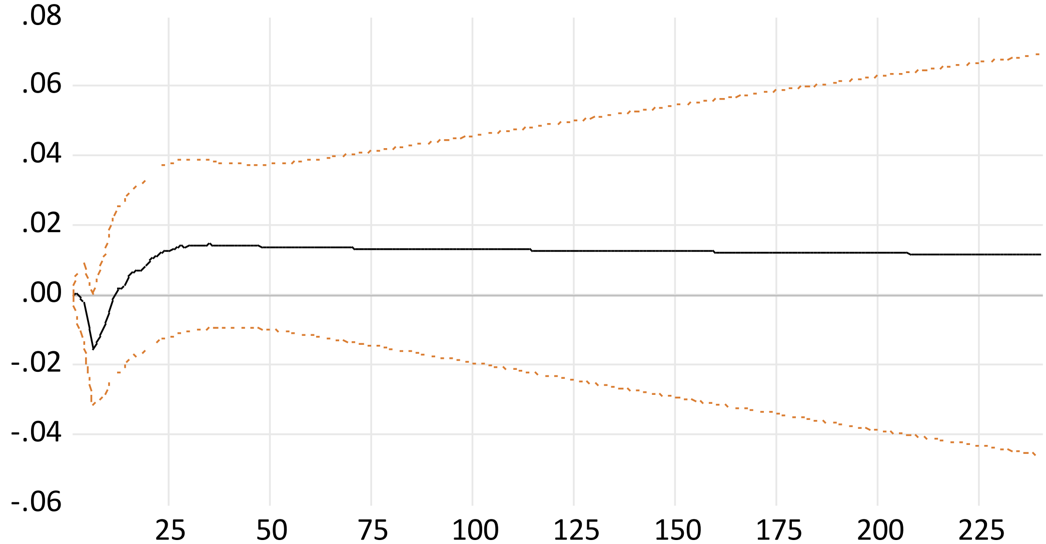

*In response to an inflation shock, stock prices reach a valley of -0.019 in period 9 and reaches par in period 22.*

Figure 20 Response of Canada Small Cap Index to Canada CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



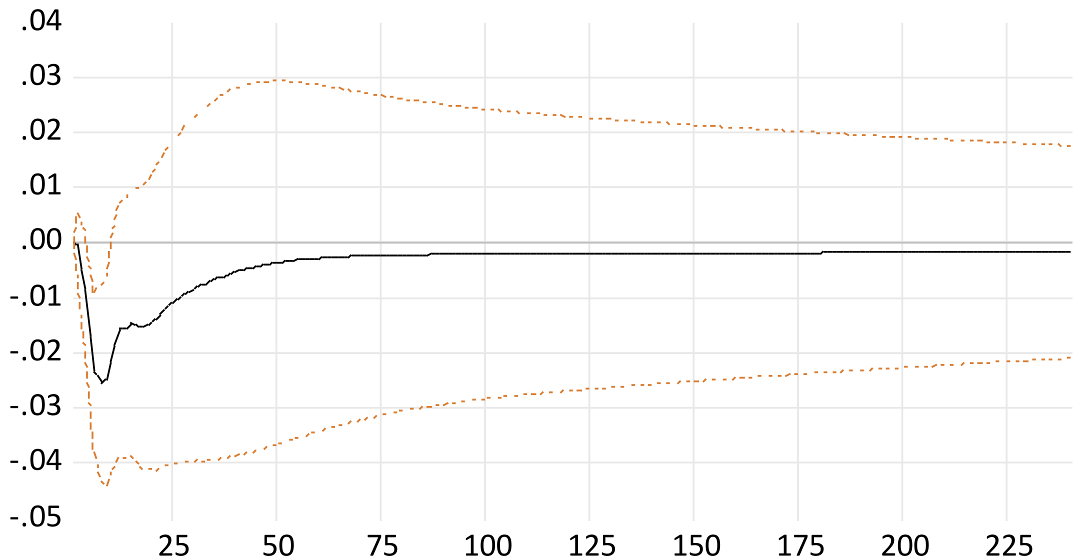

*In response to an inflation shock, stock prices reach a valley of -0.11 in period 9 never reaches par.*

Figure 21  Response of Japan Large Cap Index to Japan CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



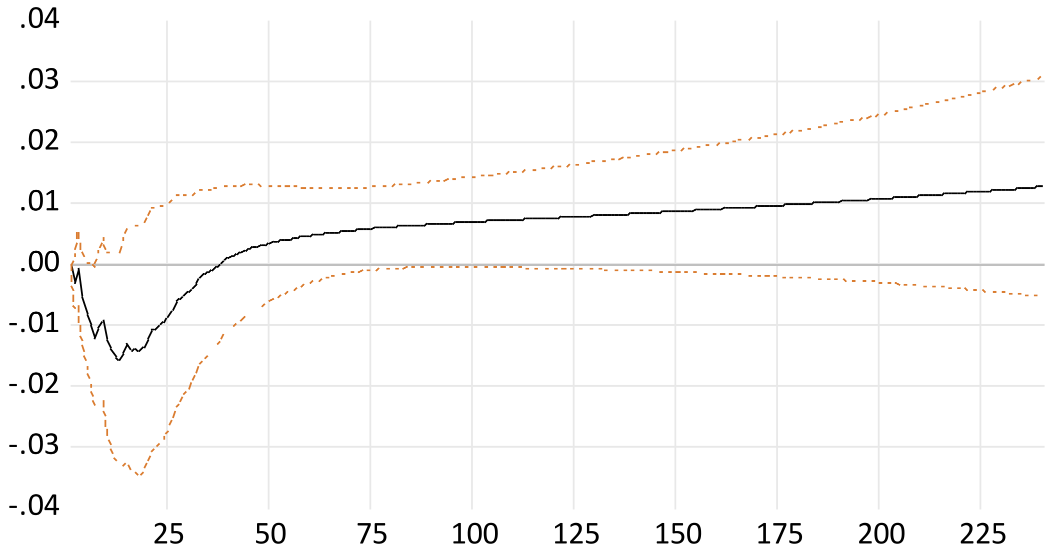

*In response to an inflation shock, stock prices reach a valley of -0.016 in period 6 and reaches par in period 11.*

Figure 22  Response of Japan Small Cap Index to Japan CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



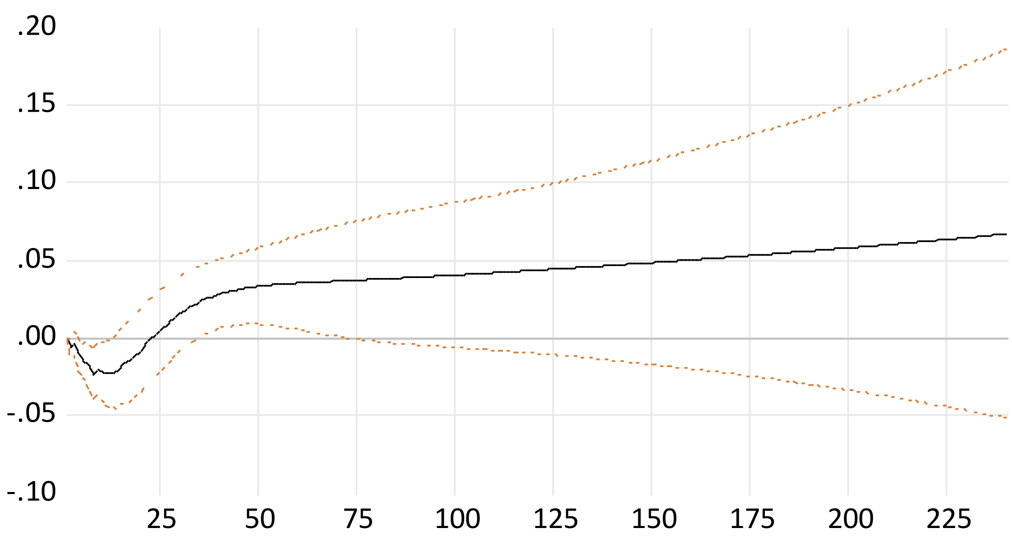

*In response to an inflation shock, stock prices reach a valley of -0.016 in period 6 and reaches par in period 11.*

Figure 23 Response of UK Large Cap Index to UK CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



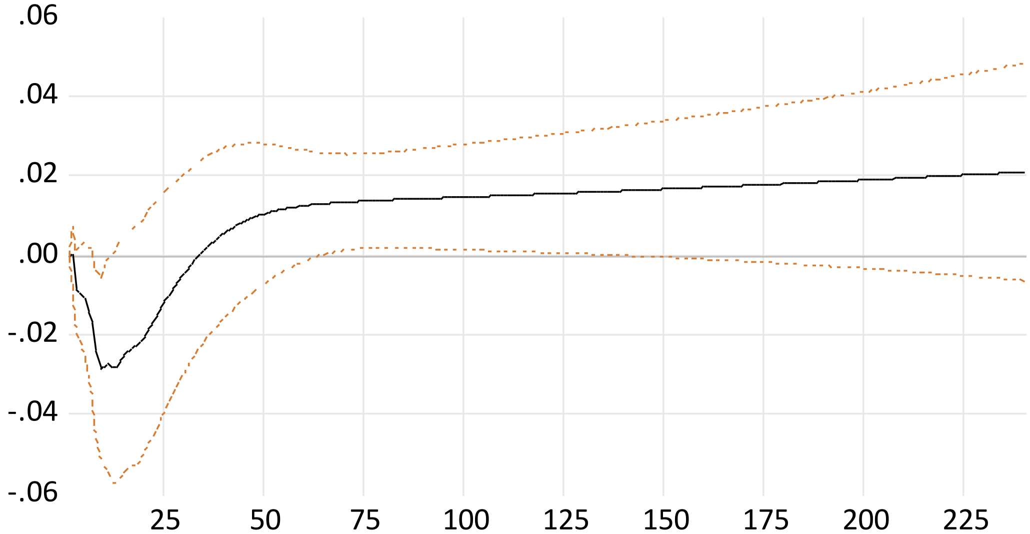

*In response to an inflation shock, stock prices reach a valley of -0.016 in period 12 and reaches par in period 37.*

Figure 24 Response of UK Small Cap Index to UK CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



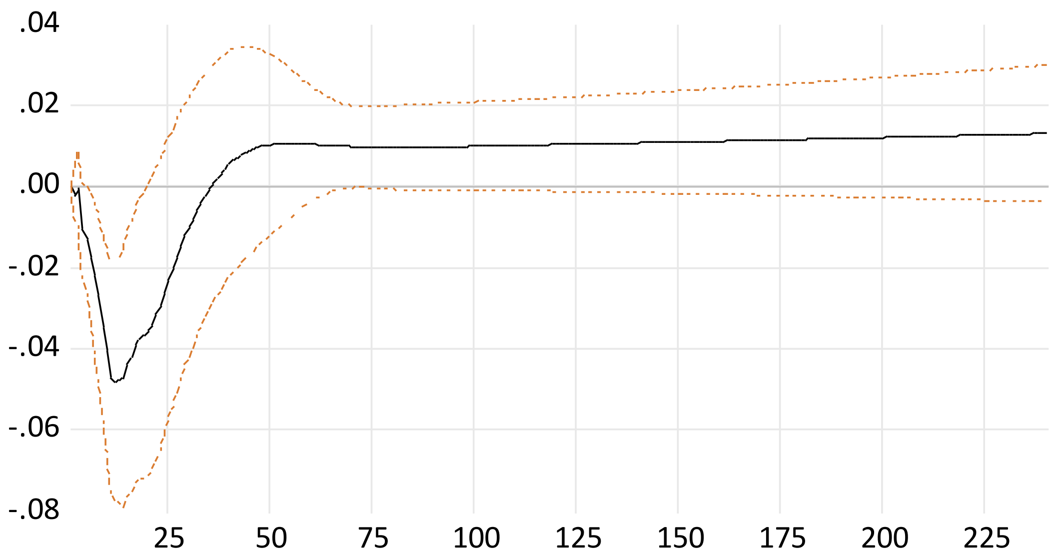

*In response to an inflation shock, stock prices reach a valley of -0.021 in period 10 and reaches par in period 23.*

Figure 25 Response of Germany Large Cap Index to Germany CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



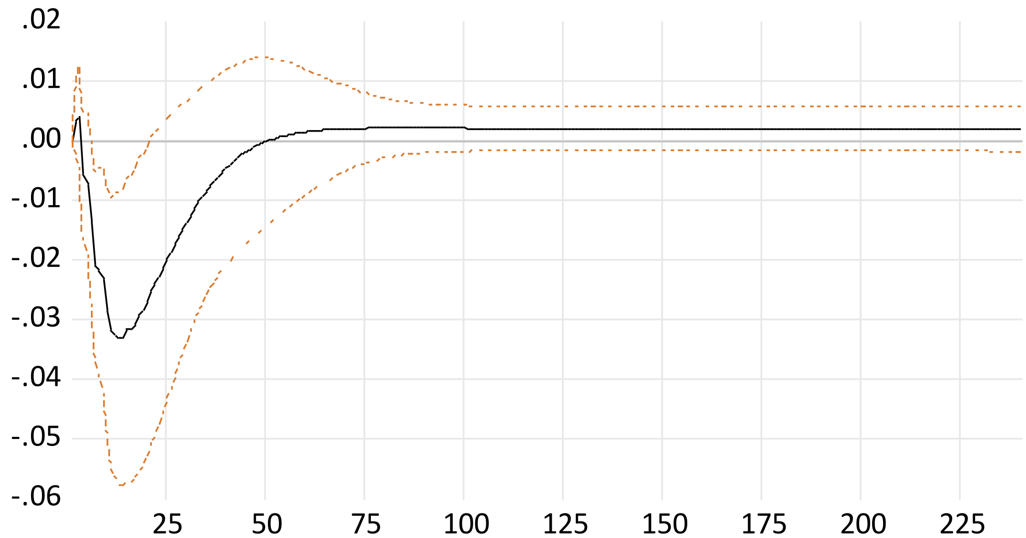

*In response to an inflation shock, stock prices reach a valley of -0.029 in period 9 and reaches par in period 34.*

Figure 26 Response of Germany Small Cap Index to Germany CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



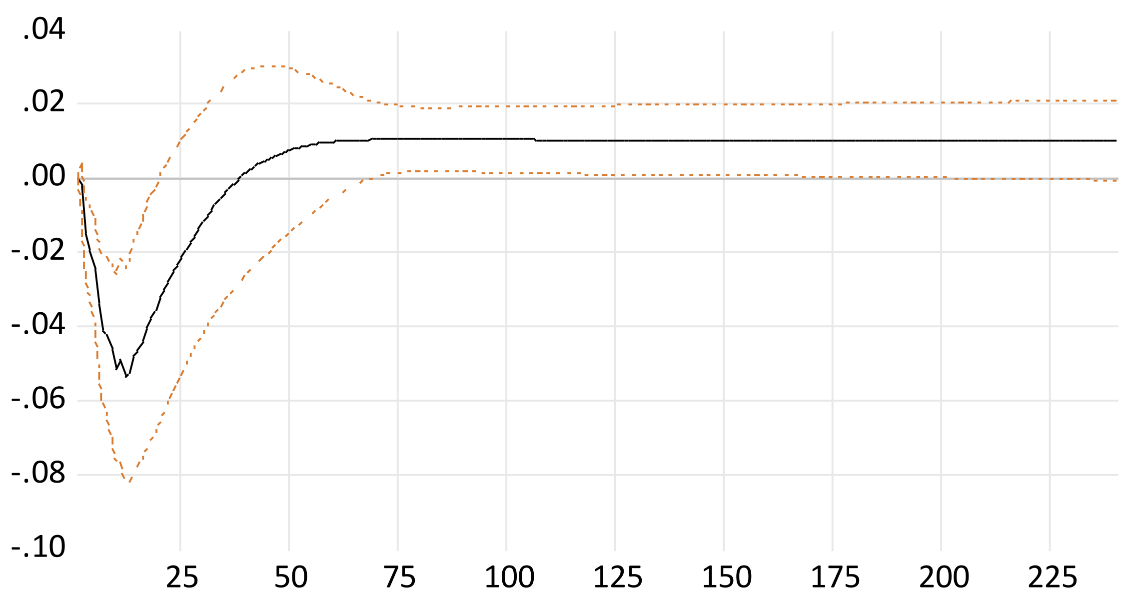

*In response to an inflation shock, stock prices reach a valley of -0.048 in period 13 and reaches par in period 35.*

Figure 27  Response of France Large Cap Index to France CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



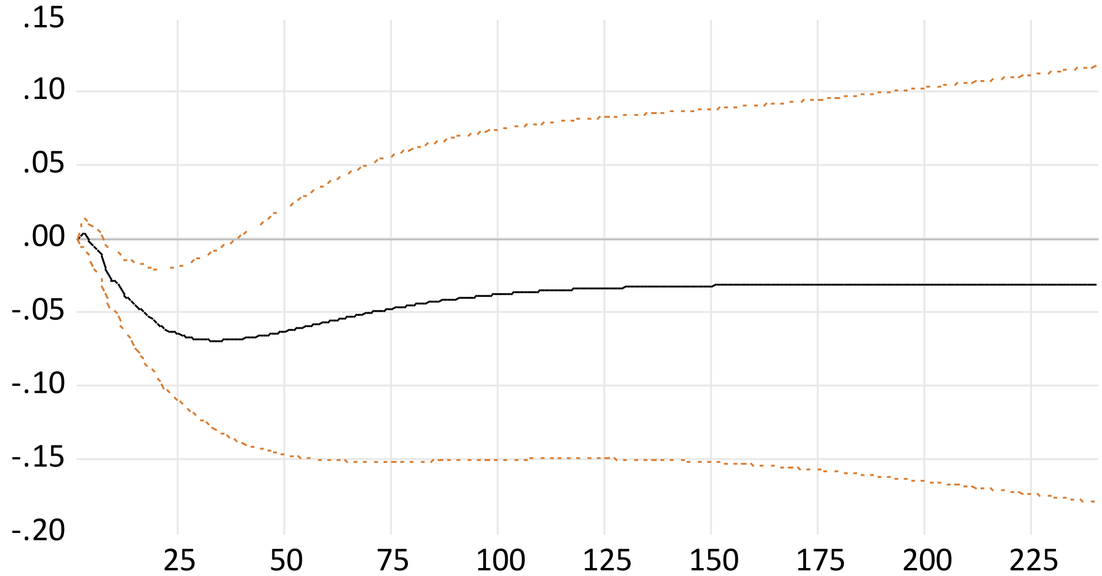

*In response to an inflation shock, stock prices reach a valley of -0.033 in period 13 and reaches par in period 49.*

Figure 28  Response of France Large Cap Index to France CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s



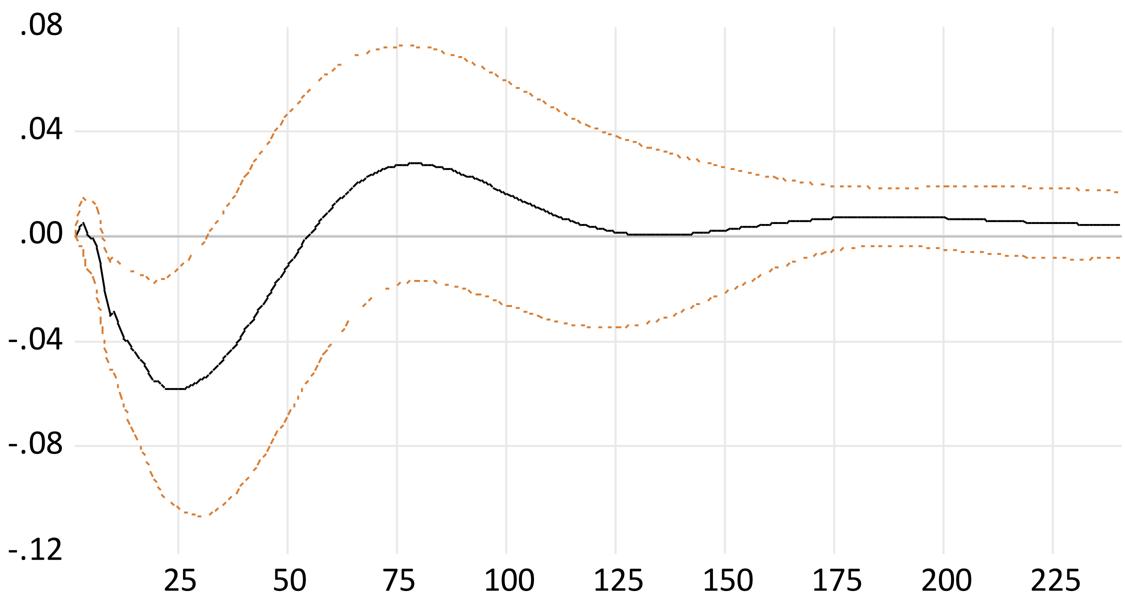

*In response to an inflation shock, stock prices reach a valley of -0.053 in period 12 and reaches par in period 38.*

Figure 29  Response of Italy Large Cap Index to Italy CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s




*In response to an inflation shock, stock prices reach a valley of -0.069 in period 32 and does not return to par.*

Figure 30  Response of Italy Small Cap Index to Italy CPI
Cholesky One S.D. (d.f. adjusted) Innovation
± 2 analytic asymptotic S.E.s




*In response to an inflation shock, stock prices reach a valley of -0.058 in period 23 and reaches par in period 54*

## Discussion

The estimated statistics from the Markov Switching Regression models had mixed results amongst the G7 countries. The outputs for Germany, the UK, and France had highly unstable states. For the UK and France, the expected durations of each state were nearly the same between two regimes which made interpretation of these results difficult. For Germany, where the assumed strong economic state has an expected duration of around 11 months, the only significant variable was the default risk, which had a negative and near-zero relationship with small-cap premiums. Small-cap premiums experienced state changes during the period corresponding to the GFC and Covid-19 market crash, but also is seen to have several other state changes during some periods of market pullbacks, such as in late 2018. The CPI and default risk factors are significant but weakly negative only during the weak economic regime, and not significant for the strong economic regime, indicating that these do not have a strong impact on small-cap prices but when they do it is during periods of possible higher market uncertainty. This single regime significance is consistent with the results in Connolly et al. (2022) who found strong negative relationship between inflation and stock returns for the US in weak economic times only. The coefficients for the US dividend yield and the risk-free rate are as expected, weakly positive for the former, and strongly negative for the latter. Similarly, the results for Canada show a significant CPI variable in both regimes, but a strongly negative relationship with small-cap premiums in the regime associated with a weak economic state, while the negative relationship in the strong economic state is only weakly negative, coefficients being -10.0836 and -0.7455 respectively. Default risk for Canada was not a significant variable in either state but the coefficient was near zero in both cases. The dividend yield in Canada was strongly negatively related to small-cap premiums, at -9.7965, showing that the market punishes small-cap performance when excess cashflows are returned to investors from the market. The coefficients for CPI and default risk in the weak economic state for Italy are both significant but the state is completely unstable with a near zero probability of remaining in the state. That being said, the timing of the regime switches does correspond with periods of economic uncertainty, such as the early 2000s which was right after the adoption of the Euro, during the GFC and during the Covid-19 market crash. The coefficient for dividend yield is also strongly negative for Italy as it is for Canada. The model for Japan did not have the default risk variable, and the risk-free rate was removed to improve the model. The CPI variable is significant and weakly positive only in the regime that corresponds with weak economic states. Dividend yield and term risk are significant, with dividend yield being strongly positive with a coefficient of 3.1709. In summary, the results for the G7 countries in the date period analyzed show that coefficients for CPI seem to be more meaningful during weak economic times and is negative when it is the case. The results for the dividend yield factor show significant but contradictory results country to country and default risk, where it is significant, shows a weak or near zero relationship with small-cap premiums.

Testing for long run relationships and Fisher Elasticities found results consistent with those in the literature which find linear ADF univariate testing insufficient for confirming the stationarity of the time series. The modified-ADF test yielded the strongest statistical power, and the ADF test with breaks showed inconsistent power that was not always greater than the standard ADF test. This result could be due to the lack of large structural breaks in our sample, and we found value to having the results of both the modified-ADF test and ADF test with breaks. However, ~~U~~univariate testing did not find any significant long-run coefficients amongst G7 countries, with the exception of Italy, which is in direct contrast to Boamah (2017) who found the exact opposite to be true, albeit over a shorter time horizon. If the analysis were to end here, the results would indicate that there is no long-run relationship between stocks and goods prices at all, regardless of market capitalization. The Italy Small-cap Index was the only series with a significant relationship with goods prices at 5% significance, however the sign of the error correction term is positive which indicates this is not a trend stationary process. Consistent with our results, Cook (2009) finds I(0) stationarity amongst

OECD countries from 1985-1994, except with Japan, suggesting that ~~inflation~~ goods prices may in fact be a trend stationary process.

The panel analysis results in all I(1) stationary panel groups with cointegrating relationships to have significant Fisher Coefficients, with betas ranging from 1.08 to 1.2269, showing a greater-than-unit Fisher relationship between goods prices and stock prices for these groups (summarized in Table X). The panel groups for all G7 countries, both large-cap and small-caps have significant coefficients in the post GFC period, with small-caps having a greater coefficient, indicating that G7 stocks are a hedge to G7 inflation in years following the crash and small-caps provide greater returns than large-caps per unit increase in inflation. The same is true for isolating for North America (US and Canada) Small-caps, the beta is 1.12 in the post GFC period. Looking at the G7 countries outside of North America; Germany, France, Japan, Italy, and the UK, we see that the Fisher coefficient is only significant in the years prior to the GFC. Interestingly, countries that were shown to not be First-Order stationary I(1) through univariate unit root tests were I(1) stationary when analysed as a panel group; France, Germany, and the UK, the coefficients are significant for both large-caps and small-caps, small-caps again outperforming their large-cap counterparts over the full sample 1994-2022. These results reinforce the conclusion from Omay (2015), which found that ignoring cross-sectional correlation in panel tests has substantial impacts on results. The overall overperformance of small-caps in hedging inflation is supported by the results of Switzer and Fan (2007) who find that small-cap returns from some G7 countries from 1984-2000 are a significant asset class separate from large-caps, which provide diversification benefits through superior risk-adjusted returns.

The short-run VAR analysis shows all coefficients to be significant at the 5% level. As expected, the response to a shock in goods prices is an immediate negative response from stock prices. Canada and Japan's small-cap indices both show an asymmetrically negative reaction to an inflation shock relative to their large-caps, and both do not recover the losses in the 240-month period. The Italian large-cap stock index also does not recover fully, though the small-cap index does. All small-cap indices have a greater negative reaction to the shock than the large-caps, but we see that for the US, UK and France, the small-caps recover in fewer periods than the large-caps. From this we can confirm that the consensus on the negative short-run relationship between stocks and inflation shocks is seen here. The relative underperformance of Canadian and Japanese small-caps, and the relative outperformance of the US, UK, and France small-caps provide an interesting insight into potential investment strategies in the face of inflationary risk.

**Conclusion**

This paper was able to address the two important questions about macroeconomic factors for small-cap premiums, and the inflation hedging properties of small-cap stocks in G7 countries. In agreement with the Connolly (2022) paper, the regime switching analysis reveals that annual CPI rates may have a negative impact on small-cap premiums during weak economic states only, especially in Canada and the US. Default risk had a weak relationship with small-cap premiums in most regimes which seems to contradict some of the literature on the subject. Dividend yields had contradictory signs and large coefficients across the G7, indicating rates of reinvestment for a country have differing impacts on the overperformance of their small-cap stocks.

The Fisher Elasticity tests show that panel unit root tests with cross-sectional dependence yields superior results in terms of establishing first order I(1) stationarity of goods prices and stock prices, which makes intuitive sense as G7 country markets are likely interdependent due to the globalization of economies. The panel Fisher Elasticities in the results are all greater than unit without accounting for taxation, with small-cap panels having greater betas, showing that where a long-run relationship exists, it is an effective hedge against inflation. The panel of all G7 countries' stock prices only had a long run relationship with goods prices in the post-GFC period for both large-caps and small-caps. The short-run response by small-cap stock prices to goods prices agrees with the literature on the asymmetrically negative reaction relative to large-cap stocks. However, some differences between countries reveal potential important areas for further investigation for building an investment portfolio strategy for recovery from inflation shocks.

There were some limitations to the study which could be addressed to gather further insights into these questions. The date range for the series in the regime analysis was limited to post 2003, which meant that data spanned fewer periods of inflationary shocks as inflation has been relatively low for most of the G7 countries in the last few decades. The Fisher Elasticities show that small-caps do demonstrate hedging properties that are greater than large-caps in the panel analysis, however the number of periods required to fully materialize this hedge may be mismatched with the investment horizon of a portfolio manager. This study revealed some important insights on the potential benefits of international diversification for hedging risk from inflation and shows that market capitalization can be incorporated into a style rotation strategy in this context.

## References

Anari, A., Kolari, J. (2001) “Stock Prices and Inflation”. The Journal of Financial Research, Vol XXIV, No. 4, pp. 587-602.

Amenc, N., Esakia M., Luyten, B. (2019) “Macroeconomic risks in equity factor investing”. Journal of Portfolio Management, Vol 45, No. 6, pp. 39-60.

Bampinas, G., Panagiotidis, T. (2016) “Hedging Inflation with individual US Stocks: A long-run portfolio analysis”. The North American Journal of Economics and Finance. Vol 37, pp. 374-392.

Boamah, M.I. (2017) “Common Stocks and Inflation: An Empirical Analysis of G7 and BRICS”. Atlantic Economic Journal Vol 45, pp. 213-224.

Bodie, Z. (1976) “Common Stocks as a Hedge Against Inflation”. The Journal of Finance, Vol 31, No. 2 pp. 459-470.

Chan, K., Gup, B. E., Ming-Shiun, P. (2003) “International Stock Market Efficiency and Integration: A study of eighteen nations”. Journal of Business Finance & Accounting. Vol. 24, No. 6, pp. 803-813.

Chen, S.-W. (2010) “Regime Non-stationarity and Non-linearity in Inflation Rates: Evidence from OECD Countries”. International Research Journal of Finance and Economics. Vol. 46, pp. 47-57.

Chen, X. (2022) “Influence of structural break on the power of ADF unit root test”. Journal of New Economics and Finance. Vol. 2, No. 1, pp. 1-10.

Cheung, Y.-W., Kon, S.L. (1995)“Lag order and critical values of a modified Dickey-Fuller test”. Oxford Bulletin of Economics and Statistics. Vol. 57, No. 3 ,pp. 411-419.

Ciner, C. (2015) “Are equities good inflation hedges? A frequency domain perspective”. Review of Financial Economics. Vol. 24, pp. 12-17.

Cook S. (2008) “Maximum likelihood unit root testing in the presence of GARCH: A new test with increased power”. Communications in Statistics B: Simulation and Computation. Vol 37, No. 4, pp. 756–765

Cook, S. (2009) “A re-examination of the stationarity of inflation”. Journal of Applied Econometrics. Vol. 24, No. 6, pp. 1047-1053.

Connoly, R.A., Stivers, C., Licheng, S., (2022). “Stock returns and inflation shocks in weaker economic times”. Journal of Financial Management. Vol. 51, pp. 827-867.

Culver S., Papell D. (1997) “Is there a unit root in the inflation rate”. Journal of Applied Econometrics. Vol.12 pp.436–444.

Dickey, D. A., Fuller, W.A., (1979) “Distribution of the estimators for autoregressive time series with a unit root”. Journal of the American Statistical Association . Vol. 74 , pp. 427–43

Elliott, G., Rothenberg, T.J., Stock, J.H. (1992) “Efficient Tests for an Autoregressive Unit Root”. National Bureau of Economic Research Technical Working Paper. No. 130.

Fama, E. F., Schwert, W. (1979) “Inflation, Interest and Relative Prices”. The Journal of Business, Vol 52. No. 2 pp. 183-209.

Granger, C. W. J., & Teräsvirta, T. (1993). “Modelling nonlinear economic relationships”. Oxford: Oxford University Press.

Gregoriou, A., Kontonikas, A. (2010) “The long-run relationship between stock prices and goods prices: New evidence from panel cointegration”. Journal of International Financial Markets, Institutions & Money, Vol. 20 , pp. 166-176.

Hamilton, J.D. (1989) “A new approach to the economic analysis of nonstationary time series and the business cycle”. Econometrica Vol. 57, pp. 357–384.

Harvey, D.,I., Leybourne, S.J., Taylor, A.M. (2013) “Testing for unit roots in the possible presence of multiple trend breaks using minimum Dickey-Fuller statistics”. The Journal of Econometrics. Vol 177, No. 2, pp. 265-284.

Hoesli, M., Lizieri, C., MacGregor, B. (2008) “The inflation hedging characteristics of US and UK Investments: A Multi-Factor Error Correction Approach”. Journal of Real Estate Financial Economics. Vol. 36, pp.183-206.

Kremer, S., Bick, A., Nautz, D. (2013) “Inflation and growth: new evidence from a dynamic panel threshold analysis”. Empirical Economics: Journal of the Institute for Advanced Studies, Vienna, Austria. Vol 44, No. 2, pp. 861-878.

Kenneth R. French - Data Library. [online] Available at: <http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html> [Accessed September 2022].

Kremer, S., Bick, A., Nautz, D., (2013) “Inflation and growth: new evidence from a dynamic threshold analysis”. Empirical Economics. Vol. 44, pp. 861-878.

Lee, K. (1999) “Unexpected Inflation, inflation uncertainty, and stock returns”. Applied Financial Economics, Vol. 9, No. 4 pp. 315-328.

Levin, Lin and Chu (2002) "Unit root tests in panel data: Asymptotic and finite-sample properties", Journal of Econometrics, vol 108, Issue 1, pp. 1–24

Liu, J., Serletis, A. (2022) “The complex relationship between inflation and equity returns”. Journal of Economic Studies. Volume 49, No. 1, pp. 159-184

Madapour, S., Asgari, M. (2019) “The puzzling relationship between stocks return and inflation: a review article”. International Review of Economics. Vol. 66, pp. 115-145.

Narayan, P. K., Narayn, S. (2008) “Is there a unit root in the inflation rate? New evidence from panel data models with multiple structural breaks”. Journal of Applied Economics, Vol 42, No. 13, pp.: 1661-1670.

Omay, T. , Hasanov, M. (2010a) “The effects of inflation uncertainty on interest rates: a nonlinear approach”. Applied Economics. Vol. 42 No. 23, pp. 2941-2955.

Omay, T., Kan, E. O. (2010b) “Re-examining the threshold effects in the inflation-growth nexus with cross-sectionally dependent non-linear panel: evidence from six industrialized economies”. Economic Modelling. Vol. 27, No. 5, pp. 996-1005.

Omay, T., Yuksel, As., Yuksel, Ay. (2015) “An empirical examination of the generalized Fisher effect using cross-sectional correlation robust tests for panel cointegration”. Journal of International Financial Markets, Institutions & Money. Vol 35, No. 3, pp. 18-29.

Parikh, H., Malladi, R.,K., Fabozzi, F.J. “Preparing for higher inflation: Portfolio solutions using U.S. equities”. Review of Financial Economics, Vol. 38, No. 3, (2019), pp. 542-554.

Pesaran , H., Shin, Y. (1999) “An Autoregressive Distributed Lag Modelling Approach to Cointegration Analysis”. In Econometrics and Economic Theory in the 20th Century: the Ragnar Frisch Centennial Symposium: Cambridge: Cambridge University Press.

Pesaran, M. H., Shin, Y., Smith, R. J. (2001) “Bounds testing approaches to the analysis of level relationships”. Journal of Applied Econometrics. Vol 16, Issue 3, pp.289-326.

Im K.S., Pesaran M.H., Shin, Y. (2003) “Testing for unit roots in heterogenous panels.” Journal of Econometrics, Vol. 115, Issue 1., pp. 53-74.

Pesaran M.H. (2007) “A simple panel unit root test in the presence of cross-section dependence.” Journal of Applied Econometrics,Vol 22, Issue 2, 265–312.

Said, S.E. , Dickey, D.A. (1984) “Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order”. Biometrika, Vol. 71, pp. 599-607.

Sami, J. (2021) “Stock Market Investment and Inflation: evidence from the United States and Canada”. Review of Economic Analysis. Vol 13 pp. 339-365.

Switzer, L.N. (2010). “The behavior of small-cap vs. large-cap stocks in recessions and recoveries: Empirical evidence for the United States and Canada”. The North American Journal of Economics and Finance. Vol. 21, Issue 3, pp. 332-346

Switzer, L.N., Fan, H. (2007) “Spanning Tests for Replicable Small-Cap Indexes as separate asset classes”. Journal of Portfolio Management, Vol. 33, No. 4, pp. 102-110.

Switzer, L. N, Picard, A. (2020) “The cyclical behavior of the small-cap premium: A regime-switching approach”. Journal of Business, Accounting, and Finance Perspectives. Vol. 2, Issue 1, pp.

Teräsvirta, T. (1994) “Specification, estimation, and evaluation of smooth transition autoregressive models”. Journal of the American Statistical Association,Vol. 89, pp. 208–218.

Tiwari, A.K., Adewuyi, A.O., Awodumi, O.B., Roubaud, D. (2022) “Relationship between stock returns and inflation: New evidence from the US using wavelet and causality methods”. International Journal of Finance and Economics. doi: 10.1002/ijfe.2384

Tsong, C.-C., Lee, C.-F., Lee, C.-C. (2012) “A revisit of the stationarity of OECD inflation: evidence from panel unit-root tests and the covariate point optimal test” The Japanese Economic review. Vol. 63, No. 3, pp. 380-396.

U.S. Bureau of Labor Statistics (USBLS), (2022) “Consumer Price Index for All Urban Consumers: All Items in U.S. City Average [CPIAUCSL]”, retrieved from FRED, Federal Reserve Bank of St. Louis; https://fred.stlouisfed.org/series/CPIAUCSL.

Zhang, Q.J., Hopkins, P., Satchell, S.E., Schwob, R. (2009) “The link between macro-economic factors and style returns”. Journal of Asset Management. Vol. 10, Issue 5, , pp. 338-355