Do political and economic uncertainties separate stock markets?

Hamid Babaei[a,*], Georges Hübner[a], Aline Muller[b]

Affiliations
[a] University of Liège, HEC Management School, Finance and Law, 14 rue Louvrex, 4000 Liège, Belgium
[b] Luxembourg Institute of Socio-Economic Research, Luxembourg
Corresponding author.
E-mail addresses: hbabaei@uliege.be

Abstract

This study investigates the dynamic pattern of the interdependence among G7 stock markets over the 1990-2021 period. The state-space formulation of the time-varying cointegrating coefficient allows us to examine the potential drivers of disruption in the long-run bivariate co-movement of markets. Our results reveal that variations in a number of financial risk factors, including economic policy uncertainty (EPU) and world geopolitical risk (GPR), have a significant impact on the cointegrating coefficient. Further analysis on the comovement of the augmented and the non-augmented cointegrating coefficients suggests that globalization has reduced market segmentation causes to our risk factors.

Keywords: time-varying cointegration, Kalman filter, uncertainty, globalization

Classification codes : F360, F020, E520, D89

Abbreviations : Nominal price return (NPR), Real total return (RTR)

1. Introduction

A number of integrated series are said to be cointegrated if a linear combination of them produces a series with a lower degree of integration. Stock market cointegration has been the object of extensive research in recent decades. Two seminal papers by Engle and Granger (1987) and Johansen (1988) introduced the basic theory and estimation procedure of cointegration relations. Since then, the literature has provided a broad strand of empirical works investigating the long-run equilibrium relation among the world's stock markets, with a particular focus on those of G7 countries (the US, the UK, France, Germany, Italy, Canada and Japan). A number of empirical works find weak evidence supporting time-invariant cointegration among G7 stock market indices. They suggest that there might be a structural break in the cointegrating coefficients. Taylor and Tonks (1989), Arshanapalli and Doukas (1993), Kanas (1998) and Menezes et al (2012) find distinct time-invariant cointegrations in different subperiods based on known economic events. Alternatively, to investigate the evolution of long-run market co-movements over time, some authors have also applied methods based on recursive as well as rolling time windows (see Mylonidis and Kollias, 2010; Pascual, 2003; Rangvid, 2001). A number of theoretical works develop statistical tests of instability in cointegration relations (see Gregory and Hansen, 1996: Hansen, 1992, Kejriawal and Perron, 2010; Quintos and Phillips, 1993). The developments dedicated to the estimation of smoothly varying cointegrating vectors have also been in the spotlight of a number of theoretical works (Bierens and Martins, 2010; Park and Hahn, 1999; Li et al., 2020).

Despite empirical evidence of time-varying cointegration, most works have focused on the estimation of and test procedures for time-varying cointegration. However, the reasons for repeated disruptions in the long run equilibrium relation of stock markets have not been specifically investigated.

Our paper takes this latter point of view: Considering the extant empirical evidence on the variability of the cointegration levels between the main stock markets over time, we aim to investigate the influence of two of the main potential causes of this dynamic behaviour, namely political and economic sources of uncertainty.

Long run equilibrium relations are subject to instantaneous changes as a result of repeated changes in the economic circumstances. We address here the important factor of uncertainty, as evidenced in the economics and finance literature, to substantially influence economic aggregates and, in turn, to drive security prices. There is indeed a broad range of papers investigating the effects of economic uncertainty on a number of subjects including investment, economic growth, firm-level managerial decisions and risk management. Even though uncertainty is used as a proxy for all those variables which themselves may impact cointegration levels, there is no paper that tries to make an explicit and direct link between uncertainty and cointegration. This study contributes to the literature on long-run stock market co-movements by shedding light on the causes of market segmentation that lead to mixed conclusions in the literature on bivariate cointegration between the world’s developed stock markets. We form our hypothesis relying on the findings in the literature implying that uncertainty can be a significant driver of economic aggregates as well as security prices. This fact could be of special interest particularly because expectations of future economic activity can be significantly shifted during periods of rising or declining uncertainty. The state-space formulation of the time varying parameters allows us to study the potential drivers of the instability in the bivariate long-run equilibrium relation.

Among the variables that have been proposed as proxies for economic uncertainty, the index of economic policy uncertainty (EPU) proposed by Baker et al (2016) has received extensive attention. This variable aims to overcome some identified weaknesses of earlier attempts to measure uncertainty. It introduces a news-based index that covers articles containing three categories of pre-specified words pertaining to uncertainty, the economy and policy. Since its introduction, EPU has been widely used in the literature to analyse the effects of uncertainty on macroeconomic aggregates (He et al., 2021; Prüser and Schlösser, 2020; Suh and Yang, 2021; Zhou et al., 2021). A number of works also rely on this EPU variable to conduct analysis relating uncertainty to a range of indicators namely bank credit (Nguyen et al., 2020), exchange rate volatility (Zhou et al., 2020), asset prices (Brogard and Detzel, 2015) and crude oil return volatility (Ma et al., 2019).

The literature has also investigated the effects of the uncertainty associated with political crises and tensions on the economy and financial markets (see Bloom, 2009 and

Cheng and Chiu, 2018, Le and Tran, 2021, Omar et al., 2017). Since the introduction of the geopolitical risk (GPR) index developed by Caldara and Iacoviello (2021), a number of studies have been used it as a proxy for the uncertainty associated with political instabilities, as it is found to be a key determinant of stock market dynamics (Antonakakis et al., 2017; Balcilar et al., 2018; Das et al., 2019; Kannadhasana and Das, 2020). As indicated by Caldara and Iacoviello (2021), being exogenous to business and financial cycles, the GPR index can provoke financial volatility and policy uncertainty. For the same reasons as those justifying the use of the EPU variable as a direct explanatory factor of the time variations in market cointegration levels, we adopt the GPR index introduced by Caldara and Iacoviello (2021) as the most convincing proxy to date for the uncertainty cast by geopolitical instability.

We apply a Kalman filter model to two sets of monthly stock market indices during the 1990-2021 period. As most of the literature uses either nominal price return (NPR) indices or real total return (RTR) indices, we use both sets of the indices to further explore their reactions to uncertainty innovations. EPU and GPR are found to be two significant causes of repeated disruptions in the cointegrating vector, as they can impact the expectations of investors. Our empirical findings build on the literature on uncertainty implying that a shock to EPU and GPR leads to market segmentation. The dynamics of the cointegrating coefficients are found to significantly depend on the variations in the uncertainty and financial indicators. This outcome suggests the existence of a path dependence of market equilibrium relations, as human decision making drives the uncertainty indicators of both economic policy and geopolitical tensions.

The structure of this paper is as follows. Section 2 briefly provides a review of the literature on cointegration, and explains empirical works and theoretical developments in the stability analysis of cointegration relations. The section also proposes a review of the literature on applications of the EPU and GPR indices. Section 3 presents the Kalman filter model of this paper to study the drivers of time-varying cointegration. Section 4 explains our data and preliminary analysis on bivariate cointegration. In section 5, a stability analysis on bivariate cointegration is conducted. Section 6 discusses the drivers of cointegration dynamics. The seventh section examines in detail the incremental information brought by the augmented model. Finally, section 8 concludes the paper.

2. Literature review

2.1. Evidence of time-varying cointegration

The conjoint evolution of worldwide financial markets hinges on a broad range of economic and political events. Factors such as the abolition of exchange rate control, free capital flow, market deregulation, developing trade ties, fiscal and monetary policies, economic crises, international political conventions and political conflicts are widely accepted as having significant impacts on stock prices. Taylor and Tonks (1989) find that after the abolition of restricted capital movement in 1979, the UK stock market became cointegrated with that of Germany, Netherlands and Japan. They find no evidence for cointegration between the UK and US stock markets after 1979. Arshanapalli and Doukas (1993) find evidence for cointegration between the US stock market and the UK, French and Japanese stock markets from 1980 to 1990. Their study does not support cointegration between the US and Germany in the pre-crash subperiod of 1980 to 1987. Kanas (1998) applies three econometric methods to study the cointegration of the US stock market from 1983 to 1986 with six major European markets, namely, those of the UK, France, Germany, Italy, the Netherlands and Switzerland. They find no evidence of cointegration for the entire period. Their pre- and post-crash subperiod analysis confirms this conclusion as well.

Mylonidis and Kollias (2010) adopt a rolling time window approach to examine the instability of the long-run cointegration relation in the post-euro era stock markets of Germany, France, Spain and Italy. They find only weak evidence of convergence: Their rolling trace statistic remains generally below the critical value. Rangvid (2001) recursively applies the Johansen rank test to the UK, French and German stock markets from 1960 to 1999. The test statistics reveal upward sloping after 1982. In particular his analysis reveals that a decreasing number of stochastic trends drives stock markets. Pascual (2003) recursively estimates the Johansen cointegration model for the UK, French and German stock markets from 1960 to 1999. Starting from a five-year time window of quarterly data, his estimation continues by adding new quarterly data to the end of the window. Despite the upward slope, the trace statistic does not exceed the 5% critical value. Menezes et al. (2012) study the stock market integration of G7 countries in both price levels and returns from 1973 to 2009. Applying the methodology developed by Gregory and Hansen (1996), they find a structural break in the cointegration relation of G7 stock markets.

In line with the poor empirical evidence of stable cointegration in the world's largest stock markets, a broad range of papers has developed necessary techniques as well as proofs to test and estimate time-varying cointegration. A number of works address sudden structural changes in the parameters and develop procedures and methodologies to test and detect structural breaks. Hamilton (1989) assumes discrete shifts in the dynamics of the cointegration relation. He uses Goldfeld and Quandt's (1973) Markov switching regression to identify structural changes in the cointegration parameters. Gregory and Hansen (1996) provide a residual-based test for the null hypothesis of no cointegration versus a regime change in either the intercept alone or the cointegrating vector. They assume that the timing of the structural shift in the cointegrating vector is unknown. Hansen (2003) models the structural change in the Johansen cointegration model. His framework allows for structural change in any subset of the parameters. He assumes that the change point and the number of cointegration relations are known *a priori*.

Instead of positing structural discontinuities, some studies assume that the cointegration parameters change smoothly over time. Hansen (1992) extends the Lagrange multiplier test of parameter change proposed by Nyblom (1989) to the context of fully modified cointegration of Phillips and Hansen (1990). He formulates a test in which, under the alternative hypothesis, the parameters follow a random walk. He argues that, since such hypothesis of a random walk in the intercept implies no cointegration relation, the test statistic represents a null hypothesis of cointegration against the alternative of no cointegration. Park and Hahn (1999) investigate the smooth evolution of cointegrating coefficients using a nonparametric method. Hansen and Johansen (1999) use the recursive estimation of cointegration parameters to derive the limiting distribution of the fluctuation test proposed by Ploberger et al. (1989) and the Lagrange multiplier test proposed by Nyblom (1989). The fluctuation test is used to test the constancy of recursively estimated eigenvalues. Bierens and Martins (2010) use Chebyshev time polynomials to estimate smoothly changing cointegrating vectors. Li et al. (2020) use a nonparametric approach to estimate a time dependent cointegration relation. They apply the Nadaraya-Watson kernel method to estimate time varying cointegrating vector.

As shown in the papers above, the time variations in the level of market cointegrations are considered either to be abruptly discontinuous, or continuous but smooth. The investigation of a middle path between them, i.e., continuous but volatile changes, would represent a potential way to reconcile these two presumably conflicting approaches.

Moreover, by proposing a dynamic time-varying cointegration framework, we can study the causes of the instabilities in bivariate long-run relations. As the literature emphasizes that economic and political events form a turning point in the history of financial markets, hence leading to structural breaks in the equilibrium relations, we use two recently quantified indicators to represent the uncertainty stemming from human decision making. By relating a state-space model of the joint evolution of stock markets to a set of uncertainty and financial variables, we investigate whether cointegration relations are susceptible to reflecting more repeated changes caused by an uncertain world rather than exogenous structural breaks.

2.2. The role of uncertainties in business cycles and security prices

As mentioned above, it is widely accepted that financial and investment decisions are profoundly affected by uncertainties about economic policies. A major strand of papers investigates the effects of uncertainty on macroeconomic aggregates, namely, investment, economic growth and business cycles. As indicated in the literature, economic uncertainty can dramatically decrease investment and cause business cycles through the real option channel because of the irreversibility of investment decisions. Bernanke (1983) offers an explanation for the fluctuations in cyclical investments by arguing that an optimizing investor postpones economically irreversible investments under uncertainty. Assuming irreversible investment in a competitive market, Caballero and Pindyck (1996) conclude that uncertainty adversely affects firm investment by decreasing the marginal profitability of capital. Bloom (2009) structurally analyses the impact of uncertainty shocks on the hiring rate and economic growth in the aftermath of a shock. Although he admits that uncertainty increases during economic and political shocks, he uses volatility-related measures of uncertainty in his work. Using an endogenous cross-sectional volatility-based measure of uncertainty, Bloom et al (2018) conclude that uncertainty shocks lead to a drop in economic growth. Ludvigson et al. (2021) argue that the conclusions of the works conducted by Bloom (2009) and Bloom et al. (2018) are robust to the use of exogenous measures of uncertainty. The conclusion drawn by Bloom et al. (2018) is also emphasized by Baker et al. (2016) using their news-based exogenous uncertainty measure.

Uncertainty has also been documented to be a significant cause of financial market volatility. Generalized autoregressive conditional heteroscedasticity (GARCH)-type models have been used in the literature to analyse the predictability of the stock market or exchange rate volatility based on uncertainty variables. Brogard and Detzel (2015) find that changes in EPU and market returns have contemporaneous negative correlations and that the current level of EPU and future market excess returns are positively correlated. Having estimated factor-mimicking portfolios for EPU and two other uncertainty indicators, they find significant exposure to uncertainty. Liow et al. (2018) study the interaction between the EPU spillover and financial market volatility spillover of major economies of the world. They find that EPU spillover drives the financial market volatility spillover. Ma et al (2019) use GARCH-mixed data sampling (GARCH-MIDAS) modelling, which can combine low-frequency EPU data with high-frequency return data, to study the impact of EPU in major crude oil importing and exporting countries on oil price volatility. They find that EPU indicators significantly forecast oil price volatility. Using an error correction model (ECM) Aladesanmi et al (2019) investigate the evolution of the conditional correlation between the US and UK stock markets and find that the conditional correlation between the EPU of the two countries significantly influences the conditional correlations of the stock markets. Zhou et al. (2020) study the impact of the ratio between the EPU of China and the US on Chinese exchange rate volatility using the GARCH-MIDAS approach. They find that the ratio significantly predicts long-term and short-term exchange rate volatility. In a similar study, Yu and Huang (2021) apply GARCH-MIDAS modelling to study the impact of Chinese EPU on China stock market volatility. They find that a change in the EPU level significantly provokes volatility.

3. Time-varying cointegration model

To the best of our knowledge only very few papers apply the Kalman filter in the context of long-run equilibrium relations between economic or financial variables. Haldane and Hall (1991) apply the Kalman filter to study the central role of the Deutchmark-Dollar exchange rate on the Dollar-Sterling and Deutchmark-Sterling exchange rates. They find a high degree of variation in bivariate cointegration parameters. Serletis and King (1997) measure stock market integration by regressing the difference between a certain market index with Germany on a constant and the difference between Germany and the United States. They estimate the time-varying parameters of this bivariate regression model using a Kalman filter.

Despite the growing body of theoretical and empirical literature on time-varying cointegration, the reasons for the underlying causes of the variation in the cointegration relation over time between the largest stock markets worldwide remain an open question. A state-space formulation of bivariate equilibrium relation allows us to tackle this question by relating the varying cointegrating coefficient to the presumed driving factors of the dynamic interdependence between market indices.

To shed more light on the mixed findings in the literature, we study two sets of indices, namely, the NPR and the RTR, for each country, as both sets of variables are used in the literature.[1] The use of RTR indices is motivated by a number of works that discuss the cointegration levels of return indices. Richards (1995) argues that return indices are not cointegrated because of permanent country-specific components that drive their behaviour. In their theoretical work, Lence and Falk (2005) investigate the interrelationship between integrated markets and cointegrated prices. Having defined integrated markets based on the absence of arbitrage opportunities, they conclude that regardless of market integration, the necessary (but not sufficient) condition for two price indices to be cointegrated is that their supply processes are themselves cointegrated. They theoretically derive the same conclusion for the return indices.

We address three potential variables that can presumably change investors' expectations about firms' future cash flows, hence leading to a new equilibrium relation. As noted in the literature, the EPU and GPR indices are candidate proxy variables for economic and political uncertainty conditions, which can potentially cast serious doubts on future economic activity.

In our time-varying cointegration model, the innovations in the cointegrating coefficient are driven by the innovations in the uncertainty variables. This is represented by the following two variations of the state-space model:

$$\log(US_t^{RTR}) = a + b_t \log(X_t^{RTR}) + \varepsilon_t \qquad (1\text{-a})$$

$$\log(US_t^{NPR}) = a + b_t \log(X_t^{NPR}) + \varepsilon_t \qquad (1\text{-b})$$

$$b_t = b_{t-1} + \sum_{l=0}^{L} \boldsymbol{\gamma}_{\mathbf{t-l}}^{\prime} \mathbf{U}_{\mathbf{t-l}} + \boldsymbol{\omega}_{\mathbf{t}}^{\prime} \mathbf{F}_{\mathbf{t}} + \nu_t \qquad (2)$$

---

[1] Taylor and Tonks (1989), Arshanapalli and Doukas (1993), Rangvid (2001), Mylonidis and Kollias (2010), and Menezes et al (2012) are among those who use price indices. Richards (1995) and Gilmore et al (2008) are among those who use total return indices.

Equations (1-a) and (1-b) represent Engle-Granger cointegration relations between RTR and NPR indices, respectively. In both equations US_t and X_t represent the US and the domestic indices, a is a constant, b_t represents the time-varying cointegrating coefficient, $\varepsilon_t \sim N(0, \sigma_\varepsilon)$ and $\nu_t \sim N(0, \sigma_\nu)$. In equation (2) which applies to both the RTR and NPR models, the line vector $\boldsymbol{\gamma}_{\mathbf{t-l}}^{'} = (\gamma_{US,l} \quad \gamma_{EU,l} \quad \gamma_{GPR,l})$ gathers the coefficients of the l-th lag uncertainty variable and $\mathbf{U}_{\mathbf{t-l}}^{'} = [\Delta\log\,(EPU_{US,t-l}) \quad \Delta\log\,(EPU_{EU,t-l}) \quad \Delta\log\,(GPR_{t-l})]$ forms the column vector of the uncertainty variables. The variables $EPU_{US,t}$ and $EPU_{EU,t}$ represent the economic policy uncertainty (EPU) in the US and European Union, respectively, at time t, and GPR_t represents the world geopolitical risk index at time t. The vector $\mathbf{F_t}$ contains a number of stock return determinants, including the US bond-to-equity ratio (BEER), US credit spread, US dividend yield, change in the US implied volatility index (VIX), US and domestic inflation rates, and US and domestic term spreads.

As equation (2) indicates, we assume that a shock to the uncertainty variables can trigger a disruption in the cointegrating coefficient. Moreover equation (2) allows the cointegrating coefficient to fluctuate based on a random walk process when the uncertainty variables remain constant. We also include lagged values of the innovations in the uncertainty variables because we let the earlier values of the uncertainty innovations impact the cointegrating coefficient. The intercept in equations (1-a) and (1-b) is constant because if it is added to the state-space, one can no longer refer to equation (1-a) or (1-b) as a cointegration relation.

4. Data and preliminary evidence

We use monthly stock market indices of the US (S&P 500), France (MSCI index), Germany (MSCI Germany), the UK (FTSE all-share index), Italy (MSCI Italy), Canada (S&P TSX composite index), and Japan (TOPIX all share index) retrieved from DataStream during the period from January 1990 to January 2021. These indices are selected to cover a large part of the market capitalization throughout the sample period. All indices are expressed in terms of US dollars as a common numeraire to avoid any contamination by a currency effect. The sample paths of the time series of these indices, with normalized starting points, are represented in Figure 1. The US credit spread is calculated as the Moody's Aaa minus Baa corporate bond yield. The US terms spread is calculated as the 10-year minus 3-month treasury constant maturity rates and that of the other G7 member countries are calculated as the long term government bond yield minus 3 month interbank rates. The US BEER is calculated as the ratio of the US 10-year treasury constant maturity rate to S&P 500 earnings yield. The VIX, the inflation rates and all the interest rates used in the calculation of the financial factors are downloaded from the federal reserve bank of st. louis.

[insert Figure 1 here]

Table 1 represents the results of unit root tests.

[insert Table 1 here]

The unit root tests are performed using the augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. The null of a unit root in the data are not rejected for any of the indices.

Given this set of results, we subsequently perform standard time-invariant cointegration tests depending on the choice of the dependent variable. The results of these tests, which we interpret as the base case, are displayed in Table 2.

[insert Table 2 here]

The Engle-Granger cointegration test does not reject the null hypothesis of the absence of cointegration in most of the bivariate relations. Only in the single case of the bilateral relation of the US stock market with Germany, does the z-statistic weakly reject the null hypothesis at 10% significance level. As our sample covers a relatively long period, such results are expected. During the era starting from 1990 to 2021, numerous political and economic crises have impacted the evolution of the world's stock markets, some of which are supposed in the literature to cause a structural break in the cointegration relations. We do not consider any abruptly and exogenously discontinuous cointegration structure here because one cannot necessarily refer to a certain moment of time as the onset or the ending point of an economic or political crisis.

5. Evidence of dynamic cointegration

Prior to the estimation of time-varying cointegration through the state-space formulation of equations (1) and (2), we apply the Hansen stability analysis of bivariate cointegration relations.[2]

Hansen (1992) develops a distributional theory for three test statistics namely F_{nt} (or L_c test first proposed by Gardner (1969)), $SupF$ (proposed by Quandt (1960)) and $MeanF$, in a regression with integrated processes. The underlying null hypothesis posits constant cointegration coefficients, whereas the alternative hypotheses are an unknown structural break and a random walk in the parameters. Hansen emphasizes that rejecting the null hypothesis implies that the standard cointegration model with stable coefficients is rejected. We focus here

[2] There are a number of tests aimed at examining parameter stability, most of which assume a regime change or smooth variation of the parameters. The Hansen stability analysis fulfils our objective, as it tests the null hypothesis of static cointegration against the alternative of a gradual change in the cointegrating vector.

on the $MeanF$ and L_c statistics as in the underlying alternative hypothesis, the cointegration parameters follow a random walk. Hansen (1992) argues that since an intercept following a random walk implies the absence of cointegration, the corresponding alternative hypothesis of L_c is no cointegration. We apply the Hansen stability test to the bivariate cointegration of the US stock market with each of the 6 other selected developed markets. As Table 3 shows, the null hypothesis of time invariant cointegration is strongly rejected in the majority of the relations. The $MeanF$ test statistics indicate that the null hypothesis of static cointegration is rejected in 4 bivariate relations at the 5% significance level. The L_c statistic rejects the null of cointegration against the alternative of no cointegration, at the 5% level, in 3 bivariate relations for the RTR indices. It rejects the null hypothesis at the 5% significance level in 2 bivariate relations for NPR indices, and for two other bivariate relations, the null hypothesis is rejected at the 10% significance level. For two bivariate relations, i.e., those of the US with Germany and Canada, the L_c statistic does not reject the null hypothesis in either of the indices.

[insert Table 3 here]

Both sets of indices demonstrate a high degree of instability in the cointegrating coefficient. As $SupF$ indicates, the bivariate cointegration between the US and most of the other nominal indices strongly incurs a sudden structural break in an unknown time in the cointegrating coefficient. Only in the single relation between the US and Canada do none of the statistics reject the null hypothesis of static cointegration. The $MeanF$ statistic of five equilibrium relations between nominal indices evidence an unstable cointegration relation over the entire time horizon.

Rather than always displaying the same expectations irrespective of the economic conditions, it is more likely that rational investors adjust their investment decisions based on the evolution of the economic benefits of securities. As RTR indices evolve mostly based on the realized economic benefits, the investment decisions of a rational investor will probably drive real RTR indices towards equilibrium relations. However, in five bivariate relations the p-values of $SupF$ strongly support an unknown sudden structural break in the relation. The corresponding p-value of the $MeanF$ statistic in four bivariate relations between the NPR indices and in five bivariate relations between the RTR indices displays particularly significantly low values, pointing to a dynamic equilibrium relation. In line with the NPR indices, the $MeanF$ statistics of the RTR indices for the US-Canada connection do not reject the null hypothesis of a static cointegration relation. Given that evidence we maintain this pair of indices in our analysis of the state space estimation.

The state-space variable in equation (2) is augmented by the vectors $\mathbf{U_{t-1}}$ and $\mathbf{F_{t-1}}$ whereas it can also be formulated as a simple random walk. Estimating the time-varying cointegration relation excluding the vectors $\mathbf{U_{t-1}}$ and $\mathbf{F_{t-1}}$ results in a model that we refer to as the non-augmented model. The estimated state-space of the augmented and non-augmented models is denoted by $\hat{\hat{b}}_t$ and $\hat{b}_t$ respectively.

As Figure 2 shows, the cointegrating coefficients for all the bivariate relations appear to follow an increasing trend until the mid-1990s, from whence they stabilize until the economic recession of the early 2000s. The second increasing trend is associated with the economic growth preceding the global financial crisis of 2008. Following that crisis, the estimated cointegrating coefficients follow a rather steep upward trend until the end of the sample period. It seems that the Covid-19 crisis is not, unlike the earlier crises, associated with a drop in the cointegrating coefficients. Since the indices are in terms of the natural logarithms, one may address the differences as the returns and try to relate the gradual adjustments in the long-run equilibrium relation to the rate of returns. The first differences have to be expressed in terms of the ECM, in which we cannot relate the contemporaneous rates of returns of the underlying markets. Thus, if we are to relate the contemporaneous rates of returns of two markets, we have to use the estimated equilibrium relations (1-a) or (1-b) and refer to them as the relation between long-run rates of returns. The steadily disrupted cointegrating coefficients over the three periods showing an upward trend are the result of the higher long-run risk premium that the US market receives compared to the other G7 markets. According to Figure 2, during economic crises, the US market is also more penalized in the long run than the other markets. We can refer to the augmented model as the long-run risk-adjusted equilibrium relation in which the equilibrium relation subsumes the long-run relative risk-adjusted rate of return of both markets.

[insert Figure 2 here]

In Figure 3, we report the dynamics of the difference between the augmented and the original time-varying cointegration coefficients. During bearish periods, the augmented (i.e., the long-run risk-adjusted) cointegrating coefficient falls faster than the non-augmented coefficient. As the indices are in natural logarithms, the discrepancy between both sets of cointegrating coefficients yields insight into the relative long-run risk-adjusted return of the US market vis-à-vis other markets. An upward (downward) sloping discrepancy variable is associated with a downward (upward) sloping risk-adjusted cointegrating coefficient. This, in turn, leads to a lower (higher) relative long-run risk-adjusted rate of return of the US market. The US market has higher exposure in the long-run to uncertainty and financial risk factors than the other G7 markets as the discrepancy variable takes a downward trend during periods of economic growth and follows an upward trend during periods of economic recession.

[insert Figure 3 here]

The relative long-run risk premium of the US-Japan relation behaves in a distinct manner. During the years before the global financial crisis, it exhibits substantial fluctuations.

During the Asian crisis, the recession in the early 2000s and the global financial crisis, it spiked and then followed a downward trend until 2021. The relative long-run risk premium of the US to that of Japan is much higher than that of the other bivariate relations. Figure 3 suggests that the US market declined more sharply than all the other G7 markets during the global financial crisis, especially Japan. Moreover, the overall downtrends of the discrepancy variables imply a faster increasing trend in the relative long-run risk-adjusted rate of return of the US market. The discrepancy variable corresponding to the bivariate relation with Canada appears quite distinct, as it follows an upward trend from 2012 until 2016 and then stabilizes.

6. Drivers of the cointegration dynamics

Although the literature largely documents a significant impact of EPU and GPR on stock market evolution, the impact of these variables on the cointegration dynamics remains largely unexplored to date. Through our modelling approach, we associate, in equation (2), the level of market segmentation with not only financial stock return determinants but also, more importantly, innovations in the uncertainty associated with economic policy and political crises. This section investigates the results of our estimation.

As most of the literature evidences the interrelationship between uncertainty and macroeconomic as well as financial variables through the real option channel (see Bernanke, 1983; and Bloom, 2009), we base our hypothesis on the implications of these conclusions for the dynamics of the interdependence of stock markets. Investment, demand and other macroeconomic aggregates can be adversely affected by shocks to uncertainty, causing financial markets to react to these macroeconomic shocks. This can presumably lead stock market indices to depart from long-run co-movement relations. In other words, our hypothesis is motivated by the argument that a varying level of uncertainty can change optimal investment decisions and hence lead to time-varying equilibrium relations, which in turn will affect the level of market cointegration.

6.1. Overall evaluation

Table 4 presents the results of our estimation of equations (1)-(2) up to the first lagged values of innovations in EPU and GPR.

[insert Table 4 here]

The results of the ADF test reveal that the residuals are strongly stationary. As Table 4 indicates, the innovations in the uncertainty variables are generally negatively correlated with the state-space variable when the coefficients are significant. This finding suggests that the US market is usually more penalized in the long-run than the markets of the other G7 members by positive innovations in either of the uncertainty variables. That is the US stock market investors incur larger risks than investors in the other G7 markets.

The estimation results of the RTR indices reveal that GPR is a significant driver in all six bivariate relations. The association is strongest in the bivariate relation with the UK and Japan. The bivariate relation with Canada is found to be significantly affected by innovations in GPR and by innovations in EPU in the European Union and in the US. This relation is also significantly affected by the uncertainty variables when applying our model to the NPR indices. To some extent, this fact positions Canada in a specific situation. The bivariate relation with three European countries, i.e., the UK, France and Germany, remains unaffected by innovations in either of the EPU indicators. The instability in the equilibrium relation of the US with two non-European countries, i.e., Canada and Japan, is more responsive than others to innovations in EPU. The results of the estimation using NPR indices reveal that the US EPU induces a significant disruptive impact on the equilibrium relations with Canada and

Italy. In the bivariate relation with Germany, none of the uncertainty variables is found to drive the cointegrating coefficient.

Regarding the financial variables, the DVIX is found to be strongly significant across all bivariate relations. Used as an endogenous measure of uncertainty in the literature (Bloom, 2009), implied volatility turns out to significantly influence market co-movement. The negative correlation of the volatility indicators with the cointegrating coefficient suggests that when markets become more nervous, this penalizes the US stock market in the long-run to a greater extent than it penalizes the other markets. The US credit spread is also found to be a major driver of the dynamics in the bivariate cointegration relations, except for the UK. This finding tells us that the US market is riskier than the other G7 markets with respect to our state-space indicators.[3] The US dividend yield also has a positive significant correlation with the cointegrating coefficient across all bivariate relations. It is found to be a significant driver of time-varying cointegration when using either the RTR or NPR indices.

7. Incremental information of the augmented model

Augmenting the time-varying cointegration model with the EPU, and GPR indicators and stock return determinants appears to be informative regarding the significance of the associated coefficients. Nevertheless, it is necessary to compare the evolution of the augmented state-space with that of a Kalman filter incorporating a simple random walk as the state space. The fit of

[3] Naturally, since the BEER, DIV, and DVIX variables are measured relative to the US market, we expect their impact to be greater on the US stock exchange than on the other stock exchanges. Nevertheless, the sign of their coefficients is informative regarding how the discrepancy reacts to their evolution.

the augmented model vis-à-vis the non-augmented model can be examined by conducting an F test on the variance in the error terms of the measurement as well as the state-space variables. As indicated in Table 5, augmenting the state space leads to a significant increase in the variance in the measurement error term in all the bivariate cointegration relations. On the other hand, the augmented model significantly reduces the variation in the cointegrating coefficients. In other words, augmenting the state-space reduces the error variance in the cointegrating coefficient whereas it increases the variance in the dependent variable and its error term in the cointegration equation.

[insert Table 5 here]

So far, the non-augmented model has been shown to explain much of the variation in the dependent market index by significantly increasing the variability of the cointegrating coefficient. The next question that one must necessarily ask is whether the cointegrating coefficients of the non-augmented and augmented models subsume any information about gradual market cointegration or segmentation over time as globalization has significantly developed in recent decades. Further analysis of the evolution of the non-augmented and augmented coefficients can provide insight into the effects of globalization on the gradual convergence of the two coefficients. We refer to the augmented and the non-augmented coefficients as the risk-adjusted and the unadjusted cointegrating coefficients, respectively. As Table 5 also suggests we can use the non-augmented state-space as the best fitting time-varying cointegrating coefficient. Equation (2) presumes that both augmented and non-augmented cointegrating coefficients are unit root processes. Further analysis of the long run co-movement of the $\hat{b}_t$ and the $\hat{\hat{b}}_t$ can shed more light on the effects of globalization.

We know from Table 3 and Figure 3 that the cointegrating coefficients are driven by uncertainty and stock return determinants. Moreover, the risk-adjusted cointegrating coefficient fluctuates substantially in periods of market turmoil compared to the unadjusted coefficients. This finding motivates the question as to whether globalization has driven both cointegrating coefficients to follow a common stochastic trend. Markets can possibly be segmented as a result of a number of phenomena including different technological developments, restrictions on capital flow, etc. The causes of market segmentation form a number of unknown factors for which we do not have inclusive indicators. Globalization can dominate these obstacles by reducing the segmentation causes to a certain number of risk factors. If globalization has led to stronger market cointegration, then the non-augmented cointegrating coefficient has to display a common trend with the risk-adjusted cointegrating coefficient. In Table 6, we check whether both types of processes associated with a certain bivariate relation follow a common trend.

[insert Table 6 here]

A bivariate cointegration analysis of the augmented and the non-augmented cointegrating coefficients reveals that in most cases, the two time-varying cointegrating coefficients follow distinct stochastic trends. Regarding the RTR indices, Table 6 shows that in the case of the bivariate relation of the US with Germany, the two augmented and non-augmented cointegrating coefficients follow common trends. When using NPR indices, the coefficients associated with the relations of the US with Germany and Japan follow common trends. In these limited cases of a common trend, we suspect that the time-varying equilibrium relations of the markets are uniquely driven by a number of common risk factors. In addition, the state-space variables estimated by two different specifications follow two distinct stochastic trends in most of the cases. Thus the best fitting time-varying equilibrium relation (i.e., the non-augmented one) follows a distinct unit root process probably driven by a number of unknown segmentation causes. This result implies that markets can be segmented beyond what can be explained by our known set of risk factors.

7.1. On the impact of globalization

Despite considerable advancement in globalization, Figure 2 suggests that long-run market co-movement is subject to short-term adjustments in equilibrium relations, which can be referred to as market segmentation. The adjustment in the *a posteriori* update of the state-space by the Kalman gain matrix can be referred to as a disruption in the cointegration relation. We cannot conclude that markets become more strongly cointegrated over time because, according to Figure 2, both cointegrating coefficients are subject to repeated adjustments. Instead, we can analyse the evolution of the augmented and non-augmented cointegrating coefficients to draw conclusions on markets segmentation. Table 6 suggests that markets are segmented beyond a unit root process driven by uncertainty and risk factors. This entails that there must be a number of unknown factors other than our uncertainty and financial factors that drive the time-varying cointegrating coefficients. The question is whether the effect of these factors diminish as a result of globalization. If globalization has truly exercised such an effect on markets, then the augmented and non-augmented cointegrating coefficients must have gradually followed a common stochastic trend over time. The results of Table 6 tell us that the augmented and non-augmented cointegrating coefficients do not share a stochastic trend in common over the entire period of our analysis; nevertheless, globalization may have gradually caused the two variables to share a common trend. If the non-augmented cointegration coefficients do not evolve far away from the augmented coefficients, then one can conclude that the stochastic trends of the time-varying cointegrating coefficients are those whose first difference is explained by uncertainty and risk factors. We have to examine the gradual convergence of the non-augmented and augmented cointegrating coefficients to detect whether globalization has led both series to share a unique stochastic trend.

To further explore the evolution of the two cointegrating coefficients, we estimate the following state-space formulation, which allows us to observe the time-varying evolution of the equilibrium relation between the non-augmented and the risk-adjusted cointegrating coefficients. The time varying standard deviation of the u_t term can be informative in that it can reveal how the cointegrating coefficients may evolve to a more stable equilibrium relation. We add a GARCH(1,1) effect to the disturbance term of the state-space to investigate the time-varying evolution of the variance.

$$\hat{b}_t = \alpha + \gamma_t \hat{\hat{b}}_t + w_t \tag{3}$$

$$\gamma_t = \gamma_{t-1} + u_t \text{ with } u_t \sim N(0, \sigma_t^2) \tag{4}$$

$$\hat{\sigma}_t^2 = c_1 + c_2 \hat{\sigma}_{t-1}^2 + c_3 \hat{u}_{t-1}^2 \tag{5}$$

The results of the Kalman filter specification on the regression of the non-augmented cointegrating coefficients on the risk-adjusted coefficient including a GARCH effect on the residuals of the state-space are presented in Figure 4. It shows that the regression slope gradually stabilizes over time, particularly in the period following the global financial crisis. Though the estimated conditional standard deviation of the β_t increases during bearish markets, it seems that this standard deviation bears less variation in the recent economic crises. One potential implication of Figure 4 is that globalization has reduced market segmentation sources to the uncertainty and the risk factors in equation (2). One can then conclude that the non-augmented cointegrating coefficient converges gradually to the same stochastic trend as that of the risk-adjusted coefficient. As confirmed in Figures 2 and 3, one cannot definitively conclude that globalization has led to stronger market cointegration. Rather, as Figure 4 suggests, globalization has restrained market segmentation by leading the evolution of the two cointegrating coefficients to follow an identical stochastic trend.

[insert Figure 4 here]

7.2. Country specific results

The results of Table 4 together with those of Figures 3 and 4 suggest that the influence of the EPU and GPR indices on the stability of the bilateral relations of the US stock exchange with the stock markets of the two non-European G7 countries (in addition to the US) warrants further discussion.

As the results of the NPR indices in Table 4 indicate, the GPR index has the largest influence on the Japanese stock market. Moreover, the results of the estimation based on the RTR indices suggest that the equilibrium relation of the US market with that of Japan is significantly destabilized by EPU in the European Union. This conclusion is also in line with Figures 3 where the US market co-movement with that of Japan shows considerable fluctuation in the long-run relative risk-adjusted return over time. During the crisis in the early 1990s following the Japanese asset price bubble, Japan's market comovement with the US becomes sharply more volatile than what can be attributed to the identified sources of market segmentation. However, as Figure 4 shows, the unidentified causes of segmentation in this relation are considerably reduced over time. The vulnerability of the US-Japan equilibrium could be explained by considering its particular geopolitical situation and also the role of growing China as its rival. The US and Japan have maintained their political and security ties during post-Cold War era as firmly as the raging Cold War era. However, the economic and trade frictions between the US and Japan have remained an issue distinct from national security ties (Curtis, 2001; Urata, 2020). The US-Japan trade policy was not a foreign policy priority until the administration of President Clinton. Japan's limited bargaining power in trade negotiations with the US resulted in moderation of its trade account surplus (Curtis, 2001). The well-known section 232 of the Trade Expansion Act, which allows the US to take protectionism measures in the name of national security, plays a major role in the trade frictions between the two nations. Despite the growing security ties between these two countries, our results indicate that the unique geopolitical situation of Japan puts its equilibrium relation with the US in a position that is vulnerable to both economic and geopolitical uncertainty. However, in the recent decade the globalization has reduced the causes of market segmentation to our identified factors.

Japan's specific geopolitical situation and historical antagonism with China have possibly rendered its stock market comovement with that of the US vulnerable to GPR. Although Japan's trade with China has gradually grown over recent decades, the historical political conflict between the two countries has remained a major obstacle in the mutual relation between the two world powers (Smith, 2009).

China's military expansion in light of the historical enmity between the two nations could have left a destabilizing impact on the long-run US-Japan equilibrium relation. The absence of a trilateral political alliance among the world's three major powers makes uncertainty factors a more significant driver of the dynamics of the co-movement of the two markets.

The relative long-run risk-adjusted rate of return of the US in the US-Canada relation as Figure 3 displays, rises after 2010 such that it peaks in early 2016, whereas that of the other bivariate relations is downward sloping. The economic downturn in the early 1990s has adversely affected the US-Canada equilibrium relation. As Figure 4 displays, the time-varying standard deviation of the state space spikes during this recession, whereas during the economic recession of the early 2000s it is only slightly increased. Moreover, according to Figure 4, the time varying standard deviation of γ_t corresponding to the US-Canada relation substantially decreases in the period following the global financial crisis.

These findings are surprising because despite Canada's geographical proximity to the US, one can hardly expect such an unsteady equilibrium relation. Given that Canada's economy has greater dependence on foreign trade with the US than vice versa, where almost 75% of Canadian exports are destined to the US, its economy can be more vulnerable to uncertainty variables, as documented in the literature (see Smith et al. (2018)). The substantial reliance of the Canada's economy on that of the US puts the Canadian economy in a more precarious position, making the bivariate equilibrium relation with the US market more vulnerable to both economic and geopolitical uncertainty.

Despite the specific instabilities in the bivariate relation of the Canadian and US markets, as Figure 4 shows, γ_t has substantially stabilized in the last decade. This finding is also in line with the specific bivariate equilibrium relation with Japan according to which γ_t has also stabilized over time. As indicated above, a potential explanation would be the positive impact of the globalization on reducing the unknown causes of market segmentation.

8. Conclusions

Rapidly growing globalization suggests stronger interconnectedness of world stock markets whereas the stock markets of the G7 member countries do not exhibit a time-invariant equilibrium relation. Although the theoretical developments in time-varying cointegration are quite mature, the causes of time-varying cointegration among stock markets remain unexplored. The literature theoretically and empirically supports the affective role of uncertainty in the economy as well as in financial markets. Time-varying cointegration could be a consequence of repeated shocks to the uncertainty prompted by economic policy makers or by geopolitical risks. As the literature has provided evidence that uncertainty affects the evolution of stock markets, we investigated the hypothesis that a shock to uncertainty is likely to distort the co-movement of markets. Our empirical findings imply that the time-varying nature of the long-run equilibrium relation is driven by changing economic policy uncertainty and geopolitical risks and by a number of stock return determinants.

Since the disruptions of the augmented cointegrating coefficient are driven by uncertainty and financial risk determinants we refer to the augmented state-space as the risk-adjusted cointegrating coefficient. The results of the estimation indicate that the US market has a higher (lower) long-run risk-adjusted return that the other G7 countries in periods of economic growth (recession). This finding implies that having higher exposure to certain risk factors leads to market segmentation. Moreover, our results confirm that the unknown factors driving markets out of cointegration have been dominated as a results of globalization.

The bivariate cointegration of the US market with two non-European G7 markets, namely, those of Japan and Canada, has, more than other relations, been subject to segmentation by the underlying variables. The specific geopolitical situation of Japan is in compliance with our empirical findings whereas the fragile equilibrium of the US and Canada is a surprising finding that must be closely scrutinized in future works.

Declarations of interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper**.**

References

Aladesanmia, O., et al. (2019). "Stock market integration between the UK and the US: Evidence over eight decades." Global Finance Journal 41: 32-43.

Antonakakis, N., et al. (2017). "Geopolitical risks and the oil-stock nexus over 1899–2016." Finance Research Letters 23: 165-173.

Arshanapalli, B. and J. Doukas (1993). "International stock market linkages: Evidence from the pre- and post-October 1987 period." Journal of Banking & Finance 17(1): 193-208.

Baker, S. R., et al. (2016). "Measuring Economic Policy Uncertainty*." The Quarterly Journal of Economics 131(4): 1593-1636.

Balcilar, M., et al. (2018). "Geopolitical risks and stock market dynamics of the BRICS." Economic Systems 42(2): 295-306.

Bernanke, B. S. (1983). "Irreversibility, uncertainty and cyclical investment." Quarterly Journal of Economics 98(1): 85-106.

Bierens, H. J. and L. F. Martins (2010). "Time-varying cointegration." Econometric Theory 26(5): 1453-1490.

Bloom, N. (2009). "The impact of uncertainty shocks." Econometrica 77(3): 623–685.

Bloom, N., et al. (2018). "Really Uncertain Business Cycles." Econometrica 86(3): 1031-1065.

Brogaard, J. and A. Detzel (2015). "The Asset-Pricing Implications of Government Economic Policy Uncertainty." Management Science 61(1): 3–18.

Caballero, R. J. and R. S. Pindyck (1996). "Uncertainty, investment, and industry evolution." International Economic Review 37: 641–662.

Caldara, D. and M. Iacoviello (2021). "Measuring geopolitical risk." Working Paper, Board of the Governors of the Federal Reserve Board.

Cheng, C. H. J. and C.-W. J. Chiu (2018). "How important are global geopolitical risks to emerging countries?" International Economics 156: 305-325.

Curtis, G. (2001). "New Perspectives on U.S.-Japan Relations." Japan Center for International Exchange.

Das, D., et al. (2019). "Do the emerging stock markets react to international economic policy uncertainty, geopolitical risk and financial stress alike?" The North American Journal of Economics and Finance 48: 1-19.

Engle, R. F. and C. W. J. Granger (1987). "Co-Integration and Error Correction: Representation, Estimation, and Testing." Econometrica 55(2): 251-276.

Gardner, L. A. (1969). "On Detecting Changes in the Mean of Normal Variates." The Annals of Mathematical Statistics 40(1): 116-126.

Gilmore, C. G., et al. (2008). "The dynamics of Central European equity market comovements." The Quarterly Review of Economics and Finance 48(3): 605-622.

Goldfeld, S. M. and R. E. Quandt (1973). "A Markov model for switching regressions." Journal of Econometrics 1(1): 3-15.

Gregory, A., W. and B. E. Hansen (1996). "Residual-based tests for cointegration in models with regime shifts." Journal of Econometrics 70(1): 99-126.

Haldane, A. G. and S. G. Hall (1991). "Sterling's Relationship with the Dollar and the Deutschemark: 1976-89." The Economic Journal 101(406): 436-443.

Hamilton, J. D. (1989). "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle." Econometrica 57(2): 357-384.

Hansen, B. E. (1992). "Tests for Parameter Instability in Regressions with I(1) Processes." Journal of Business & Economic Statistics 20(1): 45-59.

Hansen, H. and S. Johansen (1999). "Some tests for parameter constancy in cointegrated VAR-models." The Econometrics Journal 2(2): 306-333.

Hansen, P. R. (2003). "Structural changes in the cointegrated vector autoregressive model." Journal of Econometrics 114(2): 261-295.

He, L., et al. (2021). "Institutional conditions, economic policy uncertainty and foreign institutional investment in China." Emerging Markets Review: 100823.

Johansen, S. (1988). "Statistical analysis of cointegration vectors." Journal of Economic Dynamics and Control 12(2-3): 231-254.

Kanas, A. (1998). "Linkages between the US and European equity markets: further evidence from cointegration tests." Applied Financial Economics 8(6): 607-614.

Kannadhasan, M. and D. Das (2020). "Do Asian emerging stock markets react to international economic policy uncertainty and geopolitical risk alike? A quantile regression approach." Finance Research Letters 34: 101276.

Kejriwal, M. and P. Perron (2010). "Testing for Multiple Structural Changes in Cointegrated Regression Models." Journal of Business & Economic Statistics 28(4): 503-522.

Le, A.-T. and T. P. Tran (2021). "Does geopolitical risk matter for corporate investment? Evidence from emerging countries in Asia." Journal of Multinational Financial Management: Article in press.

Lence, S. and B. Falk (2005). "Cointegration, market integration, and market efficiency." Journal of International Money and Finance 24(6): 873-890.

Li, D., et al. (2020). "Kernel-based Inference in Time-Varying Coefficient Cointegrating Regression." Journal of Econometrics 215(2): 607-632.

Liow, K. H., et al. (2018). "Dynamics of international spillovers and interaction: Evidence from financial market stress and economic policy uncertainty." Economic Modelling 68: 96-116.

Ludvigson, S. C., et al. (2021). "Uncertainty and Business Cycles: Exogenous Impulse or Endogenous Response?" American Economic Journal: Macroeconomics 13(4): 369-410.

Ma, R., et al. (2019). "The forecasting power of EPU for crude oil return volatility." Energy Reports 5: 866-873.

Menezes, R., et al. (2012). "On the globalization of stock markets: An application of Vector Error Correction Model, Mutual Information and Singular Spectrum Analysis to the G7 countries." The Quarterly Review of Economics and Finance 52(4): 369-384.

Mylonidis, N. and C. Kollias (2010). "Dynamic European stock market convergence: Evidence from rolling cointegration analysis in the first euro-decade." Journal of Banking & Finance 34(9): 2056-2064.

Nguyen, C. P., et al. (2020). "Economic policy uncertainty and credit growth: Evidence from a global sample." Research in International Business and Finance 51: 101118.

Nyblom, J. (1989). "Testing for the Constancy of Parameters Over Time." Journal of the American Statistical Association 84(405): 223-230.

Omar, A. M. A., et al. (2017). "Diversifying away the risk of war and cross-border political crisis." Energy Economics 64: 494–510.

Park, J. Y. and S. B. Hahn (1999). "Cointegrating Regressions with Time Varying Coefficients." Econometric Theory 15(5): 664-703.

Pascual, A. G. (2003). "Assessing European stock markets (co)integration." Economics Letters 78(2): 197-203.

Phillips, P. C. B. and B. E. Hansen (1990). "Statistical Inference in Instrumental Variables Regression with I(1) Processes." The Review of Economic Studies 57(1): 99-

125.

Ploberger, W., et al. (1989). "A new test for structural stability in the linear regression model." Journal of Econometrics 40(2): 307-318.

Prüser, J. and A. Schlösser (2020). "On the time-varying effects of economic policy uncertainty on the US economy." Oxford bulletin of economics and statistics 82(5): 0305–9049.

Quandt, R. E. (1960). "Tests of the Hypothesis that a Linear Regression System Obeys Two Separate Regimes." Journal of the American Statistical Association 55(290): 324-330.

Quintos, E. C. and P. C. B. Phillips (1993). "Parameter constancy in cointegrating regressions." Empirical Economics 18: pages675–706.

Rangvid, J. (2001). "Increasing convergence among European stock markets?: A recursive common stochastic trends analysis." Economics Letters 71(3): 383-389.

Richards, A. J. (1995). "Comovements in national stock market returns: Evidence of predictability, but not cointegration." Journal of Monetary Economics 36(3): 631-654.

Serletis, A. and M. King (1997). "Common Stochastic Trends and Convergence of European Union Stock Markets." The Manchester School 65(1): 44-57.

Smith, M. J., et al. (2018). "Long-term lessons on the effects of post-9/11 border thickening on cross-border trade between Canada and the United States: A systematic review." Transport Policy 72: 198-207.

Smith, P. J. (2009). "China-Japan Relations and the Future Geopolitics of East Asia." Asian Affairs: An American Review 35(4): 230-256.

Suh, H. and J. Y. Yang (2021). "Global uncertainty and Global Economic Policy Uncertainty: Different implications for firm investment." Economics Letters 200: 109767.

Taylor, M. P. and I. Tonks (1989). "The Internationalisation of Stock Markets and the Abolition of U.K. Exchange Control." The Review of Economics and Statistics 71(2): 332-336.

URATA, S. (2020). "US–Japan Trade Frictions: The Past, the Present, and Implications for the US–China Trade War." Asian Economic Policy Review 15(1): 141–159.

Yu, X. and Y. Huang (2021). "The impact of economic policy uncertainty on stock volatility: Evidence from GARCH–MIDAS approach." Physica A: Statistical Mechanics and its Applications 570: 125794.

Zhou, K., et al. (2021). "The economic policy uncertainty and the choice of entry mode of outward foreign direct investment: Cross-border M&A or Greenfield Investment." Journal of Asian Economics 74: 101306.

Zhou, Z., et al. (2020,). "Can economic policy uncertainty predict exchange rate volatility? New evidence from the GARCH-MIDAS model." Finance Research Letters 34: 101258.

Table 1. Unit root tests of the indices transformed by logarithms

	RTR indices				NPR indices			
	ADF		PP		ADF		PP	
	Test statistic	Lag	Test statistic	Bandwidth	Test statistic	Lag	Test statistic	bandwidth
US	-0.58	(0)	-0.79	(5)	-0.71	(0)	-0.72	(5)
UK	-1.75	(0)	-1.77	(4)	-2.03	(0)	-2.07	(5)
France	-2.03	(0)	-2.03	(4)	-2.57	(0)	-2.57	(4)
Germany	-1.79	(0)	-1.80	(0)	-2.03	(0)	-2.02	(2)
Italy	-2.18	(0)	-2.19	(7)	-2.29	(0)	-2.30	(4)
Canada	-1.61	(0)	-1.52	(8)	-1.63	(0)	-1.55	(8)
Japan	-1.59	(0)	-1.83	(6)	-2.32	(0)	-2.56	(6)

All indices are in terms of US dollars and transformed by natural logarithms. A unit root test is performed using the augmented Dickey-Fuller (ADF) and Phillip-Perron (PP) test methods. *, ** and *** represent significance at the 10%, 5% and 1% levels, respectively. The numbers in parentheses represent the lag length of the ADF test and the bandwidth of the PP test. ADF: the lag length is automatically specified using the Schwarz criterion. PP: The Bartlett kernel is used and the bandwidth is specified using the Newey-West method.

Table 2. Bi-variate time invariant cointegration test

	US-UK		**US-France**		**US-Germany**		**US-Italy**		**US-Canada**		**US-Japan**	
	RTR indices											
Dependent	US	UK	US	France	US	Germany	US	Italy	US	Canada	US	Japan
tau-statistic	-0.84	-1.53	-2.14	-2.74	-2.54	-2.99	-1.27	-2.45	-0.78	-1.69	-1.95	-1.95
	(0.93)	(0.75)	(0.46)	(0.19)	(0.26)	(0.12)	(0.84)	(0.30)	(0.93)	(0.68)	(0.68)	(0.56)
z-statistic	-3.07	-6.23	-9.16	-12.35	-14.02	-17.35	-3.55	-11.94	-1.33	-5.67	-2.29	-8.29
	(0.87)	(0.63)	(0.41)	(0.24)	(0.18)	(0.09)	(0.84)	(0.26)	(0.96)	(0.68)	(0.92)	(0.47)
	NPR indices											
Dependent	US	UK	US	France	US	Germany	US	Italy	US	Canada	US	Japan
tau-statistic	-0.82	-1.70	-1.95	-2.88	-2.14	-2.82	0.05	-2.07	-0.89	-1.61	-0.53	-2.27
	(0.93)	(0.68)	(0.56)	(0.14)	(0.46)	(0.16)	(0.98)	(0.49)	(0.92)	(0.72)	(0.96)	(0.39)
z-statistic	-3.06	-7.28	-7.91	-12.87	-10.50	-15.01	0.07	-8.64	-1.63	-5.16	-0.84	-11.33
	(0.87)	(0.55)	(0.50)	(0.22)	(0.33)	(0.14)	(0.98)	(0.45)	(0.95)	(0.72)	(0.97)	(0.29)

All indices are in terms of US dollars and transformed by natural logarithms. Engle-Granger cointegration is applied to examine bivariate cointegration relations. The numbers in parentheses are p-values. Each bivariate cointegration relation is examined twice depending on the choice of dependent variable.

Table 3. Hansen stability analysis of bivariate cointegration of the US and other G7 members

	UK	France	Germany	Italy	Canada	Japan
			RTR indices			
L_c	0.53	0.19	0.21	0.42	0.04	0.35
	(0.03)	(>0.2)	(>0.2)	(0.07)	(>0.2)	(0.10)
$MeanF$	12.69	5.09	6.40	5.75	0.92	3.55
	(0.01)	(0.04)	(0.01)	(0.02)	(>0.2)	(0.12)
$SupF$	41.10	16.34	16.22	19.50	1.68	12.46
	(0.01)	(0.01)	(0.01)	(0.01)	(>0.2)	(0.06)
			NPR indices			
L_c	0.76	0.27	0.17	139.76	0.05	0.36
	(0.01)	(0.18)	(>0.2)	(0.01)	(>0.2)	(0.10)
$MeanF$	16.20	7.60	5.50	3396.71	1.27	3.75
	(0.01)	(0.01)	(0.03)	(0.01)	(>0.2)	(0.09)
$SupF$	57.45	20.00	19.32	7419.32	2.5	17.66
	(0.01)	(0.01)	(0.01)	(0.01)	(>0.2)	(0.01)

All indices are in terms of US dollars and transformed by natural logarithms. The Hansen test is applied to the bivariate cointegration relations between the US and other G7 member countries. The null hypothesis is a cointegration with constant parameters against the alternative of a random walk in the cointegrating coefficient. The numbers in parentheses are p-values. P-values of more than 0.2 are indicated by 0.2. The k_1 and k_2 parameters are both taken to be zero. The null hypothesis of the three tests is constant cointegration. The alternative hypothesis corresponding to L_c and $MeanF$ is a random walk in the cointegrating coefficient. The alternative hypothesis corresponding to $SupF$ is an unknown structural break in the cointegration relation.

Table 4. Coefficients of the state-space regressors in the cointegration of the US market with 6 other developed markets

	UK	France	Germany	Italy	Canada	Japan
			RTR indices			
a	2.600***	3.8121***	5.7604***	5.7317***	7.0198***	6.8968***
EPU_0^{US}	0,0001	-0,0002	-0,0011	-0.0008	-0,0015*	-0,0005
EPU_0^{EU}	0,0002	0,0002	-0,0002	0.0000	-0,0016*	-0,0038*
GPR_0	-0,0010***	-0,0006*	-0,0009**	-0.0009**	-0,0010*	-0,0046***
EPU_1^{US}	-0,0002	-0,0005	-0,0006	-0.0009	-0,0010	-0,0006
EPU_1^{EU}	0,0000	0,0000	0,0005	0.0005	-0,0002	-0,0004
GPR_1	-0,0002	-0,0005	-0,0003	-0.0009*	-0,0003	0,0001
BEER	-0,0003	-0,0004	-0,0002*	-0.0005	-0,0003	-0,0003
Credit spread	-0,0288	-0,0603**	-0,1090**	-0.0601*	-0,1189***	-0,2923***
DIVY	0,0251***	0.0293***	0.0345***	0.0425***	0.0488***	0.1081**
DVIX	-0.0002***	-0.0003***	-0.0005***	-0.0005***	-0.0006***	-0.0016***
US inflation	-0.0004	0.0000	-0.0003	-0.0014	-0.0002	0.0012
Domestic inflation	-0.0001	0.0000	-0.0003	0.0010	-0.0002	0.0015
US Term spread	-0.0021	-0.0015	-0.0053	-0.0052	-0.0178	-0.0335
Domestic term	-0.0036	0.0068	0.0180	-0.0060	-0.0246	0.0231
SLL	**844.44**	**838.42**	**774.74**	**782.41**	**746.72**	**747.44**
ADF	**-18.77*****	**-18.73*****	**-18.80*****	**-18.81*****	**-18.92*****	**-17.76*****
			NPR indices			
a	2.1351***	3.7079***	4.4189***	5.3035***	6.3087***	6.1743***
EPU_0^{US}	0.0000	-0.0002	-0.0008	-0.0009	-0.0017*	-0.0008
EPU_0^{EU}	0.0005	0.0003	0.0007	-0.0003	-0.0016*	-0.0033
GPR_0	-0.0010**	-0.0007*	-0.0005	-0.0010*	-0.0009*	-0.0041***
EPU_1^{US}	-0.0004	-0.0007	-0.0011	-0.0014*	-0.0012	-0.0017
EPU_1^{EU}	0.0000	0.0000	0.0002	0.0007	-0.0002	-0.0003
GPR_1	-0.0002	-0.0006	-0.0004	-0.0010*	-0.0001	0.0006
BEER	-0.0000	-0.0004	0.0006	-0.0006	0.0007*	0.0035***
Credit spread	-0.0297	-0.0824**	-0.1014**	-0.0789*	-0.1924***	-0.5580***
DIVY	0.0263***	0.0345***	0.0387***	0.0479***	0.0596***	0.1872***
DVIX	-0.0002***	-0.0030***	-0.0004***	0.0006***	-0.0007***	-0.0017***
US inflation	-0.0005	-0.0005	-0.0008	-0.0008	-0.00320	-0.0119**
Domestic inflation	-0.0007	0.0140	-0.0008	0.0016*	0.0003	0.0020
US Term spread	-0.0065	0.0011	-0.0102	-0.0101	-0.01250	-0.0825
Domestic term	-0.0009	-0.0001	0.0070	-0.0009	-0.0010	-0.1203
SLL	**844.92**	**822.02**	**818.27**	**785.52**	**744.80**	**753.84**
ADF	**-18.93*****	**-18.77*****	**-18.78*****	**-18.86*****	**-18.80*****	**-18.04*****

All indices are in terms of US dollars and transformed by natural logarithms. Each column corresponds to the bivariate cointegration relation of the US with G7 member countries. [*], [**] and [***] represent significance at 10%, 5% and 1% levels, respectively. The numbers in parentheses are p-values. The parameters $\gamma_{US,0}$ and $\gamma_{EU,0}$ are the response of the change in the cointegrating coefficient to contemporaneous innovations in EPU in the US and European

Union respectively. $\gamma_{GPR,0}$ represents the response of the change in the cointegrating coefficient to contemporaneous innovations in the GPR index. SLL stands for the maximized sum of the log likelihood function. ADF represents the augmented Dickey-Fuller test statistic of the residuals of the cointegration relation.

Table 5. Variance equality test of the residuals

	UK	France	Germany	Italy	Canada	Japan
Variance term			**RTR indices**			
σ_{ε}	1.4385 (0.00)	1.6875 (0.00)	4.8142 (0.00)	3.3181 (0.00)	2.7636 (0.06)	1639907.514 (0.00)
σ_{v}	0.7277 (0.00)	0.4983 (0.00)	0.5463 (0.00)	0.3987 (0.00)	0.4978 (0.00)	0.4507 (0.00)
$Y_t - \bar{Y}_t$	1.3205 (0.00)	1.4684 (0.00)	3.5840 (0.00)	2.3489 (0.00)	1.0543 (0.30)	6.5361 (0.00)
Variance term			**NPR indices**			
σ_{ε}	1.7362 (0.00)	2.3456 (0.00)	2.5085 (0.00)	3.1741 (0.00)	2.9182 (0.00)	1848176.426 (0.00)
σ_{v}	0.7625 (0.00)	0.6796 (0.00)	0.71378 (0.00)	0.5521 (0.00)	0.5289 (0.00)	0.4607 (0.00)
$Y_t - \bar{Y}_t$	1.5745 (0. 00)	2.0222 (0. 00)	1.9912 (0. 00)	2.0112 (0. 00)	1.1157 (0.15)	7.3189 (0. 00)

The error terms of the augmented model and that of the non-augmented model are tested for equality of variance. The F statistic is obtained by dividing the estimated variance in the error terms of the augmented bivariate cointegration model to those of the non-augmented model. Each column corresponds to the bivariate cointegration relation of the US with a G7 member country. The numbers in parentheses indicate p-values. The p-values of the first and third rows are right-tail values, and those of the second row are left-tail values.

Table 6. Cointegration test of augmented and non-augmented cointegrating coefficients

	UK		France		Germany		Italy		Canada		Japan	
	RTR indices											
Dependent	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$
tau-statistic	-2.28	-2.35	-2.39	-2.37	-3.20	-3.17	-2.56	-2.54	-1.67	-1.70	-2.74	-2.68
	(0.38)	(0.35)	(0.33)	(0.34)	(0.07)	(0.08)	(0.26)	(0.26)	(0.69)	(0.68)	(0.19)	(0.21)
z-statistic	-10.20	-10.41	-11.36	-11.11	-19.87	-19.57	-12.73	-12.65	-5.71	-5.82	-15.22	-14.86
	(0.35)	(0.34)	(0.28)	(0.30)	(0.05)	(0.06)	(0.22)	(0.23)	(0.67)	(0.66)	(0.14)	(0.15)
	NPR indices											
Dependent	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$	$\hat{b}_t$	$\hat{\hat{b}}_t$
tau-statistic	-2.38	-2.44	-2.62	-2.65	-3.25	-3.25	-2.18	-2.19	-1.67	-1.66	-3.24	-3.20
	(0.34)	(0.31)	(0.23)	(0.22)	(0.06)	(0.06)	(0.44)	(0.43)	(0.69)	(0.70)	(0.07)	(0.07)
z-statistic	-11.29	-11.52	-13.26	-13.32	-21.17	-21.09	-9.65	-9.71	-5.60	-5.53	-21.07	-20.78
	(0.29)	(0.28)	(0.20)	(0.20)	(0.04)	(0.04)	(0.38)	(0.38)	(0.68)	(0.69)	(0.04)	(0.04)

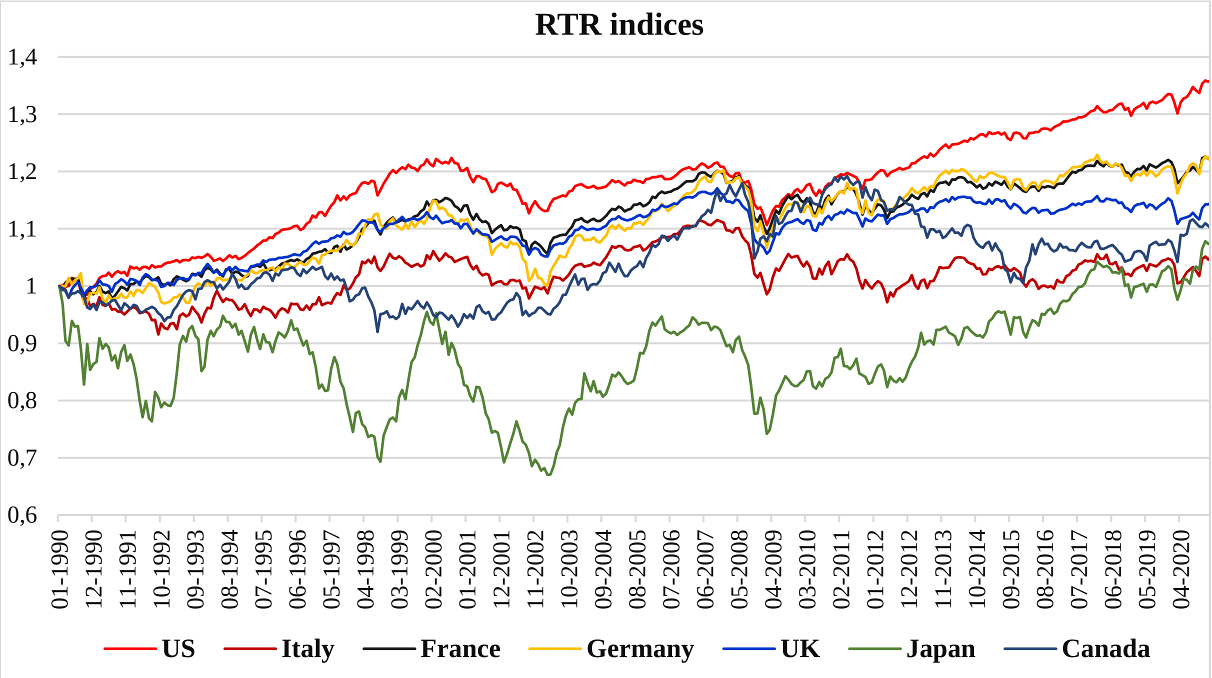
Engle-Granger cointegration is applied to examine the bivariate cointegration relations between augmented and non-augmented cointegrating coefficients. $\hat{b}_t$ represents the non-augmented cointegrating coefficient and $\hat{\hat{b}}_t$ represents the augmented cointegrating coefficient. The numbers in parentheses are p-values. Each bivariate cointegration relation is examined twice depending on the choice of dependent variable.

Figure 1. RTR and nominal NPR indices for G7 countries. All indices are in terms of US dollars and are transformed by natural logarithms. The first sub-graph exhibits inflation-adjusted total return indices. The second sub-graph exhibits nominal price return indices.

Figure 2. Cointegrating coefficient of the bivariate relations between the US and each of the G7 members. The solid lines represent the results of the augmented model. The dashed lines represent the results of the non-augmented model. All indices are in terms of US dollars and are transformed by natural logarithms. The first sub-graph exhibits the results estimated using inflation-adjusted total return indices. The second sub-graph exhibits the results estimated using nominal price return indices.

Figure 3. The discrepancy between the non-augmented and augmented cointegrating coefficients. Discrepancy is defined as $\hat{\hat{b}}_t - \hat{b}_t$, where $\hat{\hat{b}}_t$ and $\hat{b}_t$ represent the augmented and the non-augmented state-space variables, respectively. The first sub-graph exhibits the results estimated using inflation-adjusted total return indices. The second sub-graph exhibits the results estimated using nominal price return indices.

Figure 4. The time varying cointegrating coefficients of the non-augmented and augmented models. The non-augmented cointegrating coefficient is regressed on a constant and the augmented cointegrating coefficient, whereas the slope follows a random walk. The first sub-graph exhibits the results estimated using inflation-adjusted total return indices for Germany, Canada and Japan. The second sub-graph exhibits the results using the same indices for the UK, France and Italy.


RTR indices
1,4
1,3
1,2
1,1
1
0,9
0,8
0,7
0,6
01-1990
12-1990
11-1991
10-1992
09-1993
08-1994
07-1995
06-1996
05-1997
04-1998
03-1999
02-2000
01-2001
12-2001
11-2002
10-2003
09-2004
08-2005
07-2006
06-2007
05-2008
04-2009
03-2010
02-2011
01-2012
12-2012
11-2013
10-2014
09-2015
08-2016
07-2017
06-2018
05-2019
04-2020
US
Italy
France
Germany
UK
Japan
Canada

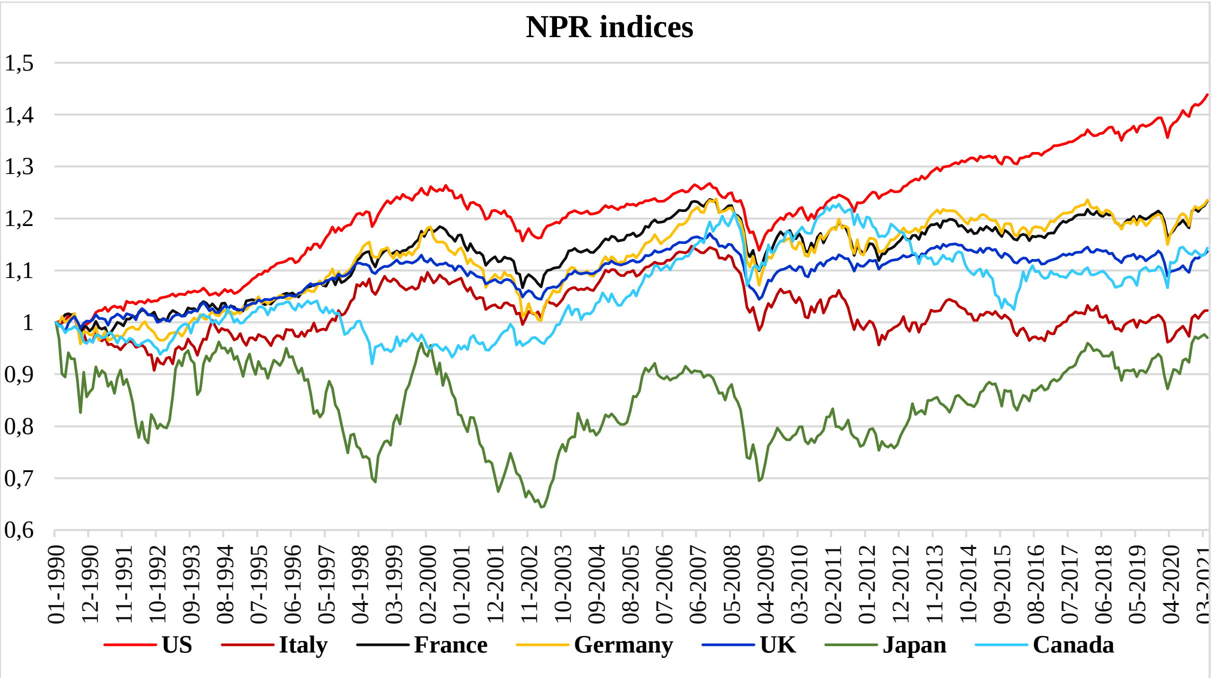
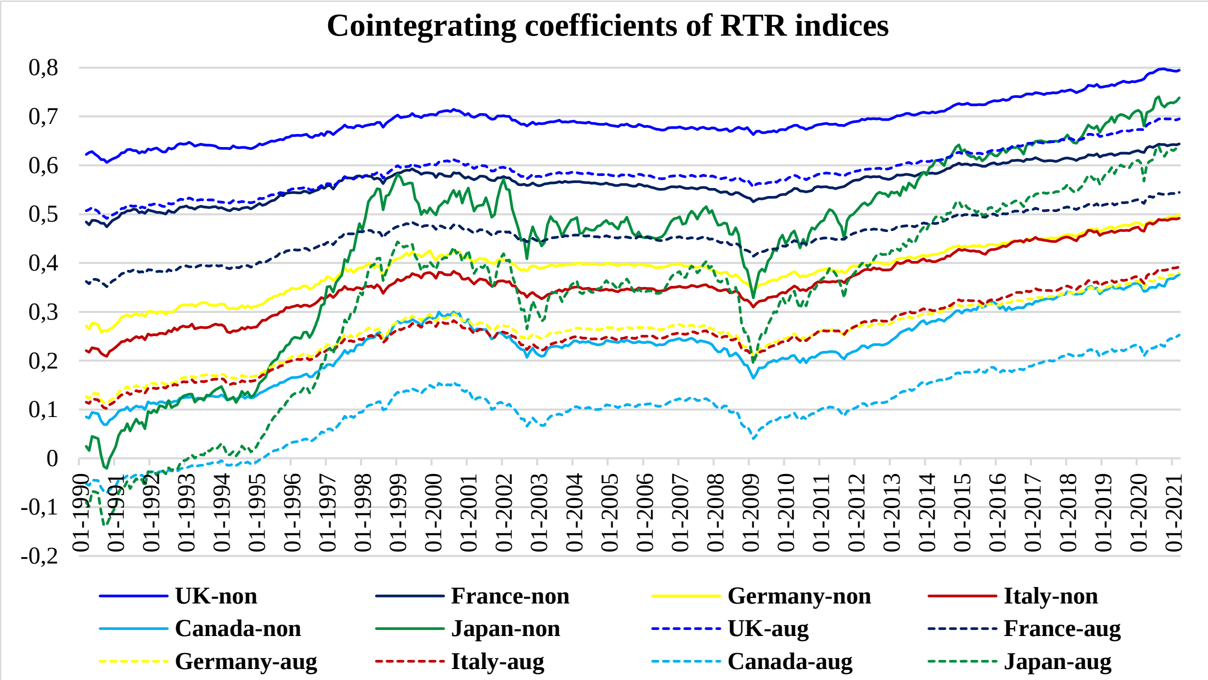

NPR indices
1,5
1,4
1,3
1,2
1,1
1
0,9
0,8
0,7
0,6
01-1990
12-1990
11-1991
10-1992
09-1993
08-1994
07-1995
06-1996
05-1997
04-1998
03-1999
02-2000
01-2001
12-2001
11-2002
10-2003
09-2004
08-2005
07-2006
06-2007
05-2008
04-2009
03-2010
02-2011
01-2012
12-2012
11-2013
10-2014
09-2015
08-2016
07-2017
06-2018
05-2019
04-2020
03-2021
US
Italy
France
Germany
UK
Japan
Canada

Figure 1


Cointegrating coefficients of RTR indices
0,8
0,7
0,6
0,5
0,4
0,3
0,2
0,1
0
-0,1
-0,2
01-1990
01-1991
01-1992
01-1993
01-1994
01-1995
01-1996
01-1997
01-1998
01-1999
01-2000
01-2001
01-2002
01-2003
01-2004
01-2005
01-2006
01-2007
01-2008
01-2009
01-2010
01-2011
01-2012
01-2013
01-2014
01-2015
01-2016
01-2017
01-2018
01-2019
01-2020
01-2021
UK-non
France-non
Germany-non
Italy-non
Canada-non
Japan-non
UK-aug
France-aug
Germany-aug
Italy-aug
Canada-aug
Japan-aug

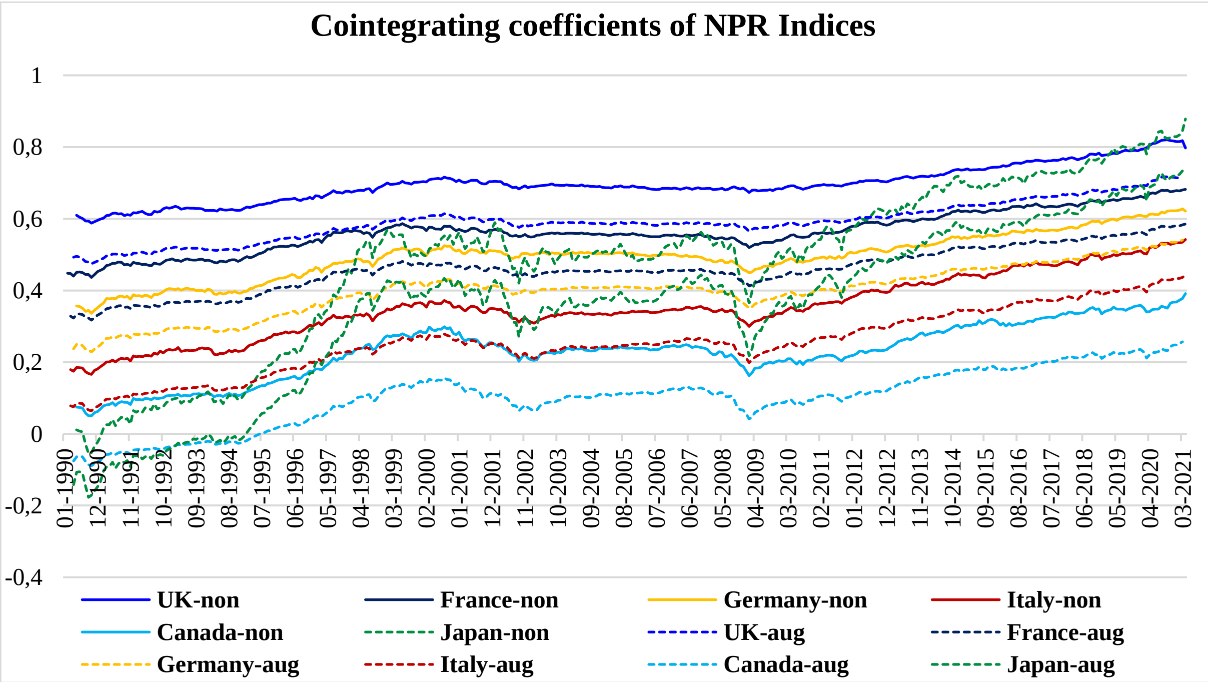
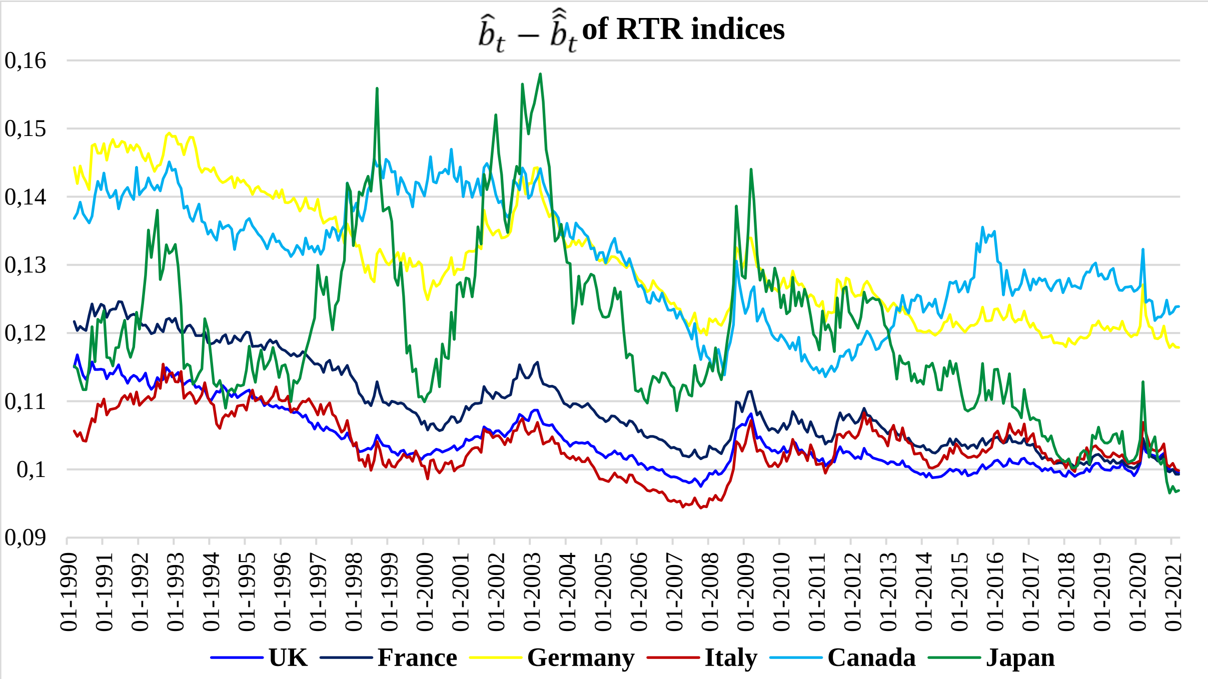

Cointegrating coefficients of NPR Indices
1
0,8
0,6
0,4
0,2
0
-0,2
-0,4
01-1990
12-1990
11-1991
10-1992
09-1993
08-1994
07-1995
06-1996
05-1997
04-1998
03-1999
02-2000
01-2001
12-2001
11-2002
10-2003
09-2004
08-2005
07-2006
06-2007
05-2008
04-2009
03-2010
02-2011
01-2012
12-2012
11-2013
10-2014
09-2015
08-2016
07-2017
06-2018
05-2019
04-2020
03-2021
UK-non
France-non
Germany-non
Italy-non
Canada-non
Japan-non
UK-aug
France-aug
Germany-aug
Italy-aug
Canada-aug
Japan-aug

Figure 2


$\hat{b}_t - \hat{\hat{b}}_t$ of RTR indices
0,16
0,15
0,14
0,13
0,12
0,11
0,1
0,09
01-1990 01-1991 01-1992 01-1993 01-1994 01-1995 01-1996 01-1997 01-1998 01-1999 01-2000 01-2001 01-2002 01-2003 01-2004 01-2005 01-2006 01-2007 01-2008 01-2009 01-2010 01-2011 01-2012 01-2013 01-2014 01-2015 01-2016 01-2017 01-2018 01-2019 01-2020 01-2021
UK
France
Germany
Italy
Canada
Japan

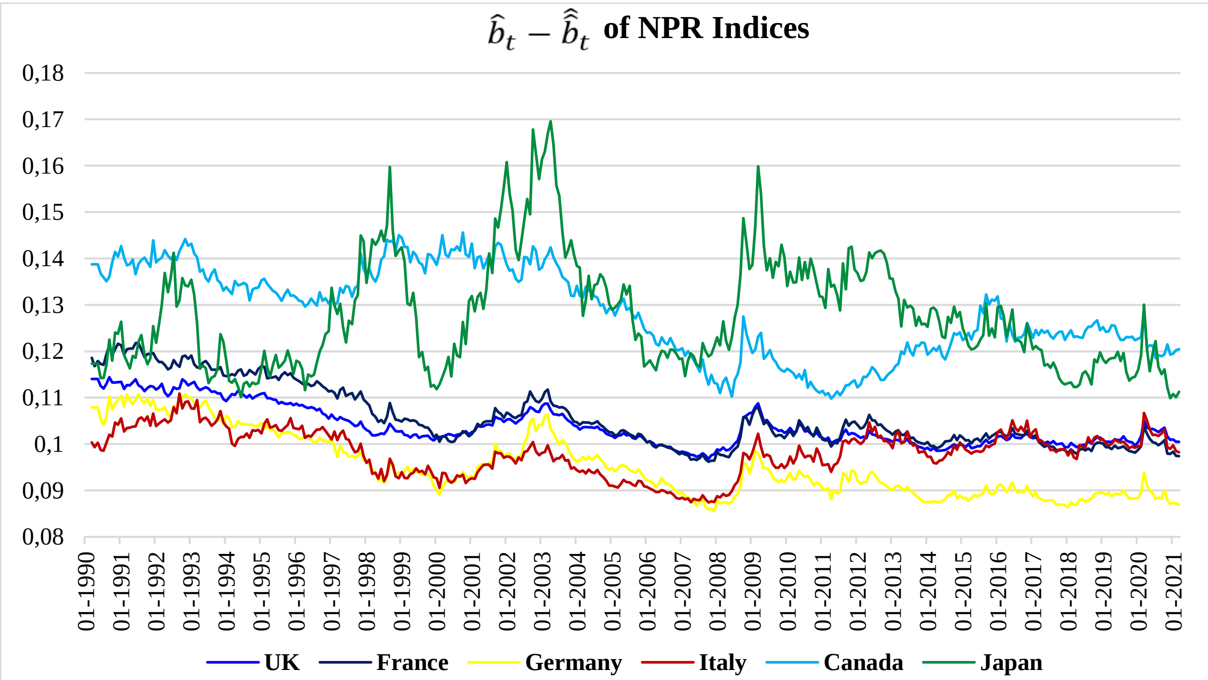
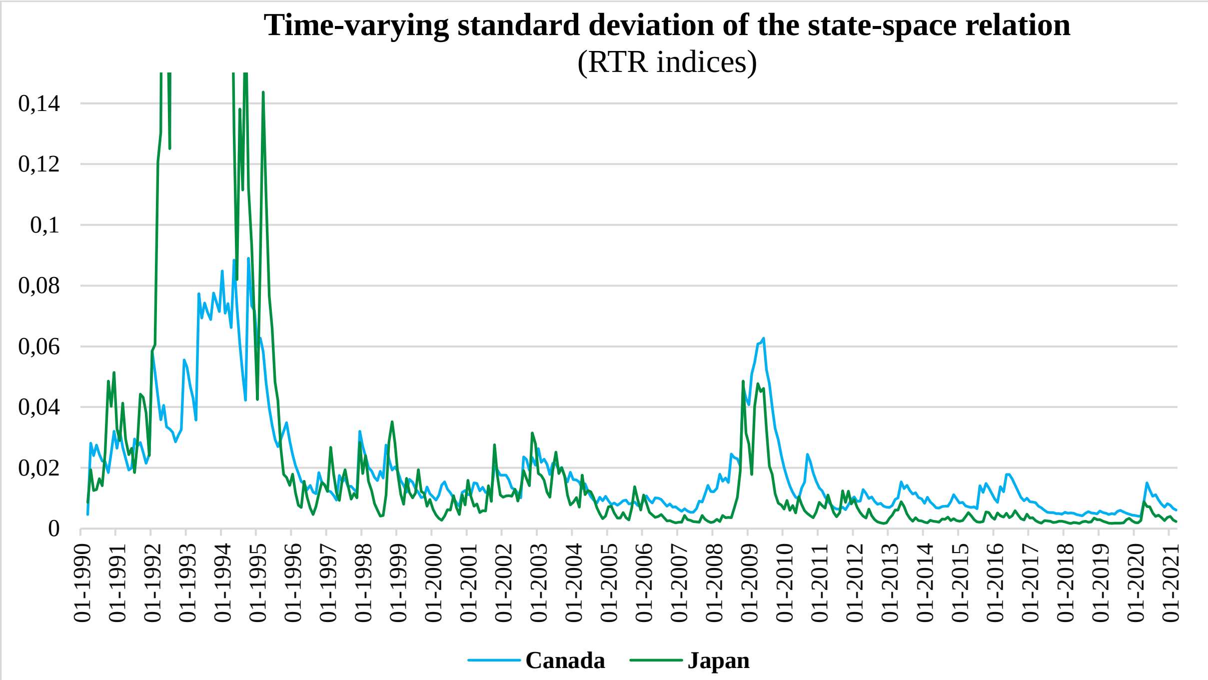

$\hat{b}_t - \hat{\hat{b}}_t$ of NPR Indices
0,18
0,17
0,16
0,15
0,14
0,13
0,12
0,11
0,1
0,09
0,08
01-1990 01-1991 01-1992 01-1993 01-1994 01-1995 01-1996 01-1997 01-1998 01-1999 01-2000 01-2001 01-2002 01-2003 01-2004 01-2005 01-2006 01-2007 01-2008 01-2009 01-2010 01-2011 01-2012 01-2013 01-2014 01-2015 01-2016 01-2017 01-2018 01-2019 01-2020 01-2021
UK
France
Germany
Italy
Canada
Japan

Figure 3


Time-varying standard deviation of the state-space relation
(RTR indices)
0,14
0,12
0,1
0,08
0,06
0,04
0,02
0
01-1990
01-1991
01-1992
01-1993
01-1994
01-1995
01-1996
01-1997
01-1998
01-1999
01-2000
01-2001
01-2002
01-2003
01-2004
01-2005
01-2006
01-2007
01-2008
01-2009
01-2010
01-2011
01-2012
01-2013
01-2014
01-2015
01-2016
01-2017
01-2018
01-2019
01-2020
01-2021
Canada
Japan

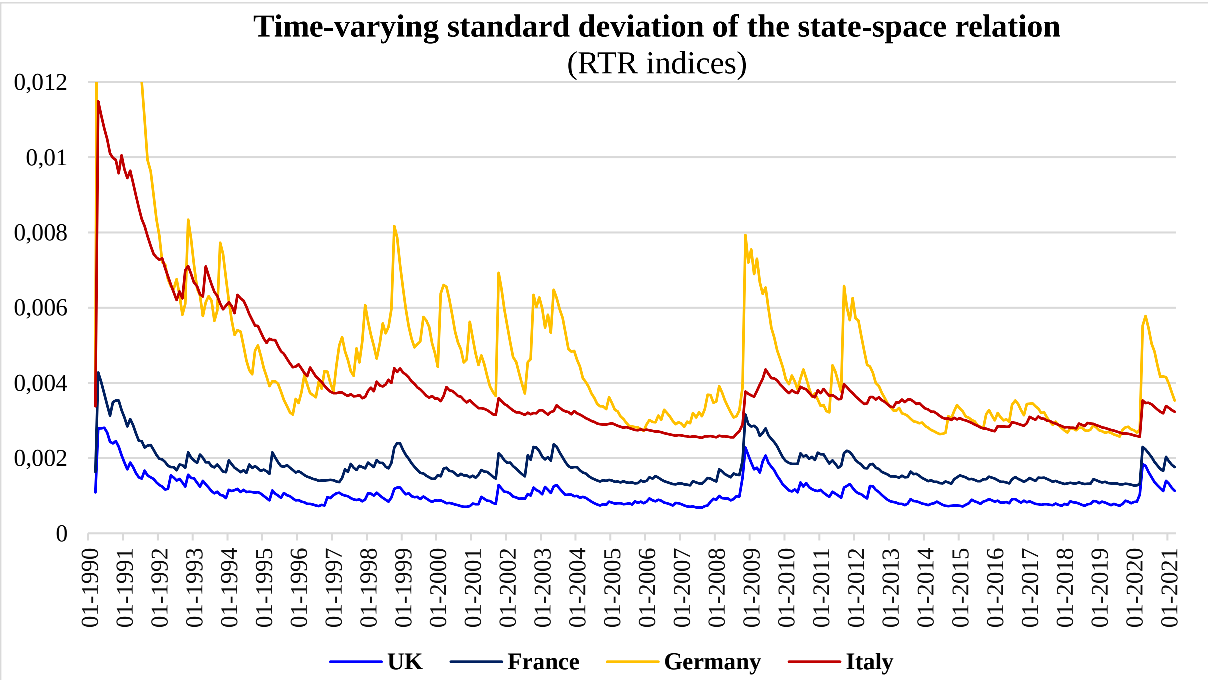

Time-varying standard deviation of the state-space relation
(RTR indices)
0,012
0,01
0,008
0,006
0,004
0,002
0
01-1990
01-1991
01-1992
01-1993
01-1994
01-1995
01-1996
01-1997
01-1998
01-1999
01-2000
01-2001
01-2002
01-2003
01-2004
01-2005
01-2006
01-2007
01-2008
01-2009
01-2010
01-2011
01-2012
01-2013
01-2014
01-2015
01-2016
01-2017
01-2018
01-2019
01-2020
01-2021
UK
France
Germany
Italy

Figure 4