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Hedging political risk in international portfolios

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Abstract

We show that internationally diversified portfolios carry sizeable political risk premia and expose investors to tail risk. We use a portfolio selection model for skewed distributions to obtain political efficient frontiers and develop a new asymptotic inference test to compare portfolio performance with and without hedging political risk. We find that politically hedged portfolios outperform a broad market index and the equally weighted portfolio for US, Eurozone, and Japanese investors and that political risk hedging is not subsumed by currency hedging. The diversification gains of politically hedged portfolios persist under currency hedging and transaction cost frictions. Hedging political risk induces equity home bias but does not fully explain the puzzle.

JEL Classification: C12, C13, C54, C61, F21, F30, G11.

Keywords: Block bootstrapping; conditional Value-at-Risk; equity home bias puzzle; international diversification; portfolio selection; higher-order moments.

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1 Introduction

We show that political risk is a significant determinant of risk and return in internationally diversified portfolios and ask what happens when we hedge it. Political risk is a well-documented determinant of financial market returns,[1] but international diversification studies focus on currency risk.[2] Political risk has been considered country-specific (Erb et al., 1996) and a determinant of market segmentation and, hence, diversifiable (Bekaert et al., 2016). However, recently, Kelly, Pástor, and Veronesi (2016) introduced political risk spillovers in asset pricing models, and Liu and Shaliastovich (2022) documented such spillovers empirically. Gala, Pagliardi, and Zenios (2023) showed that political risk is characterized by a global component, alongside country-specific shocks, priced in international equity returns; they support an APT interpretation (Ross, 1976) that political risk is a distinct factor (P-factor) from market and currency risk. These works imply that political risk may not be diversifiable, raising the question: Do internationally diversified equity portfolios carry a political risk premium? If yes, how do we manage portfolio political risk? And, importantly, do diversification benefits persist when political risk is hedged? Our paper provides answers.

Performance gains from international equities diversification over the home index or an equally weighted portfolio are well documented and persist under currency risk hedging (footnote 2). We show that such gains come with increasing exposure to political risk but hedging political risk does not eliminate the portfolio gains. This is a new result in the literature that holds both in and out of sample, robust to currency risk hedging, for US, Eurozone, and Japanese invstors. The hedged portfolios attain about 2.4% annualized higher excess returns in sample over the equally weighted portfolios, with lower tail risk, corresponding to an annualized Sharpe ratio increase of about 0.17.

To establish our findings, we develop a portfolio selection model with a political hedging constraint. We use a higher-moments performance measure to account for the skewed returns of the international markets (Ghysels, Plazzi, and Valkanov, 2016) and the fat tails of political events (Bremmer and Keat, 2010) that manifest themselves in the P-factor (Gala et al., 2023). Specifically, we obtain portfolios with varying, or zero, political risk exposures that maximize the *mean-to-CVaR* (MtC) performance measure of Martin, Rachev, and Siboulet (2003).[3]

MtC is second-order stochastic dominance (SSD) consistent (Ogryczak and Ruszczyński, 2002) so that the model portfolios are optimal for a broad class of investors with concave and non-decreasing utility functions. The model can be cast as a linear program for the case of finite and discrete distributions of returns, following Sahamkhadam et al. (2022); Stoyanov et al. (2007). Thus, we avoid making assumptions about an underlying distribution and obtain a

[1]See Bittlingmayer (1998) for evidence from 1880, and the recent works by Brogaard et al. (2020); Pástor and Veronesi (2012, 2013), among others.

[2]See the early works (Grubel, 1968; Levy and Sarnat, 1970; Solnik, 1974), recent advances (Black, 1989; Cambell et al., 2010; Driessen and Laeven, 2007; Glen and Jorion, 1993; Guidolin and Timmermann, 2008; You and Daigler, 2010), and the latest (Barroso et al., 2022; Topaloglou et al., 2020).

[3]This measure is called Stable Tail Adjusted Return Ratio (STARR) by Martin, Rachev, and Siboulet (2003) and mean/CVaR by Sahamkhadam, Stephan, and Östermark (2022). It is a coherent special case of the (α, β) ratio of Rachev, Stoyanov, and Fabozzi (2008). We use "mean-to-CVaR," avoiding references to stable distributions, which are not used in our work.

non-parametric model for international diversification abstracting from investor risk preferences without the need to specify a utility function (Guidolin and Timmermann, 2008).

To test our results' statistical significance, we need an asymptotically valid inference test to compare portfolio performance. However, such an inference test is missing for MtC. This gap in the literature limits the potential use of MtC models and casts a shadow on empirical studies using it. We develop an inference test on the equality of two MtC ratios to compare portfolio performance and derive the asymptotic null distribution of the test.

Equipped with a portfolio selection model for skewed returns and the inference test, we consider the tradeoff between political risk and portfolio performance. We estimate *country political betas* (β_P) through time-series regressions of country equity markets excess returns on the P-factor controlling for market risk; these are the country loadings on the P-factor. The betas align well with average excess returns (Appendix Figure 1), and we use them to impose a political risk constraint and trace the efficient frontier of portfolio political risk vs performance. This is the set of portfolios that achieve the maximum performance for a given level of political risk. We obtain the politically hedged portfolio by setting a zero beta political risk constraint.

It is possible to manage the political risk without portfolio optimization by *screening* ex-ante politically risky markets (Pedersen et al., 2021). However, high political risk is compensated with high expected returns compared to the sample mean (Erb et al., 1996), so screening out high political beta assets may lead to inefficient portfolios. We show that it does. We could also construct portfolios with a target political rating, as in Smimou (2014), without estimating betas on the political risk factor. We show that this does not ensure politically neutral portfolios. Higher ratings imply lower political risk, but the portfolio can be over-hedged with worsened performance. Incorporating the recent advances in pricing political risk within a portfolio selection model, we overcome these limitations and align portfolio selection with the APT.

We take the model to the data on a sample of 22 developed and 20 emerging markets spanning 1999–2019 to build internationally diversified portfolios and answer the research questions. Departing from existing literature that considers international diversification mainly from the perspective of US investors, we consider Eurozone and Japanese investors as well.[4] We first show that political risk is not diversified away, even with higher-order moments in portfolio selection, and that screening ex-ante politically risky markets is inefficient, so we need a political hedging model. We then proceed to our main empirical investigation using the inference test to compare politically hedged optimal MtC portfolios with the market index and the equally weighted portfolio (DeMiguel et al., 2009). Our main finding is that performance gains from international diversification survive political risk hedging. We also find that currency hedging does not eliminate political risk, and performance gains persist when hedging both currency and political risk. These results hold in and out of sample. A test during the period including the COVID-19 pandemic and the war in Ukraine, shows that the model effectively hedges the

[4]For the perspective of US investors see Barroso et al. (2022); Cosset and Suret (1995); De Roon et al. (2001); Glen and Jorion (1993); Grubel (1968); Guidolin and Timmermann (2008); Levy and Sarnat (1970); Smimou (2014); Topaloglou et al. (2020). Driessen and Laeven (2007); You and Daigler (2010) are exceptions that go beyond US investors.

unexpected political shocks from these events.[5] Finally, we use the model to show that restricting portfolio political exposure induces equity home bias (French and Poterba, 1991) but does not explain the puzzle.

Our findings survive several robustness tests: (i) The model is robust to tail risk estimation and data perturbations; (ii) the out-of-sample results survive a randomized test; (iii) transaction costs do not significantly alter the political risk of international portfolios or the efficacy of political hedging; (iv) short positions in developed markets lower political risk and may even achieve politically neutral portfolios when using the MtC model but not with Sharpe maximization.

The rest of the paper is organized as follows. Section 2 provides the motivation and discusses earlier literature. Section 3 gives the portfolio selection model and develops the novel inference test. Section 4 applies the model and the inference test to the problem of managing political risk in international portfolios and establishes the main empirical results and that political hedging induces equity home bias. Section 5 reports a battery of robustness tests. Section 6 concludes.

2 Motivation

We motivate the need to manage political risk by providing evidence that the benefits from international diversification come with increasing political risk. We use mean-variance (MV) optimization (Markowitz, 1952) to build international portfolios from a broad sample of 22 developed and 20 emerging markets for US, Eurozone, and Japanese investors. We use monthly returns to estimate moments from historical data spanning 1999-2019, solve for the maximum Sharpe ratio (SR) portfolio without short sales, and compare with the market index (I) and the equally weighted portfolio (EW).

We report the results in Table 1. In addition to the Sharpe ratio, we give the portfolio political risk as the weighted average of the portfolio countries' political risk. As a proxy for country political risk, we use the International Country Risk Guide-ICRG composite political ratings (PRS, 2005). This is a forward-looking rating, designed to only reflect political risk, as opposed to country risk that also embeds macroeconomic factors. It was shown to predict political risk realizations by Bekaert et al. (2014). We also report the portfolio political beta estimated as the weighted average of the portfolio countries' political betas; this is the portfolio *P-factor loading*. We finally report the political premium estimated as the product of the P-factor loading with the factor mean of 7.93%.[6]

We observe that the ICRG ratings, political betas, and political premia of the internationally diversified equally weighted and maximum Sharpe ratio portfolios are significantly larger than those of the home index. The SR portfolio of US investors has an economically and statistically significant political risk premium of 1.84% per annum (p.a.) compared to a market premium of 4.96% for the home index. The political premium for Eurozone investors is 1.34% compared to 4.99% for the market, and for Japanese, it is 2.05% compared to 3.93%. Political exposure

[5] We thank a referee for suggesting this test.

[6] With the betas and P-factor mean significant with p-values 0.00 (Table 1) and 0.01 (Gala et al., 2023, Table 1), respectively, the premia are, by the Bonferroni correction, significant at least at the 0.02 level. Henceforth, we report only the political betas.

Table 1 – Performance and political risk of international portfolios

This table reports annualized performance statistics of the MSCI home market index (I) and international portfolios in the home currency using equally weighted portfolios (EW) and the maximum Sharpe portfolio (SR). We also report the exposure of each portfolio to a global political risk factor. ICRG is the portfolio political risk, proxied as the weighted average of the countries' International Country Risk Guide composite political ratings (PRS, 2005), weighted by the portfolio proportionate holdings in each country. The portfolio political beta is the portfolio countries' political betas weighted average. Political premium is the product of the portfolio political beta with the P-factor mean of 7.93%, with a statistical significance upper bound estimated by the Bonferroni correction from the p-values of the betas and the P-factor. Also reported is the market premium, for comparison, and the differences in Sharpe ratio and political beta of I and the EW portfolios from the SR portfolios. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019. We use * to denote rejection of the null at the 10% level or less, with the p-values in parentheses.

	(a) US			(b) Eurozone			(c) Japan		
	I	EW	SR	I	EW	SR	I	EW	SR
Sharpe	0.42	0.46	0.74	0.25	0.52	0.86	0.32	0.51	0.77
Skewness	-0.64	-0.65	-1.70	-0.48	-0.53	-2.21	-0.34	-0.87	-2.61
ICRG rating	82.81	75.68	71.48	81.60	75.68	74.98	81.94	75.68	70.27
Political beta	-0.01	0.12*	0.23*	0.04	0.16*	0.17*	0.04	0.14*	0.26*
	(0.57)	(0.00)	(0.00)	(0.22)	(0.00)	(0.00)	(0.41)	(0.00)	(0.00)
Political premium	-0.08	0.93*	1.84*	0.31	1.30*	1.34*	0.32	1.13*	2.05*
	(1.00)	(0.02)	(0.02)	(0.44)	(0.02)	0.02	(0.82)	(0.02)	(0.02)
Market premium	4.96	5.91	5.53	4.99	4.35	3.77	3.93	5.97	5.65
Diff. to SR Sharpe	0.32*	0.28*	–	0.60*	0.34*	–	0.45*	0.26*	–
	(0.06)	(0.00)	–	(0.00)	(0.00)	–	(0.02)	(0.01)	–
Diff. to SR political beta	0.24*	0.11*	–	0.13*	0.00	–	0.22*	0.12*	–
	(0.00)	(0.01)	–	(0.01)	(0.92)	–	(0.00)	(0.07)	–

doubles when diversifying only in emerging markets and persists when diversifying only in developed.[7] The Sharpe ratio of international portfolios (0.74–0.86 p.a.) is double that of the home index (0.25–0.42), albeit it comes with sizeable political risk.

Maximum expected return portfolios were anticipated to have a higher political risk by Erb et al. (1996). However, the diversified maximum Sharpe portfolios are also politically exposed, showing that political risk has a non-diversifiable systematic component.

Although investors may be compensated to assume political risk, Giambona et al. (2017) use a psychometric test of managers' political risk perception to provide evidence that political risk drives investment decisions. Investors are willing to pay a premium to insure some of the political risk (Jensen, 2008), avoid investing before major political events such as elections (Jens, 2017), lower foreign direct investments (Busse and Hefeker, 2007), and hedge their political risk exposure to avoid crashes (Baur and Smales, 2020; Fisman et al., 2022). Political risk in portfolio analysis is also recognized as significant by rating agencies, the "big four" audit or

[7] The political premia when diversifying in emerging markets are 3.57% for the US and Japan and 3.89% for the Eurozone. For political risk as a determinant of developed markets returns, see Diamonte et al. (1996); Kelly et al. (2016), with lower average political betas than emerging by 0.33, and average ICRG ratings of 83 vs 67.

international management firms.[8] Our paper addresses the problem of investors who trade off portfolio performance to limit or hedge political risk exposure.

To illustrate this tradeoff, we select maximum Sharpe ratio portfolios with a target political rating, computed as the weighted average of the countries' ICRG ratings. That is, we revisit the model of Table 1, but instead of obtaining the politically unconstrained maximum Sharpe ratio portfolio, we vary the target rating to obtain the tradeoff curve illustrated by the red circles in Figure 1. We observed a monotonic relation of the maximum Sharpe ratios (left y-axis) with increasing portfolio ICRG ratings (right y-axis). However, when plotting Sharpe against the portfolio ex-post P-factor loadings (x-axis, blue curve), we note a hump-shaped relation that peaks at the politically unconstrained SR portfolio of Table 1. To the right, we have lower ICRG ratings with higher political risk, but performance worsens with returns not compensating for the additional political risk. These portfolios are inefficient. To the left, we increase ICRG ratings with reduced portfolio political betas and Sharpe ratios. These portfolios trade the ex-post portfolio political beta with the Sharpe ratio; only at the peak is the portfolio beta optimal. To obtain the *SR-political beta efficient frontier* to the left of the peak, we need to incorporate the estimated political betas in portfolio selection and maximize Sharpe subject to a target portfolio beta. Thus, we can delineate an efficient frontier similar to the blue line, which can not be identified using only the ratings (red circles). Importantly, we can identify the politically hedged (zero beta) portfolio.

However, we face a problem when using Sharpe as our performance ratio since it assumes normality, whereas higher-order moments are material for international diversification (Christoffersen et al., 2012; You and Daigler, 2010). The limitations of the normality assumption are exacerbated by the fat tails of political risk. This is demonstrated by the negative skewness of the SR portfolio in Table 1, which is three to seven times that of the home index. It motivates using a higher-moment performance measure for portfolio selection.

2.1 Related literature and contributions

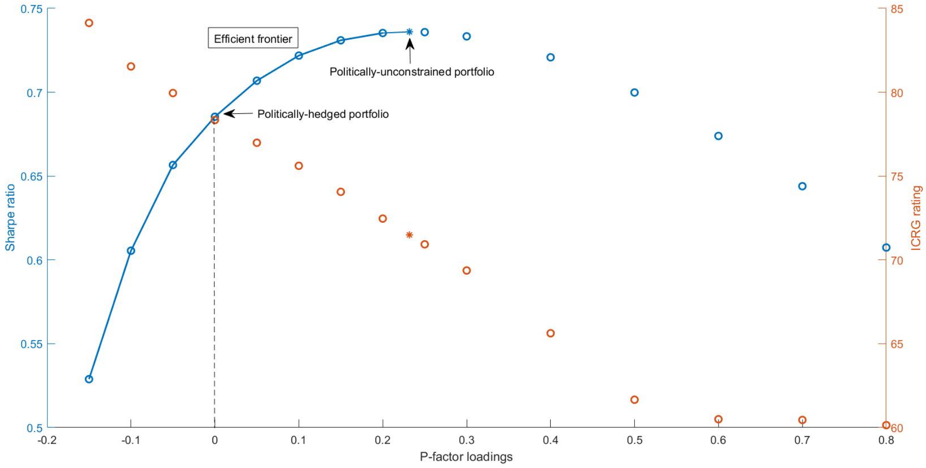
We bridge a gap between the large literature on international diversification and recent advances on political risk as a determinant of asset returns by adding political risk hedging to currency hedging. Political risk has been considered country-specific (Erb et al., 1996) and a main determinant of market segmentation, with (Bekaert et al., 2016, sec. 4.2) stating that political risk "is mainly a diversifiable risk for global investors."

Cosset and Suret (1995) were the first to study the benefits of diversifying into politically risky countries. We take the reverse and complementary vantage point to ask whether international diversification implies exposure to political risk. This question was raised by Rajan and Friedman (1997), who estimated a very wide range of political premia (9.4%–26.8%) for internationally diversified portfolios. Smimou (2014) developed a mean-variance multi-criteria

[8]See, e.g., Standard and Poors `https://www.spglobal.com/marketintelligence/en/documents/country-risk-and-sovereign-risk-1-.pdf`, Fitch `https://www.fitchratings.com/research/sovereigns/political-risk-key-driver-of-sovereign-ratings-20-03-2018`, Ernst and Young `https://esg.wharton.upenn.edu/wp-content/uploads/2022/06/ey-political-risk-and-corporate-performance-mapping-impact-final.pdf`.

Figure 1 – Political frontier of international portfolios

This figure illustrates the tradeoff between the annualized Sharpe ratio and political risk in international portfolios. Political risk is proxied by the portfolio ICRG ratings (red) and is measured by the portfolio P-factor loading (blue). The zero political beta portfolio is hedged from political risk. The portfolio political rating is the weighted average of the countries' ICRG ratings, weighted by the portfolio holdings in each country. The portfolio P-factor loading is the weighted average of the portfolio countries' political betas. The sample includes 22 developed and 20 emerging markets, spanning 1999 to 2019.



optimization model to restrict the exposure of portfolios to country political ratings and trade-off expected return, risk, and political ratings.

We advance these works in three directions. First, we estimate political betas on the recently uncovered global political risk P-factor to incorporate the market price of political risk in portfolio selection. Thus, our portfolio model is grounded on APT. We obtain a political premium of 1.84% for US investors with a tight confidence interval (p-value 0.02, Table 1) compared to the wide-ranging previous estimates of 9.4%-26.8% that far exceed the market premium of 5.53%. Note that the risk premium on the P-factor of 7.93% p.a. is higher but comparable to the market premium. Importantly, we obtain politically neutral portfolios, i.e., with zero betas (Cochrane, 2005), as opposed to arbitrarily limiting the portfolio political ratings, which may be inefficient, as demonstrated in Figure 1. Second, methodologically, we employ a portfolio selection model with higher-order moments and SSD consistency so that our results are not based on a normality assumption and are independent of a choice of a utility function. Third, we consider both currency and political risk, departing from the predominant literature that considers only currency risk (footnote 2) or Rajan-Friedman and Smimou that consider only political risk. We show that international diversification benefits persist when hedging both. We also differ from these authors in that we estimate the risk premium on global political risk using a single tradable global factor-mimicking portfolio instead of using the political rating proxies to compute the risk-return trade-off of internationally diversified portfolios.

Performance gains from optimally diversifying international portfolios have been documented by Grubel (1968); Levy and Sarnat (1970); Solnik (1974), and more recently by Barroso et al.

(2022); Topaloglou et al. (2020) with dynamic currency hedging. Despite the significance of higher-order moments for international diversification, mean-variance remains the workhorse of choice; see, among others, Cosset and Suret (1995); De Roon et al. (2001); Driessen and Laeven (2007); Glen and Jorion (1993); Grubel (1968); Levy and Sarnat (1970); Smimou (2014). Using an SSD portfolio selection model, we advance this literature. We show that diversification gains achieved without the normality assumption come with a significant increase in political risk. Still, the gains persist when hedging political and currency risks.

Some recent works incorporate political risk in financial decision-making in different contexts, using precious metals as a hedge (Baur and Smales, 2020), estimating a firm's political risk exposure through trade (Fisman et al., 2022), and establishing a relation of volatility innovations to geopolitical risk (Engle and Campos-Martins, 2020). Instead, we look at portfolios of financial assets, develop a hedging model with an associated inference test, and construct politically neutral portfolios by estimating P-factor betas.

Related to our work is the literature on inference tests for the Sharpe ratio (Lo, 2002; Schmid and Schmidt, 2010; Wright, Yam, and Yung, 2014). These tests are essential when comparing alternative portfolios to draw statistically significant conclusions. We contribute an asymptotically valid inference test for a non-parametric SSD-consistent portfolio optimization model with higher-order moments. This test is useful for other applications where deviations from normality can be a significant concern, including ESG investing (Pedersen et al., 2021), firm announcements or crashes (Chen et al., 2001), actuarial risks (Adcock et al., 2015), regime-switching (Francois et al., 2014), or political risk in currency markets (Gala et al., 2024).

Our empirical contribution is to document a statistically and economically significant political risk premium in international portfolios and show that hedging political risk does not eliminate the performance gains from international diversification. This finding holds when also hedging currency risk, in and out of sample. Importantly, we will see that we could effectively hedge the unexpected political shocks from the COVID-19 pandemic and the war in Ukraine.

We also contribute to the literature seeking to explain the equity home bias puzzle (French and Poterba, 1991). Several explanations have been advanced, but none are completely satisfactory (Ardalan, 2019; Gaar et al., 2022). Coeurdacier and Rey (2013) discuss explanations based on (i) hedging real exchange rate and non-tradable income risk, (ii) differences in transaction costs or tax treatments, and (iii) informational frictions and behavioral biases. Dahlquist et al. (2003) showed that country characteristics related to political risk could tilt portfolios towards the home market, but Guidolin and Timmermann (2008) pointed out that this explanation applies more to emerging markets and not to the limited diversification among developed economies. Smimou (2014) demonstrated that political rating constraints result in home-biased portfolios. However, this is a mechanical relation if the home market (US) has a higher political rating than the foreign. We show that even accounting for the rewards from political risk, it is optimal for investors to under-invest in emerging markets but the portfolios are tilted towards developed than the home, ruling out political risk as the sole determinant of the puzzle.

3 Portfolio selection and inference test

We define MtC efficient portfolios, add the political risk constraint, formulate the model as a linear program, and empirically compare its performance to mean-variance. We then develop an asymptotically valid test to compare MtC ratios and give an algorithm to implement it.

3.1 Mean-to-CVaR portfolios with a political constraint

Assume the random vector of n asset returns $\tilde{r} \in \mathbb{R}^n$ takes values in the probability space $(\mathbb{R}^n, \mathcal{B}(\mathbb{R}^n), \mathcal{P})$ where $\mathcal{B}(\mathbb{R}^n)$ denotes the Borel σ-field generated by the open sets of $\mathbb{R}^n$ and $\mathcal{P}$ a probability measure on $\mathcal{B}(\mathbb{R}^n)$, with expected value $\bar{r}$. The portfolio return $\tilde{r}_p = \tilde{r}^\top x$ is a function of $\tilde{r}$ and the vector of portfolio weights $x \in \mathbb{X}$, with

$$\mathbb{X} = \{x \in \mathbb{R}^n \mid x \geq 0, \sum_{i=1}^{n} x_i = 1\}. \tag{1}$$

$\mathbb{X}$ is the set of feasible portfolios with no short sales (NSS). To consider portfolios with covered short positions (long-short portfolios, LS), we set $x = x_+ - x_-$ as the difference between its long (x_+) and short (x_-) positions, with $x \in \mathbb{X}_S$, where

$$\mathbb{X}_S = \{x \in \mathbb{R}^n \mid x = x_+ - x_-,\ x_+ \geq 0,\ x_- \geq 0,\ \sum_{i=1}^{n}(x_{+i} - x_{-i}) = 1,\ \sum_{i=1}^{n} x_{-i} \leq 1\}. \tag{2}$$

Assuming a risk-free asset with return r_f, with $y \geq 0$ the proportion in the risky portfolio, the portfolio return is

$$\tilde{r}_c = y\tilde{r}_p + (1 - y)r_f. \tag{3}$$

We use CVaR (Artzner et al. 1999, see online Appendix Definition A.1) as the risk criterion in portfolio selection. The CVaR of a portfolio is a function of x, and minimizing CVaR for a varying target μ of portfolio expected returns $\bar{r}^\top x \geq \mu$, we obtain efficient frontiers in mean-CVaR space. We look at the portfolio with the maximum expected excess return to risk ratio.

From positive homogeneity and translation invariance, we have

$$\text{CVaR}_\alpha(\tilde{r}_c) = y\text{CVaR}_\alpha(\tilde{r}_p) - (1 - y)r_f, \tag{4}$$

where α is the confidence level. Solving (4) for y and substituting in (3) we get

$$\tilde{r}_c = \left(1 + \frac{\tilde{r}_p - r_f}{\text{CVaR}_\alpha(\tilde{r}_p) + r_f}\right) r_f + \frac{\tilde{r}_p - r_f}{\text{CVaR}_\alpha(\tilde{r}_p) + r_f}\text{CVaR}_\alpha(\tilde{r}_c). \tag{5}$$

Taking expectations of both sides and writing in terms of excess returns $\tilde{r}_e = \tilde{r}_p - r_f$,

$$\mathbb{E}(\tilde{r}_c) = \left(1 + \frac{\mathbb{E}(\tilde{r}_e)}{\text{CVaR}_\alpha(\tilde{r}_e)}\right) r_f + \frac{\mathbb{E}(\tilde{r}_e)}{\text{CVaR}_\alpha(\tilde{r}_e)}\text{CVaR}_\alpha(\tilde{r}_c). \tag{6}$$

The coefficient of the risk term is the *mean-to-CVaR ratio* (Martin et al., 2003)

$$\mathrm{MtC}_\alpha = \frac{\mathbb{E}(\tilde{r}_e)}{\mathrm{CVaR}_\alpha(\tilde{r}_e)}. \tag{7}$$

This performance ratio measures the expected excess return per unit of risk (Farinelli et al., 2008), akin to the Sharpe ratio of the MV efficient frontiers. We maximize this performance ratio and, dropping the subscript α, we solve

$$\mathrm{MtC}^* \doteq \max_x \frac{\mathbb{E}(\tilde{r}_e)}{\mathrm{CVaR}(\tilde{r}_e)}, \tag{8}$$

with $x \in \mathbb{X}$ or $x \in \mathbb{X}_S$ for no-short sales and short-sales portfolios, respectively.

MtC portfolios satisfy second-order stochastic dominance consistency. This follows from the SSD consistency of CVaR (Ogryczak and Ruszczyński, 2002); see online Appendix A.1. Hence, a broad class of investors with concave and non-decreasing utility functions prefer the optimal portfolios from this model.

We now add the portfolio political beta constraint. Let $\beta_P = (\beta_{P,i})_{i=1}^n \in \mathbb{R}^n$ be the vector of country political betas estimated through time-series regressions of country excess returns on the P-factor controlling for the market risk factor (MKT),

$$r_{i,t}^e = \alpha_{i,t} + \beta_{P,i}\,\text{P-factor}_t + \beta_{M,i}\,\mathrm{MKT}_t + \epsilon_{i,t}. \tag{9}$$

The portfolio beta, or P-factor loading, is given by $\beta_P^\top x$, and to limit the political risk exposure of a portfolio by a maximum value $\bar{\beta}$ we add to (8) the constraint

$$\beta_P^\top x \leq \bar{\beta}. \tag{10}$$

Solving (8) without this constraint, we obtain a portfolio unconstrained by political risk x_u^* with political beta given by $\beta_P^\top x_u^*$. Setting $\bar{\beta}$ to values lower than the unconstrained political beta reduces political risk. Solving (8) subject to (10) for different $\bar{\beta}$, we trace the *MtC-political beta efficient frontier* in mean-CVaR space akin to the frontier of Sharpe ratios in mean-variance space of Figure 1. For $\beta_P^\top x = 0$, the portfolio is politically hedged.

Our political constraint is a significant departure from Smimou (2014) that postulates a political ratings constraint in MV portfolios, e.g., $\mathrm{ICRG}^\top x \leq \overline{\mathrm{ICRG}}$ where ICRG denotes the vector of country political ratings and $\overline{\mathrm{ICRG}}$ is the target rating. A model maximizing the Sharpe ratio with this constraint generates the red circles of Figure 1. Instead, our model generates the MtC-political beta efficient frontier akin to the blue curve of Figure 1, further using a higher-moments performance ratio.

To solve the model, we cast (8) as a linear program for random returns taking discrete values from a finite set of equiprobable scenarios in $\mathbb{R}^n$ of cardinality S, using the fundamental minimization formula of Rockafellar and Uryasev (2002). The following result gives the LS model; the NSS case reduces to the linear program of Stoyanov et al. (2007) and is similar to Sahamkhadam et al. (2022).[9]

Theorem 3.1 (MtC optimization with short sales)**.** *Assuming that the* CVaR *on excess returns of every portfolio in the feasible set (2) is positive, then MtC portfolio optimization with covered short position is expressed using the transformed variables* x'_+, x'_- *as*

$$\begin{aligned} \max_{x'_+, x'_- \in \mathbb{R}^n, u' \in \mathbb{R}^S, \gamma' \in \mathbb{R}} \quad & \bar{r}^\top (x'_+ - x'_-) - r_f e^\top (x'_+ - x'_-) \\ s.t. \quad & \gamma' + \frac{1}{S(1-\alpha)} e^\top u' = 1 \\ & -R_e x'_+ + R_e x'_- - u' - e\gamma' \leq 0 \\ & e^\top x'_+ - e^\top x'_- > 0 \\ & 2e^\top x'_- - e^\top x'_+ \leq 0 \\ & u', x'_+, x'_- \geq 0, \end{aligned} \tag{11}$$

where R_e *is the matrix of excess returns of dimensions* $S \times n$*. Given the optimal solutions* x'^*_+ *and* x'^*_- *of (11), the optimal solution of maximum MtC portfolio optimization is obtained as* $x^* = \frac{1}{e^\top x'^*}$ *where* $x'^* = x'^*_+ - x'^*_-$*.*

(For the proof, see online Appendix A.2.)

3.1.1 Empirics of MtC portfolios

The MtC model appears more robust than Sharpe maximization. We solve an identical problem using mean-CVaR and mean-variance to minimize their respective risk measure for different target expected returns and obtain the efficient frontiers.[10] For each point on the frontier, we compute the optimal risk measure with each model, CVaR* and Variance*, respectively, the corresponding non-optimized risk measures denoted by Variance(CVaR*) and CVaR(Variance*), and the standardized errors:

$$\frac{\text{Variance}(\text{CVaR}^*) - \text{Variance}^*}{\text{Variance}^*} \text{ and } \frac{\text{CVaR}(\text{Variance}^*) - \text{CVaR}^*}{\text{CVaR}^*}. \tag{12}$$

The errors would be zero for normally distributed returns; for our test problem, they are displayed in Figure 2 (Panel A). The dashed line shows the error of MtC (i.e., the first term in (12), and the solid line is the error of MV (i.e., the second term). The dots indicate the MtC and Sharpe maximization portfolios. The maximum expected return portfolios are identical, and the errors are zero, but as the target expected returns are reduced, the errors increase. The mean-CVaR portfolios have lower errors than the MV portfolios, with the difference increasing

[9]Sahamkhadam et al. (2022) use MtC of returns instead of excess returns as we do, and the formulation involves a parameter that needs to be calibrated ex-ante. We use excess returns and our formulation does not require a parameter calibration and has one less variable and $n + 1$ fewer constraints than these references.

[10]We solve the model of a US investor in currency-hedged returns using our full sample with no short sales or constraints on political risk. Consistent results are obtained in developed or emerging markets as well. For details on the implementation see section 4.

for expected returns below about 22% p.a., which is large. The error is substantial at the maximum MtC portfolio, with variance 17% larger than the optimum, and the maximum SR portfolio with CVaR 44% larger than the optimum.[11]

We also find that optimal MtC portfolios better satisfy investor preferences for positive skewness (Mitton and Vorkink, 2007) than Sharpe portfolios. We block bootstrap with replacement, the time series of returns to generate several samples (see Section 3.2.1 for details) and re-optimize MtC and SR to obtain a distribution of portfolio skewness. In Figure 2 (Panels B-C), we plot the skewness of the MtC and SR portfolios for developed and emerging markets. The skewness for developed markets is -0.57 with SR and -0.42 with MtC; for emerging markets, it increases from -0.26 to -0.05.

3.2 Statistical inference test

We consider two instances of (8), with solutions x_1^* and x_0^*, and optimal values MtC_1^* and MtC_0^*. We wish to test the null hypothesis H_0 against the alternative H_1,

$$\text{H}_0 : \text{MtC}_1^* - \text{MtC}_0^* = 0 \quad \text{vs} \quad \text{H}_1 : \text{MtC}_1^* - \text{MtC}_0^* > 0. \tag{13}$$

We propose the test statistic

$$T_S = \widehat{\text{MtC}}_1 - \widehat{\text{MtC}}_0, \tag{14}$$

where for $j \in \{0, 1\}$

$$\widehat{\text{MtC}}_j = \frac{\overline{r}_j}{\widehat{\text{CVaR}}_j},$$

$$\overline{r}_j = \frac{1}{S}\sum_{t=1}^{S} r_{j,t}, \quad \widehat{\text{CVaR}}_j = -\frac{1}{(1-\alpha)}\frac{1}{S}\sum_{t=1}^{S} r_{j,t} \cdot \mathbb{1}\big(r_{j,t} \leq \widehat{\zeta}_{j,1-\alpha}\big),$$

and $\widehat{\zeta}_{j,1-\alpha}$ denotes the empirical $(1-\alpha)$ quantile of the time series $r_{j,t}$, $t = 1, 2, \ldots, S$.

We next give a theorem for the limiting distribution of the test statistic T_S if the null hypothesis H_0 is true, under some general assumptions, where $\xrightarrow{d}$ denotes convergence in distribution:

(i) The sample $r_t = (r_{1,t}, r_{0,t})^\top$, $t = 1, 2, \ldots, S$ stems from a stationary process and satisfies the moment and weak dependence conditions required so that the central limit theorem for the means sequence $\{\overline{r}_S = (1/S)\sum_{t=1}^{S} r_t, S \in \mathbb{N}\}$ holds. That is, the sequence $\{\sqrt{S}(\overline{r}_S - E(r_1)), S \in \mathbb{N}\}$ satisfies $\sqrt{S}(\overline{r}_S - E(r_1)) \xrightarrow{d} \mathcal{N}(0, \sum_{h\in\mathbb{Z}} Cov(r_t, r_{t+h}))$, as $S \to \infty$, with $0 < \sum_{h\in\mathbb{Z}} Cov(r_{j,t}, r_{j,t+h}) < \infty$, for $j \in \{1, 2\}$.

(ii) The distribution function F_j of $r_{j,t}$, $j \in \{0, 1\}$, is continuous and differentiable at $\zeta_{j,\beta}$ for any $\beta \in (0, 1)$ with positive derivative $f(\zeta_{j,\beta}) > 0$.

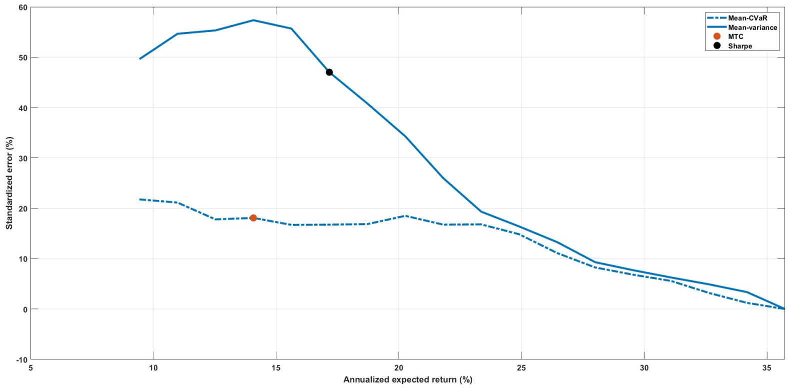
Assumption (i) is general and covers several interesting cases of processes, provided r_t has finite variance. For instance, this assumption is satisfied if the process $\{r_t = (r_{1,t}, r_{0,t})^\top, t \in \mathbb{Z}\}$

[11] The intuition behind this figure is the following: For $\alpha > 0.62$, CVaR provides an upper bound for the variance of normal distributions, and when the distribution deviates from normality but α is much larger (like 0.95 used in typical applications), CVaR again provides a bound, so, in general, a minimum CVaR portfolio has low variance. However, this is not a general property of MtC optimization.

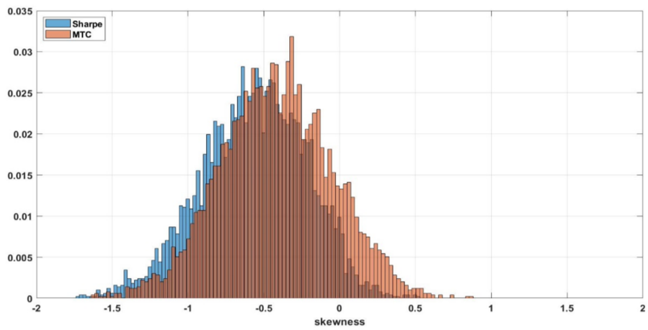
Figure 2 – Empirics of Mean-to-CVaR portfolios

This figure illustrates the relative benefits of mean-CVaR compared to mean-variance optimization. Panel A shows the standardized errors from (12) of the annualized portfolio risk measure not optimized by each model, respectively CVaR (solid line) and variance (dashed line), from its optimal value for an identical test problem over the sample of 22 developed economies and 20 emerging markets spanning 1999–2019. Panels B and C show the distribution of the annualized skewness of optimal mean-to-CVaR (MtC) and Sharpe ratio (SR) portfolios from 10000 block bootstrapped samples of asset returns for developed and emerging markets, respectively.

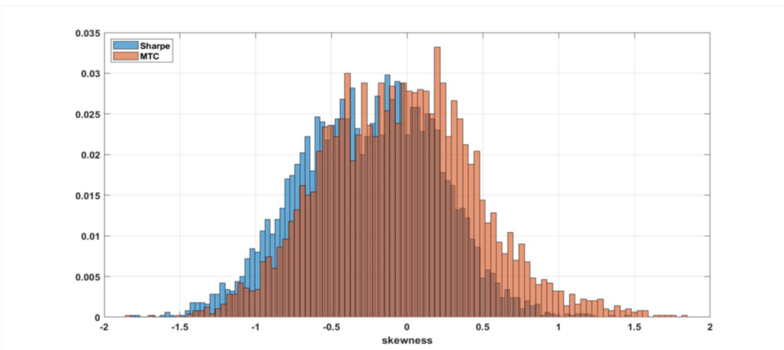
(a) Standardized errors of risk



(b) Portfolio skewness in developed markets



(c) Portfolio skewness in emerging markets



is a martingale sequence (Hall and Heyde, 1980). It is also true if the same process satisfies some mixing type conditions, like α-mixing (Ibragimov and Linnik, 1971) or other types of weak dependence conditions, including ARCH and GARCH processes. Assumption (ii) is a standard condition in CVaR analysis. Under the null, $\text{MtC}_1^* = \text{MtC}_0^*$, and we will write MtC^* for simplicity.

Theorem 3.2 (Limiting distribution of test statistic)**.** *Suppose that Assumption (i) is satisfied and that the null hypothesis in (13) is true. Then, as $S \to \infty$,*

$$\sqrt{S} \cdot T_S \xrightarrow{d} \mathcal{N}(0, \tau_0^2), \quad \tau_0^2 = \underline{c}^\top \Sigma_r \underline{c},$$

$$\textit{where } \underline{c} = \Big(\frac{1}{\text{CVaR}_1}, -\frac{\text{MtC}^*}{\text{CVaR}_1}, -\frac{1}{\text{CVaR}_0}, \frac{\text{MtC}^*}{\text{CVaR}_0}\Big)^\top$$

and $\Sigma_r = \sum_{h=-\infty}^{\infty} \text{Cov}(R_t, R_{t+h})$, and the vector R_t is defined as

$$R_t = \big(r_{1,t}, -\frac{1}{(1-\alpha)}\ r_{1,t} \cdot \mathbb{1}(r_{1,t} \leq \zeta_{1,1-\alpha}),\ r_{0,t}, -\frac{1}{(1-\alpha)}\ r_{0,t} \cdot \mathbb{1}(r_{0,t} \leq \zeta_{0,1-\alpha})\big)^\top.$$

Here, $\zeta_{j,1-\alpha}$, $j \in \{0,1\}$, denotes the $(1-\alpha)$ quantile of the distribution of $r_{j,t}$.

(For the proof, see online Appendix A.3.)

It holds that if H_0 is wrong, then $T_S \to c \neq 0$ in probability, implying that our test is consistent with power approaching unity as the sample size tends to infinity. Hence, when the null hypothesis H_0 is wrong, it is rejected at any level $\beta \in (0,1)$ with probability tending to one as $s \to \infty$, i.e., $P(\sqrt{S} \cdot T_S \geq z_{1-\beta}) \to 1$, where $z_{1-\beta}$ denotes the $(1-\beta)$ quantile of $\mathcal{N}(0, \tau_0^2)$. Implementing this test requires an estimation of τ_0^2, where the difficult part is estimating the covariance matrix Σ_r.

A corollary of Theorem 3.2 allows us to test the pair of hypotheses

$$\text{H}_0 : \text{CVaR}_1 = \text{CVaR}_0 \quad \text{vs} \quad \text{H}_1 : \text{CVaR}_1 > \text{CVaR}_0, \tag{15}$$

using the test statistic

$$C_S = \widehat{\text{CVaR}}_1 - \widehat{\text{CVaR}}_0. \tag{16}$$

Corollary 3.1. *Under the null hypothesis in (15), as $S \to \infty$ we have $\sqrt{S} \cdot C_S \xrightarrow{d} \mathcal{N}(0, v_0^2)$ with $v_0^2 = \underline{e}^\top \Sigma_r \underline{e}$, where $\underline{e} = \big(0,1,0,-1\big)^\top$ and Σ_r is given in Theorem 3.2.*

The corollary gives the limiting distribution of C_S under the null.

3.2.1 Inference test algorithm

We use the following block bootstrapping algorithm following Paparoditis and Politis (2003) to estimate the covariance matrix Σ_r of τ_0^2 and implement the MtC testing procedure of Theorem 3.2.

Step 1: Select a block size $b \in \mathbb{N}$, $b < S$ and let $k = \lceil S/b \rceil$. Assume for simplicity that S/b is an integer. Denote by $B_{t,b} = \{(r_{1,t+s-1}, r_{0,t+s-1}), s = 1, 2, \dots, b\}$ the block of b consecutive observations having starting point t, where $t \in \{1, 2, \dots, S-b+1\}$.

Step 2: Select randomly (i.e., with replacement) k such blocks $B_{t,b}$ from the set of all possible $S-b+1$ blocks and join them together in the order selected to form a bivariate set of pseudo observations denoted by $(r^*_{1,t}, r^*_{0,t})$, $t = 1, 2, \ldots, S$.

Step 3: Calculate $\overline{Y}^*_S = \frac{1}{S}\sum_{t=1}^{S} R^*_t$, where

$$R^*_t = \big(r^*_{1,t}, -\frac{1}{(1-\alpha)}\, r^*_{1,t} \cdot \mathbb{1}(r^*_{1,t} \le \widehat{\zeta}^*_{1,1-\alpha}),\ r^*_{0,t},\ -\frac{1}{(1-\alpha)} r^*_{0,t} \cdot \mathbb{1}(r^*_{0,t} \le \widehat{\zeta}^*_{0,1-\alpha})\big)^\top.$$

Step 4: Repeat Steps 2 and 3 a large number of times, say B times, and denote by $\overline{Y}^*_{S,i}$, $i = 1, 2, \ldots, B$, the replications of $\overline{Y}^*_S$ obtained by these repetitions. Calculate

$$\Sigma^*_r = \frac{1}{B}\sum_{i=1}^{B}\big(\overline{Y}^*_{S,i} - \overline{M}_B\big)\cdot\big(\overline{Y}^*_{S,i} - \overline{M}_B\big)^\top, \text{ where } \overline{M}_B = \frac{1}{B}\sum_{i+1}^{B}\overline{Y}^*_{S,i}.$$

Step 5: Let $\widehat{\tau}^2_0 = S \cdot \underline{\widehat{c}}^\top \Sigma^*_r\, \underline{\widehat{c}}$, where

$$\underline{\widehat{c}} = \Big(\frac{1}{\widehat{\mathrm{CVaR}}_1}, -\frac{\widehat{\mathrm{MtC}}}{\widehat{\mathrm{CVaR}}_1}, -\frac{1}{\widehat{\mathrm{CVaR}}_0}, \frac{\widehat{\mathrm{MtC}}}{\widehat{\mathrm{CVaR}}_0}\Big)^\top,$$

and $\widehat{\mathrm{MtC}} = (\widehat{\mathrm{MtC}}_1 + \widehat{\mathrm{MtC}}_0)/2$.

Step 6: Reject the null hypothesis H_0 if $\sqrt{S}\cdot T_S \ge z_{1-\beta}$, where $z_{1-\beta}$ denotes the $(1-\beta)$ quantile of the $\mathcal{N}(0, \widehat{\tau}^2_0)$ distribution.

We can modify the algorithm to test the CVaR null hypothesis in (15). From Corollary 3.1 it follows that the null hypothesis is rejected if $\sqrt{S}\cdot C_S \ge z_{1-\beta}$ with $z_{1-\beta}$ the upper $(1-\beta)$ percentage point of the $\mathcal{N}(0, \widehat{v}^2_0)$ distribution. Here $\widehat{v}^2_0 = S \cdot \underline{e}^\top \Sigma^*_r\, \underline{e}$, where Σ^*_r is the estimator obtained in Step 4.

3.2.2 Robustness of the inference test algorithm

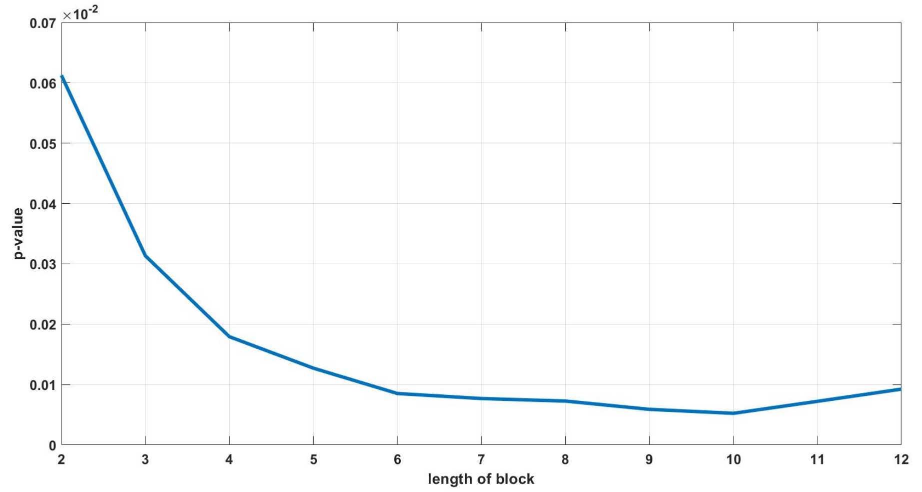
The algorithm is robust to the block size, and the test's p-values converge fast for a relatively large number of repetitions. We display in Figure 3 results using the algorithm to compare the peak MtC portfolio from Figure 1 with the US index for a sample size $S = 252$. Panel A shows the algorithm's sensitivity to the block size and Panel B the convergence of p-values.

Bootstrapping with overlapping blocks is quite efficient, and empirical evidence from the literature suggests that it works well for block sizes in the range $[S^{1/3}, S^{1/4}]$, i.e., 4 to 6 for our sample. As we observe, the null is rejected in favor of the alternative for all block sizes from 2 to 12. Panel B shows the behavior of the test concerning the number of bootstrap repetitions, and we observe that for a relatively large number of repetitions, the p-value estimates differ at the fourth decimal point. The block bootstrapped implementation is robust to the block size and for number of iterations exceeding 7000. We set $b = 6$ and $B = 10000$ in all empirical tests.

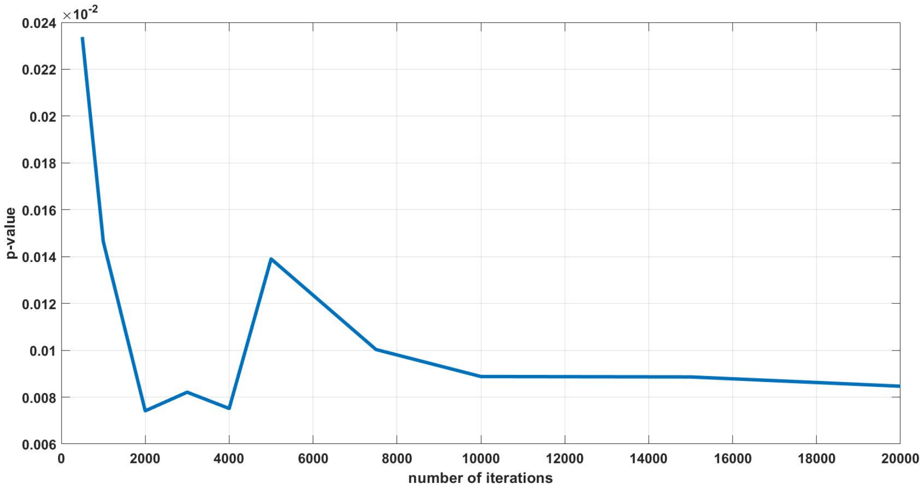
Figure 3 – Bootstrapped differences under the null hypothesis

This figure illustrates results with the inference test algorithm in comparing the MtC statistics of the maximum MtC portfolio from Figure 2 with the local (US) index benchmark, with a sample size $S = 252$. Panel A displays the sensitivity of the algorithm to varying overlapping block sizes. Panel B shows the convergence of the p-value estimates for different repetitions.

(a) Sensitivity to block size



(b) Convergence of p-value estimates



4 Global political risk management

We take the MtC model to the data. We construct international portfolios for US, Eurozone, and Japanese investors with political risk constraints, and use the inference test to compare the politically hedged portfolios with unrestricted portfolios and the benchmarks. Our tests proceed in four steps. First, we document a political risk exposure of well-diversified international portfolios even when accounting for higher order moments; this shows that the motivating results of Table 1 are not an artifact of the normality assumption of the Sharpe ratio. Second, we conduct our main tests in sample; these show that ex-ante screening politically risky countries is inefficient and that international diversification gains persist when hedging political and currency risk. Third, we test out of the sample; this shows that our results are robust and that political risk hedging can be effective during unexpected shocks like the ones following the COVID-19 pandemic and the war in Ukraine. In a final step, we consider diversification only in emerging markets and document political risk as a determinant of equity home bias.

To implement the model, we set $\alpha = 0.95$ and successfully perform additional tests in Section 5 with different α to rule out that outliers drive our results. We take the S discrete scenarios as all historically observed returns of the n assets assumed equiprobable. All tests are performed using monthly data, and results are reported monthly since we can not extrapolate annualized CVaR.

4.1 Data

We describe our data relating to political risk and asset returns.

4.1.1 P-factor and beta estimation

The P-factor is from Gala et al. (2023). It captures the effects of global political risk on asset prices by constructing a factor mimicking portfolio (Cochrane, 2005). Specifically, countries are sorted using conditional double sorts on political stability and confidence in economic policy ratings from the Ifo World Economic Survey-WES (Becker and Wohlrabe, 2007), first on the less volatile political stability dimension to maximize the spread in political stability across portfolios, and second on the policy dimension. Both variables are sorted in terciles, and the P-factor is the return of an equally weighted zero-cost tradeable portfolio, going long on countries with low ratings and short on countries with high ratings. The P-factor tracks monthly returns with portfolios rebalanced upon the release of new political ratings.

The authors also consider alternative political ratings from the literature —aggregate ICRG political rating and the World Bank political stability indicator— in constructing the P-factor. They find robust results, with the double-sorted portfolios being the most informative, with an almost additive premium compared to the single-sorted portfolios. This analysis is important since political risk is a complex multi-dimensional concept (Sottilotta, 2016), and we use the P-factor that was shown to be robust.

Gala et al. (2023) ruled out reverse causality that the experts give political ratings that are contaminated by the macroeconomic or financial conditions of the country. The politics and policy ratings correlations with sixteen other macroeconomic, financial, and trade variables from WES are low and statistically insignificant in the cross-section and inter-temporally. Importantly, when the political ratings are orthogonalized on the sixteen variables (i.e., running linear regressions with political ratings as independent variables and using the residual term in their place), they still predict country stock market returns and the P-factor remains priced. The ratings are characterized by a strong factor structure, with the first principal component explaining a sizeable fraction of the total rating variance. Global or local factors of the benchmark international asset pricing models do not span the P-factor, and adding the P-factor to any of the benchmark models increases the cross-sectional R^2 by an order of magnitude.

Importantly, for the international diversification problem, market segmentation does not drive the P-factor. In a subsequent paper (Gala et al., 2024) —where the global political risk factor is documented to be "everywhere" across stocks, bonds, and currency assets— the authors show that when ratings are demeaned, and hence all countries have the same (zero) political rating, it is still possible to construct a factor with a significant risk premium with about equal exposures to emerging and developed markets.

The P-factor has an economically and statistically significant average return, equivalently political risk premium, of 7.93% p.a., with a Sharpe ratio of 0.45 p.a. and MtC 0.081.[12] Its skewness is 2.11 with an excess kurtosis of 13.80, compared to -0.66 and 4.69 for the global market portfolio. The minimum annualized return during our sample period is -11.14%, with 0.01 return quantile of -9.14%, 0.05 quantile of -6.06%, and 0.10 quantile of -4.78%. These statistics highlight the fat tails of political risk.

We estimate the loadings of all countries on the P-factor (political betas) using (9). The cross-country dispersion of political ratings is reflected in the dispersion of P-factor loadings and lines up well with average excess returns with an R^2 of 0.41 (Appendix Figure 1).

4.1.2 Asset returns

We diversify into 22 developed and 20 emerging stock market indices spanning twenty-one years from January 1, 1999, to December 31, 2019; this sample is quite comprehensive.[13] We use the MSCI investable indices with monthly USD returns, including dividends. Using investable indices, we avoid positive biases from frictions such as illiquidity, index replicability, and countries-periods when trading may not be feasible. Summary statistics are given in the Data Appendix. For US investors, the risk-free rate is the one-month US T-Bill rate,[14] for the Eurozone we compute excess EUR returns over the one-month Euribor from Refinitiv Eikon. For Japan, we convert local USD returns into JPY using contemporaneous spot rates and calculate excess returns over the 30-day deposit of domestic banks. Data are from Datastream.

[12] The factor mean over the period 1992–2016 reported in Gala et al. (2023) is 11.02%.

[13] From earlier literature Cosset and Suret (1995) use 36 countries over eight years (1982 to 1991, excluding the market crash year 1987); De Roon et al. (2001) use 17 emerging markets over 11 years (1985–1996); Driessen and Laeven (2007) use 52 countries over 17 years (1985–2002); Christoffersen et al. (2012) use 29 to 33 countries over 36 years (1973–2009); Smimou (2014) uses 23 countries over nine years (1997–2005).

[14] Obtained from Kenneth French's website, `http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html#Developed`

Table 2 – Political risk of maximum Sharpe and mean-to-CVaR portfolios

This table reports the exposure of international portfolios to a global political risk factor β_P, the political premium computed as the product of β_P and the expected return of the P-factor, and the moments of monthly portfolio returns. Portfolios are constructed by optimizing mean-to-CVaR (MtC) and Sharpe (SR). We consider no-short-sales (NSS) and long-short strategies (LS) in developed markets. We denote by $\mu_n \doteq \mathbb{E}\left[(r - \mathbb{E}\left[r\right])^n\right]$ the n^{th} central moment of the portfolio returns. The second (third and fourth) central moments have been rescaled by multiplying the original values by 10^3 (10^5). The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

(a) US

	NSS		LS	
	SR	MtC	SR	MtC
β_P	0.23*	0.34*	0.12*	0.09
	(0.00)	(0.00)	(0.04)	(0.13)
Pol. prem.	1.84	2.66	0.97	0.68
μ	14.80	16.30	19.77	15.40
μ_2	3.35	4.20	2.49	1.83
μ_3	-10.44	-8.27	-3.34	1.68
μ_4	6.63	9.09	2.27	1.26

(b) Eurozone

	NSS		LS	
	SR	MtC	SR	MtC
β_P	0.17*	0.31*	0.15*	0.15*
	(0.00)	(0.00)	(0.01)	(0.02)
Pol. prem.	1.34	2.50	1.18	1.18
μ	13.20	15.25	20.03	18.64
μ_2	1.97	2.88	1.93	2.06
μ_3	-6.10	-3.35	-1.12	4.39
μ_4	2.11	4.09	1.19	1.56

(c) Japan

	NSS		LS	
	SR	MtC	SR	MtC
β_P	0.26*	0.42*	0.15*	0.08
	(0.00)	(0.00)	(0.04)	(0.29)
Pol. prem.	2.05	3.36	1.16	0.66
μ	17.05	19.09	23.20	20.27
μ_2	4.05	5.38	3.22	3.17
μ_3	-21.29	-15.98	-9.94	7.85
μ_4	11.44	17.62	5.22	5.01

For currency-hedged returns, we use one-month forward exchange rates from Datastream and proxy currency risk hedging by multiplying end-of-month index prices by the corresponding forward exchange rate. The absence of triangular arbitrage derives the forward exchange rates to EUR and JPY from the spot and forward rates to the USD. We do not have a complete time series of forward exchange rates for Brazil, Chile, China, Colombia, Egypt, Israel, Peru, Poland, Russia, South Korea, and Turkey. We complete the time series of hedged returns using returns from futures contracts when available or estimate synthetic replications of futures returns as the difference between the local stock market returns and the risk-free rate (Asness et al., 2013).

The descriptive statistics of excess returns in the Data Appendix show considerable differences in the moments across countries, with most indices being negatively skewed with significant tail risk. Jarque-Bera tests reject normality at conventional levels for all countries except Colombia, Japan, and South Africa. Comparing the country index statistics to the EW portfolio suggests potential diversification benefits for all investors, with larger gains for the Eurozone and Japan.

4.2 Political risk in international portfolios

We first construct politically unrestricted optimal SR and MtC portfolios with NSS restrictions or covered LS positions in developed markets. In Table 2, we report the political beta, political risk premia, and return moments.

SR and MtC NSS portfolios have economically large and statistically significant political beta and risk premia. Neither the Sharpe ratio, with the assumption of normality, nor MtC optimization, accounting for higher-order moments, diversifies away the political risk. This result is statistically significant (p-values 0.00) for all three investors.

The LS portfolios attain lower political risk by taking long-short positions in markets that hedge the political risk for each other. The reduction is much more pronounced for the MtC portfolios, where the political beta is not statistically significant for the US and Japan but remains significant for the EU. This result is explained by the high cross-sectional correlations of the moments with political betas in the range of 0.78–0.82 for all three markets, so a tail risk model that creates positive skewness and lower kurtosis can reduce political risk. However, in the interesting NSS case, since no short sales are possible in emerging markets in general, international portfolios carry significant political premia.

The SR portfolio return distribution is negatively skewed and leptokurtic, unlike the MtC portfolios that better satisfy investor preference for positive skewness (Mitton and Vorkink, 2007).

This test affirmatively answers our first research question: internationally diversified portfolios carry significant political risk premia. We now turn to political risk management and test whether diversification benefits persist when political risk is hedged.

4.3 In sample tests

We perform in-sample tests, solving the NSS MtC model without and with the political constraint, and then repeat the test using currency-hedged returns. In Figure 4, we display the β_P-MtC political frontier obtained from model (8)-(10) for US investors and observe significant tradeoffs between performance and political risk exposure. The P-factor loadings β_P can be reduced from the unconstrained 0.34 (Table 2) with an MtC performance ratio of 0.106 to -0.15 with a lower performance ratio of 0.063. The motivating evidence in Figure 1 holds when we use a performance ratio with higher order moments instead of mean-variance analysis and optimizing over the β constraint set.

4.3.1 Screening politically risky countries

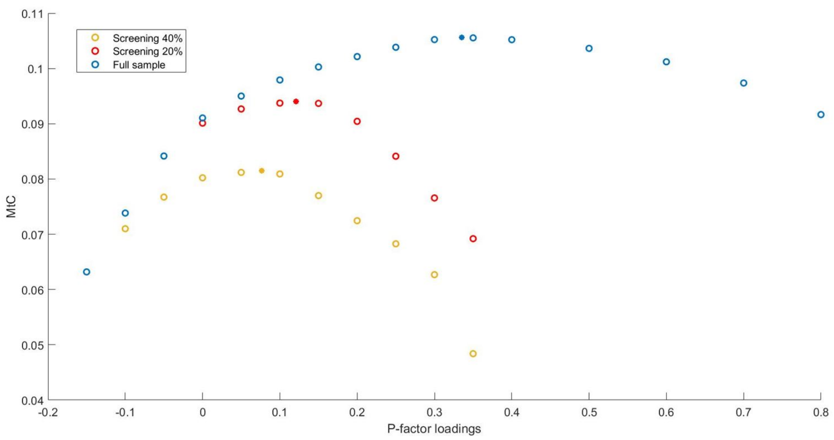
We also obtain frontiers after removing potential investments at the bottom 20% or 40% ICRG ratings, screening out the most politically risky countries. We display them with the frontier over the whole sample in Figure 4 and observe that screening leads to inefficient portfolios. Without screening, the unconstrained political beta portfolio is heavily exposed to Russia, which has the highest in-sample MtC ratio and very high political beta (0.77). Tightening the bound $\bar{\beta}$ reduces the allocation to Russia and increases the allocation to Denmark with negative beta and one of the highest MtC ratios among developed markets. The Russia-Denmark portfolio reduces political risk while preserving performance gains, whereas screening removes Russia.

4.3.2 Hedging political risk

We now turn to the central question of what happens to international portfolios with zero political betas. We start with a baseline in-sample test in the home currency without short sales and compare the unconstrained and the politically hedged portfolios with the benchmarks. In Table 3, we report the political beta, average excess return, CVaR, and MtC for each portfolio.

Figure 4 – The β_P-MtC political frontier of international portfolios

This figure illustrates the tradeoff between monthly MtC and political risk in internationally diversified portfolios, where the political beta measures portfolio political risk. It also shows the frontiers obtained after screening the set of assets to remove the worse-rated 20% (resp. 40%) by the ICRG ratings. The frontiers are obtained using the MtC model with NSS. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019.



We also report the Sharpe ratios of the optimal MtC portfolios to show that gains in the MtC ratio do not worsen Sharpe.[15] We observe the benefits of international diversification by comparing the unrestricted portfolio (U) performance ratios with benchmarks I and EW using our inference test.[16] Comparing H with the benchmarks, we observe that the benefits persist under political hedging. Comparing the CVaR of U and EW with H, we observe that crashes diminish with political hedging.

The US index's statistically insignificant political beta increases with EW diversification to a statistically significant 0.12 and 0.34 for the unconstrained MtC portfolio. Political risk premia are 0.95% for EW and 2.69% for MtC. The results with the hedged portfolios affirmatively answer our research question of whether international diversification persists when hedging political risk. We observe performance gains when diversifying internationally without any restrictions on political risk (column "U-I"). MtC doubles from 0.053, and Sharpe increases by 0.09 from 0.12, significant at conventional levels, with slightly smaller increases of 0.047 and 0.08, respectively, over EW ("U-EW"). When hedging political risk, the performance gains remain statistically significant and economically large ("H-I" and "H-EW"). MtC increases by 0.038 for I and 0.032 for EW, and Sharpe increases by 0.07 (I) and 0.06 (EW). We also

[15]The Sharpe ratios corroborate, in general, the MtC results, but they are suboptimal to be used for inferences. Optimizing Sharpe to draw inferences would ignore the higher-order moments.

[16]The Augmented Dickey-Fuller test showed no evidence of a unit root behavior of the time series of excess returns for the hedged and unhedged optimal portfolios considered in the inference test, as indicated by a statistically significant p-value (< 0.01). Simple plots also supported this fact. Additionally, analysis of the time series of returns reveals no discernible time-dependent trends. Furthermore, autocorrelation analysis of overlapped subsamples of the time series does not exhibit any significant serial correlation (p-value < 0.05). Therefore, we have no statistical evidence that our data do not satisfy the stationarity assumption, and our inference test applies.

Table 3 – Hedging political risk of international portfolios

This table reports performance statistics of the MSCI home market index I and international portfolios for US, Eurozone, and Japan investors, using equally weighted portfolios EW and mean-to-CVaR unrestricted optimal portfolios (U) and with political risk hedging (H) and no short sales. Returns are monthly in the home currency, and political risk is hedged with net zero exposure to the P-factor. Reported are the exposures to a global political risk factor β_P, and the monthly performance ratios mean-to-CVaR (MtC) and Sharpe ratio. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

	I	EW	U	H	U-I	H-I	U-EW	H-EW
				(a) US				
β_P	-0.01	0.12*	0.34*	0.00	0.35*	0.01	0.22*	-0.12*
	(0.57)	(0.00)	(0.00)	(1.00)	(0.00)	(0.82)	(0.00)	(0.01)
Av. excess return	0.52	0.71	1.36	1.02	0.83	0.50	0.64	0.31
CVaR	9.90	12.16	12.86	11.26	2.96	1.36	0.70	-0.90
MtC	0.053	0.059	0.106	0.091	0.053*	0.038*	0.047*	0.032*
					(0.06)	(0.10)	(0.01)	(0.02)
Sharpe	0.12	0.13	0.21	0.19	0.09*	0.07	0.08*	0.06*
					(0.10)	(0.15)	(0.03)	(0.04)
				(b) Eurozone				
β_P	0.04	0.16*	0.31*	0.00	0.27*	-0.04	0.15*	-0.16*
	(0.22)	(0.00)	(0.00)	(1.00)	(0.00)	(0.44)	(0.01)	(0.00)
Av. excess return	0.36	0.68	1.27	0.95	0.91	0.59	0.59	0.27
CVaR	12.16	10.80	10.69	10.64	-1.47	-1.52	-0.11	-0.16
MtC	0.030	0.063	0.119	0.090	0.089*	0.060*	0.055*	0.026*
					(0.00)	(0.00)	(0.00)	(0.04)
Sharpe	0.07	0.15	0.24	0.22	0.16*	0.14*	0.09*	0.07*
					(0.00)	(0.00)	(0.02)	(0.07)
				(c) Japan				
β_P	0.04	0.14*	0.42*	0.00	0.38*	-0.04	0.28*	-0.14*
	(0.41)	(0.00)	(0.00)	(1.00)	(0.00)	(0.55)	(0.00)	(0.02)
Av. excess return	0.47	0.88	1.59	1.16	1.13	0.70	0.71	0.28
CVaR	10.54	13.50	14.56	12.97	4.02	2.43	1.06	-0.53
MtC	0.044	0.065	0.109	0.090	0.065*	0.045*	0.044*	0.024*
					(0.04)	(0.09)	(0.01)	(0.03)
Sharpe	0.09	0.15	0.22	0.20	0.12*	0.11*	0.07*	0.06*
					(0.03)	(0.05)	(0.05)	(0.03)

observe a decrease in tail risk (CVaR) of the hedged portfolio from both the unrestricted and EW benchmarks so that crashes diminish with political hedging.

We find consistent results for the Eurozone and Japanese investors. For the Eurozone, EW diversification increases political beta to 0.16, doubling to 0.31 for the unconstrained MtC portfolio, with political risk premia 1.27% and 2.46%, respectively. The unconstrained portfolio exhibits an MtC ratio increase over the index by 0.089 from 0.030 and a Sharpe increase by 0.16 from 0.07, significant at conventional levels, with smaller but statistically significant increases over EW. The performance gains remain statistically significant and economically large when hedging political risk. MtC increases by 0.060 over I and 0.026 over EW, and Sharpe increases by 0.14 (I) and 0.07 (EW). For Japanese investors, EW diversification increases political risk with beta 0.14, which triples to 0.42 for the unconstrained MtC portfolio, with corresponding political risk premia 1.11% and 3.33%. MtC increases over the index by 0.065 from 0.044 and Sharpe by 0.12 from 0.09, significant at conventional levels, with smaller but statistically significant increases over EW. The performance gains remain significant when hedging political risk. MtC increases by 0.045 over I and 0.024 over EW, and Sharpe increases by 0.11 (I) and 0.06 (EW). We also observe lower tail risk with the hedged portfolios for both investors.

Examining the portfolio weights, we note well-diversified and balanced portfolios with 4–7 assets in all cases. The exposure to emerging markets is significant, in the range of 78-86% for the unrestricted case and 33-46% for the politically hedged portfolios.

The evidence from the Eurozone and Japan is stronger than the US with its low political risk and high returns. This finding corroborates Driessen and Laeven (2007), who document diversification benefits for non-US investors and goes further to account for higher-order moments and hedge political risk.

In conclusion, political risk is not diversifiable, and performance gains from international diversification come with increased political risk exposure. However, performance gains persist when political risk is hedged. Overall, the hedged portfolios attain about 2.4% annualized higher excess returns over the equally weighted portfolios, with significantly lower tail risk, corresponding to an annualized Sharpe ratio increase of about 0.17. Next, we show that performance gains persist when also hedging currency risk.

4.3.3 Hedging currency risk

We repeat the baseline test using currency-hedged returns and report the results in Table 4. Similarly to Table 3, we find that international portfolios have increased political beta with economically significant political premia compared to the market. For USD hedged investors, diversification increases political risk with political beta 0.13 for EW and 0.11 for the unrestricted MtC portfolio, with corresponding risk premia 1.03% and 0.87% p.a. These values are lower than in our baseline but remain statistically significant and relatively large. Currency hedging reduces without eliminating political risk, affirming the significance of political risk in international portfolios.

Table 4 – Hedging political and currency risk in international portfolios

This table reports performance statistics of portfolios with hedged currency risk, namely of the MSCI home market index (I) and international portfolios using equally weighted (EW) and mean-to-CVaR unrestricted optimal portfolios (U) and with political risk hedging (H) and no short sales. Political risk is hedged with net zero exposure to the P-factor. Currency hedging is implemented using index returns converted to the local currency as discussed in the data section 4.1. Reported are the exposures to a global political risk factor β_P, and the monthly performance ratios mean-to-CVaR (MtC) and Sharpe ratio. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

	I	EW	U	H	U-I	H-I	U-EW	H-EW
(a) US								
β_P	-0.01	0.13*	0.11*	0.00	0.12*	0.01	-0.02	-0.13*
	(0.57)	(0.00)	(0.00)	(1.00)	(0.00)	(0.80)	(0.63)	(0.00)
Av. excess return	0.52	0.55	0.82	0.92	0.29	0.40	0.26	0.37
CVaR	9.90	9.80	7.63	9.74	-2.27	-0.15	-2.18	-0.06
MtC	0.053	0.056	0.107	0.094	0.054*	0.041*	0.051*	0.038*
					(0.06)	(0.05)	(0.01)	(0.01)
Sharpe	0.12	0.13	0.21	0.22	0.09	0.10*	0.07*	0.08*
					(0.13)	(0.06)	(0.05)	(0.02)
(b) Eurozone								
β_P	0.04	0.13*	0.11*	0.00	0.07	-0.04	-0.02	-0.13*
	(0.22)	(0.00)	(0.00)	(1.00)	(0.17)	(0.46)	(0.64)	(0.00)
Av. excess return	0.36	0.51	0.89	0.90	0.52	0.53	0.38	0.39
CVaR	12.16	9.90	8.34	9.69	-3.82	-2.47	-1.56	-0.22
MtC	0.030	0.051	0.106	0.093	0.076*	0.063*	0.055*	0.042*
					(0.02)	(0.01)	(0.01)	(0.01)
Sharpe	0.07	0.12	0.21	0.21	0.13*	0.13*	0.09*	0.09*
					(0.02)	(0.01)	(0.04)	(0.04)
(c) Japan								
β_P	0.04	0.13*	0.11*	0.00	0.07	-0.04	-0.02	-0.13*
	(0.41)	(0.00)	(0.00)	(1.00)	(0.28)	(0.50)	(0.56)	(0.00)
Av. excess return	0.47	0.51	0.81	0.87	0.34		0.30	0.37
CVaR	10.54	9.85	7.96	9.77	-2.58	-0.77	-1.89	-0.08
MtC	0.044	0.052	0.101	0.089	0.057*	0.045*	0.050*	0.038*
					(0.07)	(0.09)	(0.01)	(0.01)
Sharpe	0.09	0.12	0.19	0.21	0.10	0.11*	0.07*	0.08*
					(0.13)	(0.08)	(0.06)	(0.02)

We observe significant performance gains when diversifying internationally without any restrictions on political risk. For the US investor, MtC increases over the index by 0.054 from 0.056 and Sharpe by 0.09 from 0.12, with slightly smaller increases over EW. The gains remain statistically significant and economically large when also hedging political risk. MtC increases by 0.041 over I and 0.038 over EW, and Sharpe by 0.10 (I) and 0.06 (EW). Similar gains are observed for Eurozone and Japanese investors.

Examining the portfolio weights, we note diversified and balanced portfolios with 5-8 assets for the politically unhedged portfolios in all cases. When hedging both currency and political risk, there are only 2 to 3 assets. The exposure to emerging markets remains significant but lower than without currency hedging. It is 73-77% for the unrestricted case and 36-39% for the politically hedged portfolios.

In conclusion, the increased exposure to political risk persists for currency-hedged international portfolios. The performance gains from international diversification survive political and currency risk hedging.

4.4 Out-of-sample tests

We perform two out-of-sample tests. First, we use a rolling window test to mimic international equity investors hedging political and currency risk. Second, we test the effectiveness of political hedging under the unexpected political shocks from the COVID-19 pandemic and the war in Ukraine.

4.4.1 Rolling window

We mimic how an international equity investor would hedge political and currency risk. We use selective currency hedging (Black, 1989), whereby an optimal hedge is determined by exploiting covariances among exchange rates and asset returns. More recent works proxy expected currency returns and capture correlations using various methods —e.g., currency value, carry, and momentum (Barroso et al., 2022) or scenario analysis (Beltratti et al., 2004; Topaloglou et al., 2002)— and show significant gains in performance over full currency hedging. To implement selective hedging in the MtC model, we follow Topaloglou et al. (2002). We consider a set of $2n$ assets consisting of both unhedged and hedged currency random returns $\tilde{r}_u, \tilde{r}_h$, as described in the data section. The vectors of the respective portfolio weights are $x_u, x_h \in \mathbb{R}^n$, and the asset allocation is the concatenation of the two, $x \in \mathbb{R}^{2n}$, with portfolio return $\tilde{r}_p = \tilde{r}_u^\top x_u + \tilde{r}_h^\top x_h$.

We apply the model to the sample of unhedged and hedged index returns.[17] We start with the 48-month window 1999-2002 to estimate political betas at the end of 2002 (month $t = 0$) and obtain return scenarios. We run the MtC model to get a portfolio and evaluate its performance at $t+1$. We then roll the window forward to $t+1$ to obtain new return scenarios, re-estimate the betas, re-optimize the portfolio, and evaluate its performance at $t+2$. We repeat the process 204 times until the end of 2019 and compute summary statistics of the ex-post portfolio returns, accounting for transaction costs. CVaR is computed as the expected value of the extreme 5% tail losses of the 204 ex-post returns, assumed equiprobable; see the definition in Theorem A.1.

[17] For US and Japanese investors, there are 42 hedged and 41 unhedged returns, whereas, for the Eurozone, there are 31 unhedged returns since no hedging is needed for the home index returns.

We run the test with no short sales and a one-way transaction cost 0.2%. Portfolio turnover averages 11%, 13%, and 8% for the three investors with unrestricted portfolios, increasing with political hedging to 21%, 19%, and 14%, respectively; see Table 5.[18]

The unrestricted portfolios have statistically significant average political betas for all three investors. The hedged portfolios register economically and statistically significant out-of-sample MtC gains over the index, with respective increases of 0.016, 0.063, and 0.056. The gains over EW are also significant, respectivelt 0.062, 0.078, and 0.053. We also observe a decrease in the tail risk of the hedged portfolio from both the unrestricted and EW benchmarks, signifying diminished crashes with political hedging. These results align with the in-sample test of Table 3.

We use this test to assess the potential of constraining the portfolio ICRG ratings as we did in Figure 1 (right axis). There we showed that there exists a value of ICRG that gives the zero political beta portfolio, albeit we do not know a priori if a given political rating corresponds to a politically neutral portfolio. Looking at the ICRG ratings of the 204 hedged portfolios in the out-of-sample test, we find significant time variation in the 70–90 range, with a mean of 79 and a standard deviation 5.8. Hence, we can not dismiss the political beta constraint in favor of directly constraining the ratings.

In conclusion, the main result of persistent performance gains from international diversification when political risk is hedged, survives out of sample.

4.4.2 Unexpected political shocks

We take advantage of the COVID-19 lockdowns and the war in Ukraine, after our sample period of 1999-2019 of all previous tests, to perform a truly out-of-sample test under unexpected political shocks. We test the effectiveness of the hedged portfolio of our main Table 3 compared to the unrestricted during these shocks. We obtain each portfolio's 51 monthly ex-post returns from January 2020 to March 2024.[19] The results in Table 6 show that the politically-hedged portfolio delivers consistently superior results to the unrestricted, with much lower tail risk from potential crashes, offering protection from large unexpected shocks.

The results are even stronger on the narrow five-month window around these two events. The unrestricted portfolio monthly returns for the three investors were, respectively, -0.65%, -1.51%, and -0.30% during the pandemic outbreak, with corresponding politically hedged portfolio returns of 1.22%, 0.98%, and 1.11%. Likewise, the unrestricted returns around the invasion of Ukraine window are, respectively, -4.70%, -1.51%, and -3.34%, with hedged returns of -0.91%, 1.12%, and 1.61% per month. Political hedging appears offers effective protection from large crashes during periods of unexpected political uncertainty.

[18] For the performance gains over I for the US and Japan, we report p-values after removing the 2008 outlier of the great financial crisis, as in Cosset and Suret (1995). Performance gains over EW are statistically significant on the entire sample for all countries, but gains over I for the US and Japan are significant when excluding the outlier. Whereas the outliers created a problem for two instances of inference tests, they do not drive our results as we show in a robustness test in section 5.

[19] The Russian market was dropped from the MSCI index after February 28th, 2022. We use the Table 3 portfolios, including allocations to Russia, until this data and subsequently rescaled the portfolio to allocate the Russian market weights to the remaining portfolio assets.

Table 5 – Out-of-sample test

This table reports performance statistics for 204 repetitions of out-of-sample testing on a 48-month rolling window of the MSCI home market index I and international portfolios for investors in the US, Eurozone, and Japan, using equally weighted portfolios EW and mean-to-CVaR unrestricted optimal portfolios (U) and with political risk hedging (H) and no short sales. Portfolio rebalancing incurs a one-way transaction cost of 0.2%. Currency hedging is determined by the model investing selectively in the unhedged or hedged index returns converted to the local currency, as discussed in the data section 4.1. Reported are the exposures to a global political risk factor β_P averaged in absolute value over the 204 repetitions, and the monthly performance ratios mean-to-CVaR (MtC) and Sharpe ratio. The sample includes 22 developed economies and 20 emerging markets over a rolling window spanning 2003 to 2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

	I	EW	U	H	U-I	H-I	U-EW	H-EW
(a) US								
Average $\mid \beta_P \mid$	-0.08*	0.15*	0.22*	0.00	0.30*	0.08*	0.07*	-0.15*
	(0.00)	(0.00)	(0.00)	(1.00)	(0.00)	(0.00)	(0.03)	(0.00)
Avg excess return	0.80	0.44	1.08	1.06	0.28	0.26	0.64	0.62
CVaR	9.44	11.48	12.53	10.53	3.08	1.09	1.05	-0.95
Mean-to-CVaR	0.085	0.038	0.086	0.100	0.001	0.016	0.048*	0.062*
					(0.16)	(0.15)	(0.00)	(0.00)
Sharpe	0.20	0.10	0.18	0.23	-0.02	0.03	0.09*	0.13*
					(0.90)	(0.52)	(0.07)	(0.01)
(b) Eurozone								
Average $\mid \beta_P \mid$	0.02	0.15*	0.23*	0.00	0.21*	-0.02	0.08*	-0.15*
	(0.64)	(0.00)	(0.00)	(1.00)	(0.00)	(0.63)	(0.05)	(0.00)
Avg excess return	0.65	0.47	1.12	1.01	0.46	0.35	0.64	0.53
CVaR	10.54	10.22	11.21	8.07	0.67	-2.47	0.98	-2.15
Mean-to-CVaR	0.062	0.046	0.100	0.125	0.038*	0.063*	0.053*	0.078*
					(0.10)	(0.09)	(0.01)	(0.02)
Sharpe	0.14	0.12	0.20	0.24	0.06	0.10	0.09	0.12
					(0.30)	(0.24)	(0.17)	(0.12)
(c) Japan								
Average $\mid \beta_P \mid$	0.12	0.13*	0.23*	0.00	0.11	-0.12	0.10*	-0.13*
	(0.14)	(0.00)	(0.00)	(1.00)	(0.14)	(0.11)	(0.02)	(0.00)
Avg excess return	0.62	0.47	0.97	1.01	0.35	0.62	0.50	0.54
CVaR	10.99	12.39	14.51	11.16	3.53	10.99	2.12	-1.23
Mean-to-CVaR	0.056	0.038	0.067	0.091	0.010*	0.056*	0.029*	0.053*
					(0.06)	(0.08)	(0.01)	(0.03)
Sharpe	0.12	0.09	0.16	0.21	0.04	0.08	0.07*	0.11*
					(0.17)	(0.14)	(0.09)	(0.04)

4.5 Political hedging and the equity home bias puzzle

We finally use the model to study the effects of political hedging on the equity home bias puzzle (French and Poterba, 1991). We proceed in two steps to reconcile Dahlquist et al. (2003), who

Table 6 – Out-of-sample political hedging from unexpected political shocks

This table reports performance statistics for 51 ex-post returns out-of-sample of the international portfolios from Table 3 for investors in the US, Eurozone, and Japan during the period spanning the COVID-19 and war in Ukraine. It shows results using mean-to-CVaR unrestricted optimal portfolios (U), political risk hedging (H), and no short sales. Reported are the monthly performance ratios mean-to-CVaR (MtC) and Sharpe ratio. The sample includes 22 developed economies and 20 emerging markets spanning January 1, 2020, to March 31, 2024, excluding the Russian market after February 28, 2022. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

	(a) US		(b) Eurozone		(c) Japan	
	U	H	U	H	U	H
Av. excess return	-0.04	0.98	0.65	1.55	0.78	1.96
CVaR	18.72	13.08	16.02	8.58	18.87	10.54
MtC	-0.002	0.075	0.040	0.181	0.041	0.186
MtC gains over U	–	0.077*	–	0.141*	–	0.145*
		(0.00)		(0.00)		(0.01)
Sharpe	-0.01	0.15	0.10	0.33	0.11	0.34
Sharpe gains over U	–	0.16*	–	0.23*	–	0.23*
		(0.01)		(0.01)		(0.01)

show that country political risk can tilt portfolios towards the home market, with Guidolin and Timmermann (2008), who point out that political risk applies more to emerging markets and is a less obvious explanation of limited diversification among developed economies.

We first focus on emerging markets with high political risk (Bekaert and Harvey, 2003; Diamonte et al., 1996). The political efficient frontier for emerging markets lies below those in Figure 4 (not shown) and does not extend to the zero political beta portfolios. We run tests for three points on the home and emerging markets frontier, with β_P reduced from its unconstrained value to 0.30 or 0.20; see Table 7. Comparing with Table 3, we observe that the unrestricted portfolios have larger political beta and somewhat lower performance. This is expected as we diversify into a smaller country sample with higher political risk. Still, performance gains persist when the political beta is constrained.[20]

For US investors, reducing political beta from the unconstrained 0.45 to 0.20 preserves diversification benefits, with MtC higher by 0.047 from I and 0.024 from EW. The results are stronger for the Eurozone, with a reduction of political beta from 0.49 to 0.20, having an MtC of 0.079 higher than I and 0.028 than EW. Reducing Japan's political beta from 0.45 to 0.20 achieves MtC higher by 0.054 and 0.018 over I and EW, respectively. The benefits from diversifying solely in emerging markets erode for lower values of political beta, with no statistically significant performance gains for $\beta_P = 0.10$ or lower.

Comparing the baseline with the emerging markets tests sheds light on the effect of political risk on the equity home bias puzzle. Looking at the portfolio weights underlying the results of Table 7 we find that lowering the target political beta reduces the allocation to emerging markets. For a very low $\beta_P = 0.05$ (not shown), the home allocations increase to 52%, 18%, and 89% for the three investors, respectively. This result corroborates Dahlquist et al. (2003) that political risk aversion is a factor for equity home bias.

[20]We do not report Sharpe ratios for the highly skewed emerging markets, but the results are consistent.

Table 7 – Managing political risk in emerging markets

This table reports performance statistics of portfolios diversified into emerging markets at four points of the MtC-β_P political efficient frontier with limits on the exposure to the global political risk factor. Results are obtained with the MtC model and no short sales. The monthly performance ratios mean-to-CVaR (MtC) are reported, and the gains over the index I and EW portfolios from Table 3. "U" denotes the unconstrained portfolio at the peak of the political frontier. The sample includes the home country and 20 emerging markets, spanning 1999 to 2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

	(a) US			(b) Eurozone			(c) Japan		
	U	Limited β_P		U	Limited β_P		U	Limited β_P	
β_P	0.45*	0.30*	0.20*	0.49*	0.30*	0.20*	0.45*	0.30*	0.20*
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Av. excess return	1.49	1.26	1.12	1.57	1.19	1.06	1.62	1.35	1.22
CVaR	14.26	12.11	11.17	13.40	10.46	9.77	14.92	12.85	12.51
MtC	0.104	0.104	0.100	0.117	0.114	0.109	0.108	0.105	0.098
MtC gains over I	0.052*	0.051*	0.047	0.087*	0.084*	0.079*	0.064*	0.061*	0.054*
	(0.09)	(0.09)	(0.11)	(0.01)	(0.01)	(0.01)	(0.04)	(0.05)	(0.07)
MtC gains over EW	0.028*	0.027*	0.024*	0.037*	0.034*	0.028*	0.029*	0.025*	0.018*
	(0.08)	(0.04)	(0.07)	(0.07)	(0.02)	(0.04)	(0.04)	(0.04)	(0.10)

We also note significant changes in the optimal portfolio composition underlying Table 3 when hedging political risk. The aggregate exposure to developed markets increases from 22% with unrestricted to 54% with hedged portfolios for the US, from 21% to 67% for the Eurozone, and from 14% to 57% for Japan. Likewise, computing the average exposure to developed markets during the out-of-sample test, we find that it increases from 19% with unrestricted to 38% with hedged portfolios for the US, from 26% to 30% for the Eurozone, and from 18% to 38% for Japan. Hedging political risk tilts international portfolios away from politically risky countries, but the tilt is from emerging toward developed countries and not necessarily toward the home. Our finding empirically supports the tilt in the direction anticipated by Guidolin and Timmermann (2008).

5 Robustness tests

We successfully perform a battery of robustness tests to (i) establish the robustness of the model to tail risk estimation and to data perturbations and rule out that outliers drive our results; (ii) rule out serendipitous results using a randomized test; (iii) show that transaction costs do not significantly alter the political risk of international portfolios or the efficacy of political hedging; (iv) show that political risk can be significantly reduced if short sales are possible with the model of higher-order moments. We describe each test and how it corroborates our findings, relegating the results to online Appendix B.

(i) Outliers and model robustness. One potential concern is that given the 0.95 CVaR risk measure we use, a few tail events may drive the results since CVaR is not a robust statistic. This is not usually of practical concern for large datasets, as in our case. Also, the fact that our results hold with Sharpe ratios suggests that outliers are not a problem. Nevertheless, we perform additional tests for different αs or scenario generation procedures. Specifically, we solve the model with αs in the range 0.90 to 0.99; we obtain scenarios from historical data but excluding the great financial crisis; we obtain scenarios using an in-sample randomization of returns. The results (see online Appendix B.1) confirm that our findings are robust to outliers, and the model is robust to data perturbations.

(ii) Randomized test. We perform an out-of-sample test using randomization (see online Appendix B.1). We block bootstrap the time series of our data, with replacement, to generate 1000 samples of market and benchmark portfolio returns and test the performance of the optimal in-sample politically hedged portfolio. In 98% of the runs (results not shown), the portfolio MtC outperforms the EW benchmark, and in 94%, it beats the index. These results rule out that the hedging model's success is due to chance.

(iii) Transaction costs. We introduce proportional transaction costs to assess whether diversification benefits may disappear when hedging political risk due to these costs. The results (see online Appendix B.2) show that the politically hedged portfolio registers economically and statistically significant MtC gains over both benchmark portfolios, with transaction costs as high as 0.2% for developed and 0.5% for emerging markets. Higher trading costs imply smaller gains, but the gains of politically hedged portfolios remain statistically significant for reasonable transaction costs.

(iv) Short positions. We consider short positions in developed markets. We find (see online Appendix B.3) that the unrestricted MtC portfolio political beta is not statistically significant for the US and Japanese investors, and it is half the beta of the unrestricted NSS portfolios for Eurozone investors. The optimal portfolios have long-short positions that hedge political risk, but this results from optimizing tail risk and is not achieved with Sharpe ratio maximization. Hence, political risk can be diversified away if short sales are allowed in the case of the US and Japan and are significantly reduced from the no-short-sales case for the Eurozone. Consistently with our main result, the gains from international diversification persist for politically neutral portfolios.

6 Conclusions

International diversification exposes investors to political risk, increasing the tail risk of potential crashes. We develop a mean-to-CVaR portfolio selection model accounting for the skewed return distributions of the international markets with high political risk and a political beta constraint. The optimal portfolios satisfy second-order stochastic dominance, and we derive political efficient frontiers for managing and hedging political risk for investors with non-decreasing utility functions. We also develop an asymptotic valid inference test to compare the optimal portfolios. The mean-to-CVaR model is computationally tractable and is endowed with an implementable inference test algorithm, so it is useful for other financial applications where deviations from normality can be a significant concern, such as ESG investing, firm announcements or crashes, actuarial risks, regime-switching processes, or political risk in the currency markets.

We apply the model to hedge political risk in international equity portfolios in a large sample of developed and emerging markets and use the inference test to draw conclusions. Our main empirical finding is that internationally diversified portfolios are exposed to political risk even when currency risk is hedged. Importantly, hedging political risk does not eliminate the diversification benefits and reduces tail risk from potential crashes. Existing literature finds that currency hedging does not eliminate international diversification benefits, and we show that it is possible to hedge another major source of risk (political) while preserving significant gains. These findings hold for US, Eurozone, and Japanese investors, in and out of sample, and are robust to outliers or alternative model specifications and transaction cost frictions. The model also effectively hedges the unexpected political shocks that followed the COVID-19 pandemic and the war in Ukraine.

Political hedging tilts international portfolios away from politically risky countries, but the tilt is away from emerging into developed countries rather than toward the home. This finding provides empirical support that hedging political risk induces bias in the direction anticipated by some home bias literature but does not resolve the puzzle.

Disclosure statement

The authors report there are no competing interests to declare.

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Data Appendix

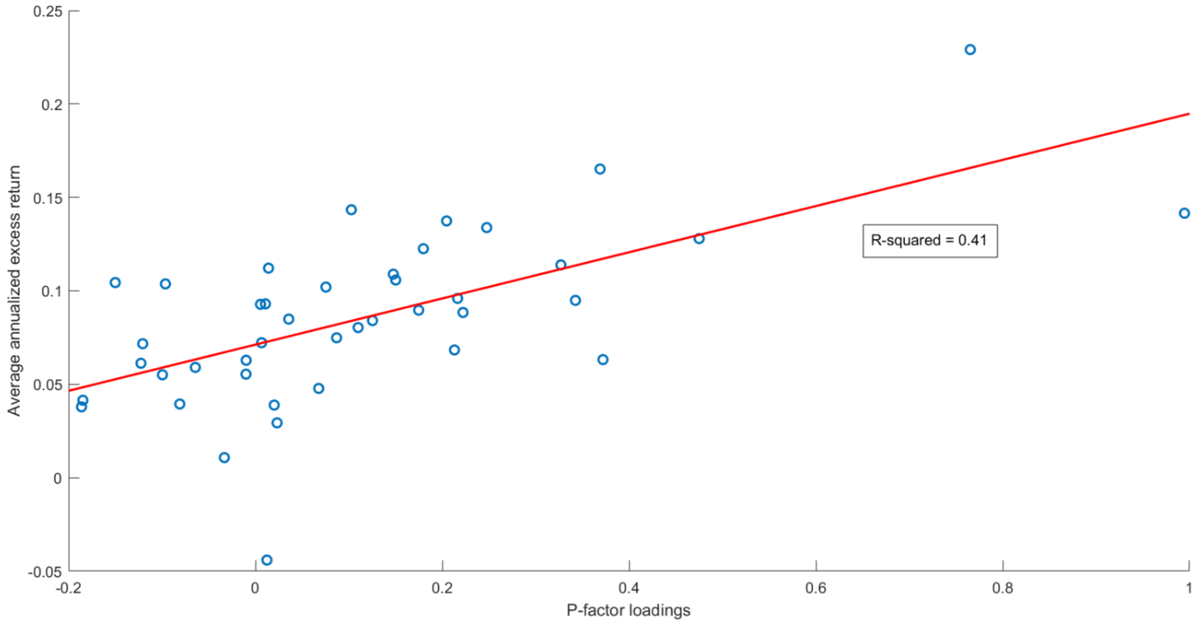
This table reports descriptive statistics for all countries in our sample, respectively, mean, standard deviation, skewness, excess kurtosis, Value-at-Risk, and Conditional-Value-at-Risk for the monthly series of each country's excess returns, denominated in USD, over the US one-month T-Bill rate. "MtC" and "Sharpe" denote the monthly mean-to-CVaR for each country's excess returns and the Sharpe ratio. VaR, CVaR, and MtC are computed at the 5% confidence level. "ICRG" is the average over time of the aggregate rating from the International Country Risk Guide. The sample period spans January 1, 1999, to December 31, 2019. All statistics are reported at a monthly frequency. Mean, StdDev, VaR, and CVaR are in percentage points.

Country	Mean	StdDev	Skew	Kurt	VaR	CVaR	MtC	Sharpe	ICRG
Australia	0.77	5.98	-0.54	1.99	8.35	13.77	0.06	0.13	85.62
Austria	0.60	6.81	-0.87	4.32	9.45	15.91	0.04	0.09	85.43
Belgium	0.35	6.00	-1.22	5.60	9.46	15.09	0.02	0.06	81.28
Brazil	1.38	10.55	-0.04	1.16	14.06	21.93	0.06	0.13	65.50
Canada	0.71	5.61	-0.53	2.62	8.39	12.09	0.06	0.13	86.61
Chile	0.67	6.26	-0.23	1.34	9.15	13.24	0.05	0.11	76.39
China	0.85	8.21	0.41	3.98	13.07	17.24	0.05	0.10	63.28
Colombia	1.15	8.20	-0.16	0.26	12.88	16.34	0.07	0.14	57.44
Czech Republic	1.02	7.43	-0.09	1.24	10.59	15.39	0.07	0.14	77.84
Denmark	0.87	5.70	-0.73	2.69	9.38	13.63	0.06	0.15	84.17
Egypt	0.79	8.93	0.07	2.14	13.41	18.50	0.04	0.09	58.12
Finland	0.60	8.11	0.10	2.07	13.42	18.13	0.03	0.07	90.82
France	0.49	5.80	-0.46	0.99	10.58	13.62	0.04	0.08	75.83
Germany	0.46	6.50	-0.37	1.64	10.25	15.48	0.03	0.07	84.50
Greece	-0.37	10.55	-0.23	0.68	18.01	24.24	-0.02	-0.03	73.47
Hong-Kong	0.70	6.04	-0.17	1.46	9.77	13.12	0.05	0.12	78.42
Hungary	0.88	9.16	-0.51	2.19	14.60	21.38	0.04	0.10	77.76
India	1.12	8.28	-0.02	2.04	13.22	17.38	0.06	0.13	60.23
Ireland	0.32	6.49	-0.70	1.94	11.78	16.44	0.02	0.05	85.97
Israel	0.62	6.26	-0.23	1.38	10.55	14.06	0.04	0.10	64.40
Italy	0.24	6.61	-0.22	0.58	11.20	14.70	0.02	0.04	76.80
Japan	0.32	4.77	-0.12	0.33	7.98	9.91	0.03	0.07	81.93
Malaysia	0.75	5.78	0.63	4.58	9.01	11.37	0.07	0.13	72.11
Mexico	0.80	6.67	-0.50	1.58	10.62	14.55	0.05	0.12	68.29
Netherlands	0.46	5.76	-0.71	1.94	9.65	14.05	0.03	0.08	86.77
New Zealand	0.93	5.74	-0.44	0.79	8.72	12.55	0.07	0.16	87.85
Norway	0.86	7.28	-0.65	2.79	9.39	16.38	0.05	0.12	88.14
Peru	1.19	7.64	-0.28	2.14	11.51	15.72	0.08	0.16	63.30
Philippines	0.57	6.95	-0.02	0.97	11.08	14.56	0.04	0.08	63.25
Poland	0.74	9.11	-0.10	0.79	13.16	18.98	0.04	0.08	77.32
Portugal	0.09	6.30	-0.33	0.82	10.03	13.97	0.01	0.01	81.20
Russia	1.91	10.59	0.55	3.44	15.09	20.26	0.09	0.18	60.55
South Africa	0.91	7.14	-0.31	0.10	10.62	14.36	0.06	0.13	66.35
South Korea	0.95	8.50	0.20	0.92	13.94	16.61	0.06	0.11	76.87
Spain	0.40	6.70	-0.14	1.04	10.08	14.31	0.03	0.06	75.74
Sweden	0.78	6.98	-0.15	1.93	11.70	16.00	0.05	0.11	88.15
Switzerland	0.51	4.43	-0.46	0.62	7.37	10.35	0.05	0.12	88.34
Taiwan	0.53	7.24	0.09	1.10	11.10	15.07	0.03	0.07	78.05
Thailand	1.07	8.47	-0.01	2.92	11.46	18.95	0.06	0.13	61.31
Turkey	1.18	13.51	0.53	3.12	17.10	27.07	0.04	0.09	57.75
UK	0.33	4.67	-0.38	1.45	7.22	10.17	0.03	0.07	82.89
US	0.52	4.33	-0.64	1.02	7.85	9.84	0.05	0.12	82.82
EW portfolio	0.71	5.37	-0.65	5.68	8.38	12.04	0.06	0.13	75.69

Appendix

Figure 1 – Country loadings on the political risk factor

This figure illustrates the positive relation between factor loadings on the P-factor and country average excess returns per annum. Factor loadings are estimated from an asset pricing model that controls for market and political risks. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019.



Online Appendix

Hedging political risk in international portfolios

S. Lotfi. G. Pagliardi, E. Paparoditis, S.A. Zenios.

A Background results and proofs

We start with some background results relating to CVaR and stochastic dominance. We point out that Rockafellar and Uryasev (2002) develop their model for a random loss variable $\tilde{z}$ and not for returns, as we do in our Definition A.1. Their CVaR of losses is the expected value *above* a threshold ζ. In contrast, we take the CVR of return as the negative of expected value *below* the $1-\alpha$ probability threshold ζ. We use their results with $\tilde{z} = -\tilde{r}_p$, and for simplicity, we drop the confidence level subscript.

Definition A.1 (Conditional Value-at-Risk, CVaR)**.** *The conditional Value-at-Risk at confidence level $\alpha \in (0,1)$ for a continuously distributed random portfolio return $\tilde{r}_p$ is*

$$\mathrm{CVaR}_\alpha(\tilde{r}_p) = -\mathbb{E}[\tilde{r}_p \mid \tilde{r}_p \leq \zeta], \tag{A1}$$

where $\mathbb{E}$ is the expectation operator and $\zeta \in \mathbb{R}$ is the Value-at-Risk, i.e., the $(1-\alpha)$-quantile of $\tilde{r}_p$ given by the highest γ such that $\tilde{r}_p$ will not exceed γ with probability $1-\alpha$,

$$\mathrm{VaR}_{1-\alpha}(\tilde{r}_p) \doteq \zeta = \max\{\gamma \in \mathbb{R} \mid \mathrm{Prob}(\tilde{r}_p \leq \gamma) \leq 1-\alpha\}. \tag{A2}$$

Theorem A.1 (Fundamental minimization formula)**.** (Rockafellar and Uryasev, 2002) *As a function of $\gamma \in \mathbb{R}$, the auxiliary function at confidence level $\alpha \in (0,1)$,*

$$F(\tilde{r}_p, \gamma) = \gamma + \frac{1}{1-\alpha}\mathbb{E}\big[\max\{-\tilde{r}_p - \gamma, 0\}\big] \tag{A3}$$

is finite and convex, with

$$\mathrm{CVaR}(\tilde{r}_p) = \min_{\gamma \in \mathbb{R}} F(\tilde{r}_p, \gamma). \tag{A4}$$

Definition A.2 (Stochastic dominance)**.** *Random variable $\tilde{X}$ dominates random variable $\tilde{Y}$ under first order stochastic dominance (FSD, $\tilde{X} \succeq_{FSD} \tilde{Y}$) if $\mathbb{E}(U(\tilde{X})) \geq \mathbb{E}(U(\tilde{Y}))$ for all non-decreasing utility functions U. Similarly, $\tilde{X}$ dominates random variable $\tilde{Y}$ under second-order stochastic dominance (SSD, $\tilde{X} \succeq_{SSD} \tilde{Y}$) if $\mathbb{E}(U(\tilde{X})) \geq \mathbb{E}(U(\tilde{Y}))$ for all non-decreasing concave utility functions U. (Ogryczak and Ruszczyński, 2002)*

Definition A.3 (Risk measure consistency)**.** *Given a stochastic order $\succeq_{SSD}$ we say that a risk measure ρ is SSD consistent if $\tilde{X} \succeq_{SSD} \tilde{Y}$ implies $\rho(\tilde{X}) \leq \rho(\tilde{Y})$. (Ogryczak and Ruszczyński, 2002)*

A.1 Second-order stochastic dominance consistency of MtC

Theorem A.2. *Let $\mathbb{X}_+$ denote the space of all feasible portfolios with a positive numerator and denominator of the MtC ratio. Then* MtC *is SSD consistent for all portfolios in $\mathbb{X}_+$.*

Proof. We use the notions of stochastic dominance of random variables and SSD consistency of risk measures; see Definitions A.2–A.3. Let us assume that portfolios x_1 and x_0 belong to $\mathbb{X}_+$, and x_1 dominates x_0, i.e., the portfolios' excess returns satisfy $\tilde{r}_{e,x_1} \succeq_{SSD} \tilde{r}_{e,x_0}$. This implies that $E(\tilde{r}_{e,x_1}) \geq E(\tilde{r}_{e,x_0}) > 0$ (Whang, 2019, Theorem 1.1.5). We also have that CVaR is SSD consistent (Ogryczak and Ruszczyński, 2002, Theorem 3.2), which implies $0 < \mathrm{CVaR}(\tilde{r}_{e,x_1}) \leq \mathrm{CVaR}(\tilde{r}_{e,x_0})$. Therefore, the ratio of CVaR-to-mean for portfolio x_1 is less than or equal to CVaR-to-mean for portfolio x_0. Hence, the inverse of the MtC ratio is consistent with SSD. Since we assume the risk measure $\rho(\cdot)$ to be positive, we replace $\rho(\tilde{X}) \leq \rho(\tilde{Y})$ with $\frac{1}{\rho(\tilde{Y})} \leq \frac{1}{\rho(\tilde{X})}$ and MtC is SSD consistent. □

A.2 Linear programming formulations for MtC optimization

First, we provide the linear programming formulation for the NSS constraint set (1) and then generalize to the LS constraint set (2). For NSS, we arrive at the linear program of Stoyanov et al. (2007), but we give it for completeness to set the stage for the proof of the LS model.

Theorem A.3 (MtC optimization with no short sales)**.** *Assuming positive CVaR on excess returns of every portfolio in the feasible set (1), the MtC maximization is expressed as*

$$\begin{aligned} \max_{x'\in\mathbb{R}^n,\, u'\in\mathbb{R}^S,\, \gamma'\in\mathbb{R}} \quad & (\bar{r} - r_f e)^\top x' \\ s.t. \quad & \gamma' + \frac{1}{S(1-\alpha)} e^\top u' = 1 \\ & -R_e x' - u' - e\gamma' \leq 0 \\ & e^\top x' > 0 \\ & u',\, x' \geq 0, \end{aligned} \tag{A5}$$

where R_e denotes the $S \times n$ matrix of excess returns. Given x'^*, the optimal solution of (A5), we obtain the optimal solution of (8) as $x^* = \frac{1}{e^\top x'^*} x'^*$.

Proof. From Theorem A.1 the CVaR of portfolio x is the optimal value of the linear program

$$\begin{aligned} \min_{u\in\mathbb{R}^S,\, \gamma\in\mathbb{R}} \quad & \gamma + \frac{1}{S(1-\alpha)} e^\top u \\ \text{s.t.} \quad & -u - e\gamma \leq Rx \\ & u \geq 0, \end{aligned} \tag{A6}$$

where e is an n-dimensional vector of 1. Given the assumption that at the optimal solution $\mathrm{CVaR}^*(\tilde{r}_e) \geq \delta > 0$, we can find a neighborhood for which $\mathrm{CVaR}(\tilde{r}_e) > 0$. We define $\xi =$

$\text{CVaR}(\tilde{r}_e) > 0$, and break the objective function (8) in two components to obtain

$$\begin{aligned} \max_{x\in\mathbb{X},\, \xi\in\mathbb{R}} \quad & \bar{r}^\top \frac{x}{\xi} - r_f \frac{1}{\xi} \\ \text{s.t.} \quad & \text{CVaR}(\tilde{r}_e) = \xi \\ & \xi > 0. \end{aligned} \tag{A7}$$

Using the definition of CVaR from (A6) and setting $x' = \frac{x}{\xi}$, $\nu = \frac{1}{\xi}$, $u' = \frac{u}{\xi}$, and $\gamma' = \frac{\gamma}{\xi}$, we rewrite the MtC maximization model as

$$\begin{aligned} \max_{x'\in\mathbb{R}^n,\, u'\in\mathbb{R}^S,\, \gamma',\, \nu\in\mathbb{R}} \quad & \bar{r}^\top x' - r_f \nu \\ \text{s.t.} \quad & \gamma' + \frac{1}{S(1-\alpha)} e^\top u' = 1 \\ & -R_e x' - u' - e\gamma' \leq 0 \\ & e^\top x' = \nu \\ & u' \geq 0,\, x' \geq 0,\, \nu > 0. \end{aligned} \tag{A8}$$

Substituting $e^\top x'$ for $\nu > 0$ in the objective function by the constraint $e^\top x' > 0$, we get (A5), completing the proof. □

We now give the proof for the MtC linear program with short sales of Theorem 3.1.

Proof. Following the same procedure as in Theorem A.3, we can define $\xi = \text{CVaR}(\tilde{r}_p - r_f) > 0$, and break the objective function from (8) in two component as below:

$$\begin{aligned} \max_{x\in\mathbb{X}_S,\, \xi\in\mathbb{R}} \quad & \bar{r}^\top \frac{x_+}{\xi} - \bar{r}^\top \frac{x_-}{\xi} - r_f \frac{1}{\xi} \\ \text{s.t.} \quad & \text{CVaR}(\tilde{r}_e) = \xi \\ & \xi > 0. \end{aligned} \tag{A9}$$

Setting $x'_+ = \frac{x_+}{\xi}$, $x'_- = \frac{x_-}{\xi}$, $\nu = \frac{1}{\xi}$, $u' = \frac{u}{\xi}$ and $\gamma' = \frac{\gamma}{\xi}$, we have

$$\begin{aligned} \max_{x'_+,\, x'_-\in\mathbb{R}^n,\, u'\in\mathbb{R}^S,\, \gamma'\in\mathbb{R}} \quad & \bar{r}^\top x'_+ - \bar{r}^\top x'_- - r_f \nu \\ \text{s.t.} \quad & \gamma' + \frac{1}{S(1-\alpha)} e^\top u' = 1 \\ & -R_e x'_+ + R_e x'_- - u' - e\gamma' \leq 0 \\ & e^\top x'_+ - e^\top x'_- = \nu \\ & e^\top x'_- \leq \nu \\ & u' \geq 0,\, x'_+,\, x'_- \geq 0,\, \nu > 0. \end{aligned} \tag{A10}$$

Substituting $e^\top(x'_+ - x'_-)$ for ν ($\nu > 0$) in the objective function, while adding $e^\top(x'_+ - x'_-) > 0$ constraint, we get (11), completing the proof. □

A.3 Proof of Theorem 3.2

We give some notations to proceed with the proof. For a sequence $\{X_n\}$ of random variables defined on the same probability space, $X_n = o_P(1)$ states for converge of $\{X_n\}$ to zero in probability and $X_n = O_P(1)$ for boundedness of $\{X_n\}$ in probability. We write $X_n \xrightarrow{P} X$ for convergence of $\{X_n\}$ to X in probability and $X_n \xrightarrow{d} X$ for convergence in distribution. For a random variable X, let $X^- = -\min\{0, X\}$ and $X^+ = \max\{0, X\}$.

Let

$$\widetilde{R}_t = \big(r_{1,t}, -\frac{1}{1-\alpha} r_{1,t}\mathbb{1}(r_{1,t} \le \widehat{\zeta}_{1,1-\alpha}), r_{0,t}, -\frac{1}{1-\alpha} r_{0,t}\mathbb{1}(r_{0,t} \le \widehat{\zeta}_{0,1-\alpha})\big)^\top,$$

and consider the sequence $\{\widetilde{Y}_S, S \in \mathbb{N}\}$ where $\widetilde{Y}_S = (1/S)\sum_{t=1}^S \widetilde{R}_t$.

Let $\mu = (\mu_1, \mathrm{CVaR}_1, \mu_0, \mathrm{CVaR}_0)^\top$, where for $j \in \{0, 1\}$,

$$\mu_j = E(r_{j,t}), \quad \mathrm{CVaR}_j = -\frac{1}{1-\alpha} E\big(r_{j,t}\mathbb{1}(r_{j,t} \le \zeta_{j,1-\alpha})\big).$$

We first show that

$$\sqrt{S}(\widetilde{Y}_S - \mu) = \sqrt{S}(\overline{Y}_S - \mu) + o_P(1), \tag{A11}$$

where $\overline{Y}_S = (1/S)\sum_{t=1}^S R_t$ with

$$R_t = \big(r_{1,t}, -\frac{1}{1-\alpha} r_{1,t}\mathbb{1}(r_{1,t} \le \zeta_{1,1-\alpha}), r_{0,t}, -\frac{1}{1-\alpha} r_{0,t}\mathbb{1}(r_{0,t} \le \zeta_{0,1-\alpha})\big)^\top.$$

Equation (A11) follows if we show that

$$\sqrt{S}(\widehat{\mathrm{CVaR}}_j - \mathrm{CVaR}_j) = \frac{1}{\sqrt{S}(1-\alpha)} \sum_{t=1}^S \big\{(r_{j,t} - \zeta_{j,1-\alpha})^- - E(r_{j,t} - \zeta_{j,1-\alpha})^-\big\} + o_P(1), \tag{A12}$$

holds. To simplify notation, notice first that for $\ell_{j,t} = -r_{j,t}$, we have that

$$-\frac{1}{1-\alpha} E\left(r_{j,t}\mathbb{1}(r_{j,t} \le \zeta_{j,1-\alpha})\right) = \frac{1}{1-\alpha} E\left(\ell_{j,t}\mathbb{1}(\ell_{j,t} \ge v_{j,\alpha})\right)$$

and

$$-\frac{1}{1-\alpha}\frac{1}{S}\sum_{t=1}^S r_{j,t}\mathbb{1}(r_{j,t} \le \widehat{\zeta}_{j,1-\alpha}) = \frac{1}{1-\alpha}\frac{1}{S}\sum_{t=1}^S \ell_{j,t}\mathbb{1}(\ell_{j,t} \ge \widehat{v}_{j,\alpha}),$$

where $v_{j,\alpha} = \inf\{x : P(\ell_{j,t} \le x) \ge \alpha\}$ and $\widehat{v}_{j,\alpha} = \ell_{\lfloor S\alpha \rfloor}$ is the empirical α quantile of the sample $\ell_{j,1}, \ell_{j,2}, \ldots, \ell_{j,S}$. Assertion (A12) is then equivalent to

$$\sqrt{S}(\widehat{\mathrm{CVaR}}_j - \mathrm{CVaR}_j) = \frac{1}{\sqrt{S}(1-\alpha)} \sum_{t=1}^S \big\{(\ell_{j,t} - v_{j,\alpha})^+ - E(\ell_{j,t} - v_{j,\alpha})^+\big\} + o_P(1). \tag{A13}$$

To establish (A13), we follow Kolla et al. (2019) and write $\widehat{\text{CVaR}}_j$ as

$$\begin{aligned}\widehat{\text{CVaR}}_j &= \widehat{v}_{j,\alpha} + \frac{1}{S(1-\alpha)}\sum_{t=1}^{S}(\ell_{j,t}-\widehat{v}_{j,\alpha})\mathbb{1}(\ell_{j,t}\geq \widehat{v}_{j,\alpha}) \\ &= v_{j,\alpha} + \frac{1}{S(1-\alpha)}\sum_{t=1}^{S}(\ell_{j,t}-v_{j,\alpha})\mathbb{1}(\ell_{j,t}\geq v_{j,\alpha}) + e_S \\ &= \text{CVaR}_j + \frac{1}{S(1-\alpha)}\sum_{t=1}^{S}\Big\{(\ell_{j,t}-v_{j,\alpha})\mathbb{1}(\ell_{j,t}\geq v_{j,\alpha}) \\ &\qquad - E(\ell_{j,t}-v_{j,\alpha})\mathbb{1}(\ell_{j,t}\geq v_{j,\alpha})\Big\} + e_S,\end{aligned}$$

where

$$\begin{aligned}e_S = &\frac{\widehat{v}_{j,\alpha}-v_{j,\alpha}}{1-\alpha}\big(\widehat{F}_{j,S}(\widehat{v}_{j,\alpha})-\alpha\big) \\ &+ \frac{1}{S(1-\alpha)}\sum_{t=1}^{S}(\ell_{j,t}-v_{j,\alpha})\big[\mathbb{1}(\ell_{j,t}\geq \widehat{v}_{j,\alpha}) - \mathbb{1}(\ell_{j,t}\geq v_{j,\alpha})\big]\end{aligned}$$

and $\widehat{F}_{j,S}$ denotes the empirical distribution function of $\ell_{j,t}$, $t = 1, 2, \ldots, S$. From the above derivations, we get

$$\sqrt{S}(\widehat{\text{CVaR}}_j - \text{CVaR}_j) = \frac{1}{\sqrt{S}(1-\alpha)}\sum_{t=1}^{S}\Big\{(\ell_{j,t}-v_{j,\alpha})^+ - E(\ell_{j,t}-v_{j,\alpha})^+\Big\} + \sqrt{S}e_S,$$

and in order to establish (A13) it suffices to show that $\sqrt{S}e_S = o_P(1)$. For this, verify first that

$$\begin{aligned}\Big|\frac{1}{S(1-\alpha)}\sum_{t=1}^{S}(\ell_{j,t}-v_{j,\alpha})\big[\mathbb{1}(\ell_{j,t}\geq \widehat{v}_{j,\alpha}) - \mathbb{1}(\ell_{j,t}\geq v_{j,\alpha})\big]\Big| \\ \leq \frac{|\widehat{v}_{j,\alpha}-v_{j,\alpha}|}{1-\alpha}\big|\widehat{F}_{j,n}(\widehat{v}_{j,\alpha}) - \widehat{F}_{j,n}(v_{j,\alpha})\big|,\end{aligned}$$

and therefore,

$$\begin{aligned}|\sqrt{S}e_S| \leq &\frac{|\sqrt{S}(\widehat{v}_{j,\alpha}-v_{j,\alpha})|}{1-\alpha}\big|\widehat{F}_{j,S}(\widehat{v}_{j,\alpha})-\alpha\big| \\ &+ \frac{|\sqrt{S}(\widehat{v}_{j,\alpha}-v_{j,\alpha})|}{1-\alpha}\big|\widehat{F}_{j,n}(\widehat{v}_{j,\alpha}) - \widehat{F}_{j,n}(v_{j,\alpha})\big|.\end{aligned} \tag{A14}$$

Notice that under Assumption (ii), $\sqrt{S}(\widehat{v}_{j,\alpha}-v_{j,\alpha}) = O_P(1)$, that is, $\widehat{v}_{j,\alpha} \xrightarrow{P} v_{j,\alpha}$; see Lemma 5.1 of Sun and Lahiri (2006). Furthermore,

$$\begin{aligned}|\widehat{F}_{j,S}(\widehat{v}_{j,\alpha})-\alpha| &\leq |\widehat{F}_{j,S}(\widehat{v}_{j,\alpha}) - F_j(\widehat{v}_{j,\alpha})| + |\widehat{F}_j(\widehat{v}_{j,\alpha})-\alpha| \\ &\leq \sup_{x\in\Re}|\widehat{F}_{j,S}(x) - F_j(x)| + |\widehat{F}_j(\widehat{v}_{j,\alpha})-\alpha|.\end{aligned}$$

Since $\sup_{x\in\Re} |\widehat{F}_{j,S}(x) - F_j(x)| \xrightarrow{P} 0$, see Dehling and Philipp (2002), the first term, goes to zero in probability. For the second term observe that $\widehat{v}_{j,\alpha} \xrightarrow{P} v_{j,\alpha}$ implies by the continuity of F_j that $|\widehat{F}_j(\widehat{v}_{j,\alpha}) - \alpha| \xrightarrow{P} 0$. Thus, the first term on the right-hand side of the bound given in (A14) converges to zero in probability as $S \to \infty$. For the second term of the same bound, we have

$$\begin{aligned}
\big|\widehat{F}_{j,n}(\widehat{v}_{j,\alpha}) - \widehat{F}_{j,n}(v_{j,\alpha})\big| &\leq \big|\widehat{F}_{j,n}(\widehat{v}_{j,\alpha}) - F_{j,n}(\widehat{v}_{j,\alpha})\big| + \big|\widehat{F}_{j,n}(v_{j,\alpha}) - F_{j,n}(v_{j,\alpha})\big| \\
&\quad + \big|F_j(\widehat{v}_{j,\alpha}) - F_j(v_{j,\alpha})\big| \\
&\leq 2\sup_{x\in\Re} \big|\widehat{F}_{j,n}(x) - F_{j,n}(x)\big| + \big|F_j(\widehat{v}_{j,\alpha}) - F_j(v_{j,\alpha})\big|,
\end{aligned}$$

which converges to zero in probability as $S \to \infty$, again, by the uniform consistency of $\widehat{F}_{j,S}$ as an estimator of F_j, the fact that $\widehat{v}_{j,\alpha} \xrightarrow{P} v_{j,\alpha}$ and the continuity of the distribution function F_j.

The previous derivations have shown that $\sqrt{S}e_S = o_P(1)$ and therefore that (A13) holds, from which we conclude the proof of assertion (A11). According to this assertion, the limiting distribution of $\sqrt{S}(\widetilde{Y}_S - \mu)$ is the same as the limiting distribution of $\sqrt{S}(\overline{Y}_S - \mu)$. To establish the limiting distribution of the last mentioned sequence, we get from Assumption (i) that, as $S \to \infty$,

$$\sqrt{S}(\overline{Y}_S - \mu) \xrightarrow{d} \mathcal{N}(0, \Sigma_r), \tag{A15}$$

where

$$\Sigma_r = \sum_{h=-\infty}^{\infty} \mathrm{E}\big((R_t - \mathrm{E}(R_t))(R_{t+h} - \mathrm{E}(R_{t+h}))^\top\big).$$

To proceed with the limiting distribution of the test statistic T_S, observe first that $T_S = g(\overline{Y}_S)$, where the function $g : \Re^4 \to \Re$ is defined by $g(x_1, x_2, x_3, x_4) = x_1/x_2 - x_3/x_4$ for $x_2 \neq 0$ and $x_4 \neq 0$. In view of (A15) and the fact that

$$\frac{\partial}{\partial x} g(x)\big|_{x=\mu} = \big(\frac{1}{\mathrm{CVaR}_1}, -\frac{MtC_1}{\mathrm{CVaR}_1}, -\frac{1}{\mathrm{CVaR}_0}, \frac{MtC_0}{\mathrm{CVaR}_0}\big)^\top \neq 0,$$

we get by an application of the delta method, see (Brockwell and Davis, 1991, Proposition 6.4.2), that

$$\sqrt{S}\big(T_S - (MtC_1^* - MtC_0^*)\big) \xrightarrow{d} \mathcal{N}(0, (\frac{\partial}{\partial x} g(x)\big|_{x=\mu})^\top \Sigma_r \frac{\partial}{\partial x} g(x)\big|_{x=\mu}). \tag{A16}$$

The assertion of the theorem follows since under the null hypothesis, $MtC_1^* = MtC_0^*$.

B Robustness Tests

We report results with several robustness tests relating to the model and our empirical findings.

B.1 Outliers and model robustness

First, we repeat the in-sample test excluding the great financial crisis of Jan. 2007–Dec. 2009 (Table B1). The benchmark portfolios perform better when the crisis is omitted, but the performance of the MtC portfolios also improves, and the benefits from international diversification persist. For the hedged portfolios, we observe, for all three countries, economically and statistically significant differences of both MtC and Sharpe ratios over both benchmarks. This test shows that outliers are not driving our results.

Second, we repeat the test by solving the model at different confidence levels α in the CVaR estimation (Table B2). The internationally diversified portfolios are exposed to economically large and statistically significant political betas for all tested α's. Likewise, the benefits from international diversification persist, and the hedged portfolios exhibit economically and statistically significant differences of both MtC and Sharpe ratios over both benchmarks. Our results are robust to the choice of α.

Third, we conducted a randomized test. We block bootstrap with replacement, the time series of data to generate 100 samples of market and benchmark portfolio returns. We run the MtC models on these bootstrapped time series, estimate the portfolio β_P, and the MtC and Sharpe ratios, and compute the proportion of significant political betas and the proportion of the hedged portfolios outperforming the benchmarks. The results (Table B3) strongly support those of Table 3. For the three countries, respectively, we obtain significant political beta for 90, 86, and 92 runs. The politically hedged portfolio's MtC and Sharpe ratios outperform the index in 97-100 runs and outperform the EW benchmark in 89-96 runs. Our findings are robust to alternative data specifications.

Finally, we use the out-of-sample test to assess the accuracy of our tail estimates. We check whether our simulations over the 204 repetitions accurately capture the tail of market returns. We estimate the VaR of portfolio returns at t using historical data and find that it brackets the ex-post observed $(t+1)$st return with the appropriate frequency: the ex-post portfolio return is in the right 0.20 quantile tail for 22% of our trials, in the 0.15 quantile with frequency 16.7%, in the 0.10 with frequency 13%, and in the 0.05 with frequency 9%. Overall, the empirically observed frequencies match the tail estimates.

B.2 Transaction costs

We introduce proportional transaction costs (Zenios, 2007) to assess whether diversification benefits may disappear when hedging political risk. The results are reported in Table B4. The politically hedged portfolio registers economically and statistically significant MtC gains over both benchmark portfolios, with transaction costs as high as 0.2% for developed and 0.5% for emerging markets. Higher trading costs imply smaller gains vis-à-vis the benchmarks. Comparing the hedged portfolios with and without transaction costs, we notice a reduction

Table B1 – Subsample analysis excluding the Great Financial Crisis

This table reports performance statistics of the MSCI home market index I and international portfolios for US, Eurozone, and Japan investors, using equally weighted portfolios EW and mean-to-CVaR unrestricted optimal portfolios (U) and with political risk hedging (H). Returns are in the home currency, and political risk is hedged with net zero exposure to the P-factor. Reported are the exposures to a global political risk factor β_P, and the monthly performance ratios mean-to-CVaR (MtC) and Sharpe ratio. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019 but excluding 2007 to 2009. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

	I	EW	U	H	U-I	H-I	U-EW	H-EW
				(a) US				
β_P	0.00	0.12*	0.30*	0.00	0.30*	0.00	0.18*	-0.12*
	(0.85)	(0.00)	(0.00)	(1.00)	(0.00)	(0.93)	(0.00)	(0.00)
Av. excess return	0.68	0.79	1.53	1.13	0.85	0.45	0.74	0.34
CVaR	8.75	9.87	10.37	8.65	1.62	-0.09	0.50	-1.21
MtC	0.078	0.080	0.148	0.130	0.070*	0.052*	0.068*	0.050*
					(0.04)	(0.05)	(0.00)	(0.00)
Sharpe	0.17	0.17	0.27	0.26	0.10*	0.09*	0.10*	0.09*
					(0.08)	(0.09)	(0.01)	(0.01)
				(b) Eurozone				
β_P	0.05	0.18*	0.26*	0.00	0.21*	-0.05	0.08	-0.18*
	(0.11)	(0.00)	(0.00)	(1.00)	(0.00)	(0.31)	(0.11)	(0.00)
Av. excess return	0.55	0.83	1.43	1.12	0.88	0.57	0.60	0.29
CVaR	10.32	8.79	7.76	7.95	-2.56	-2.36	-1.03	-0.83
MtC	0.053	0.095	0.184	0.140	0.131*	0.087*	0.089*	0.046*
					(0.00)	(0.00)	(0.00)	(0.03)
Sharpe	0.12	0.21	0.33	0.29	0.73*	0.60*	0.43*	0.30*
					(0.00)	(0.00)	(0.00)	(0.08)
				(c) Japan				
β_P	0.03	0.14*	0.34*	0.00	0.31*	-0.03	0.19*	-0.14*
	(0.52)	(0.00)	(0.00)	(1.00)	(0.00)	(0.65)	(0.00)	(0.01)
Av. excess return	0.77	1.06	1.76	1.37	0.99	0.60	0.70	0.31
CVaR	9.24	10.36	10.69	9.43	1.45	0.19	0.33	-0.92
MtC	0.084	0.102	0.165	0.145	0.081*	0.062*	0.063*	0.043*
					(0.03)	(0.07)	(0.00)	(0.02)
Sharpe	0.16	0.20	0.30	0.29	0.48*	0.43*	0.35*	0.29*
					(0.04)	(0.07)	(0.01)	(0.02)

of MtC gains over both I and EW, as expected. However, the gains of the politically hedged portfolios over the benchmarks remain statistically significant for a wide range of transaction costs.

We also noted (results not shown) that adding transaction costs tilts the portfolio towards politically risky countries due to their higher expected returns. The political premium of the unconstrained portfolio with transaction costs 0.5% as in De Roon et al. (2001), is 4.91% for the US, 4.35% for the Eurozone, and 4.45% for Japan. These values are larger than the premia

Table B2 – Solving the hedging model at different confidence levels

This table reports performance statistics of the international portfolios when solving the mean-to-CVaR model with different confidence levels with α set equal to 0.99, 0.97, 0.95, and 0.90. Reported are the exposures to a global political risk factor β_P and differences in the monthly performance ratios between the internationally diversified portfolios and the benchmarks. MtC_U, MtC_H, MtC_I, MtC_{EW} are the MtC ratios for the unrestricted, hedged, and benchmark portfolios, respectively, and SR_U, SR_H, SR_I, SR_{EW} are the corresponding Sharpe ratios. Returns are monthly in the home currency, and political risk is hedged with net zero exposure to the P-factor. The sample includes 22 developed economies and 20 emerging markets, spanning 1999–2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

α	0.99	0.97	0.95	0.90	0.99	0.97	0.95	0.90	0.99	0.97	0.95	0.90
	(a) US				(b) Eurozone				(c) Japan			
β_P	0.62*	0.39*	0.34*	0.41*	0.19*	0.43*	0.31*	0.25*	0.69*	0.42*	0.42*	0.41*
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$MtC_U - MtC_I$	0.028	0.044*	0.053*	0.072*	0.065*	0.080*	0.089*	0.114*	0.038*	0.053*	0.065*	0.085*
	(0.11)	(0.08)	(0.06)	(0.07)	(0.00)	(0.01)	(0.00)	(0.00)	(0.06)	(0.05)	(0.04)	(0.03)
$SR_U - SR_I$	0.07	0.09	0.09*	0.09	0.15*	0.15*	0.16*	0.17*	0.11*	0.13*	0.12*	0.13*
	(0.24)	(0.11)	(0.10)	(0.13)	(0.00)	(0.02)	(0.00)	(0.00)	(0.09)	(0.03)	(0.03)	(0.03)
$MtC_H - MtC_I$	0.017	0.029	0.038*	0.053*	0.051*	0.052*	0.060*	0.080*	0.023	0.035	0.045*	0.068*
	(0.17)	(0.13)	(0.10)	(0.08)	(0.00)	(0.00)	(0.00)	(0.00)	(0.14)	(0.11)	(0.09)	(0.06)
$SR_H - SR_I$	0.05	0.07	0.07	0.07	0.11*	0.14*	0.14*	0.15*	0.09*	0.12*	0.11*	0.11*
	(0.32)	(0.15)	(0.15)	(0.13)	(0.00)	(0.00)	(0.00)	(0.00)	(0.10)	(0.04)	(0.05)	(0.05)
$MtC_U - MtC_{EW}$	0.031*	0.042*	0.047*	0.061*	0.042*	0.051*	0.055*	0.072*	0.03*	0.039*	0.044*	0.053*
	(0.03)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.00)	(0.00)	(0.03)	(0.01)	(0.01)	(0.01)
$SR_U - SR_{EW}$	0.06	0.08*	0.08*	0.08*	0.07*	0.07	0.09*	0.09*	0.05	0.07*	0.07*	0.07*
	(0.24)	(0.04)	(0.03)	(0.04)	(0.04)	(0.12)	(0.02)	(0.00)	(0.29)	(0.05)	(0.05)	(0.04)
$MtC_H - MtC_{EW}$	0.02*	0.027*	0.032*	0.042*	0.028*	0.023*	0.026*	0.038*	0.015*	0.021*	0.024*	0.036*
	(0.04)	(0.02)	(0.02)	(0.01)	(0.06)	(0.03)	(0.04)	(0.02)	(0.06)	(0.02)	(0.03)	(0.02)
$SR_H - SR_{EW}$	0.04	0.06*	0.06*	0.06*	0.03	0.06*	0.07*	0.07*	0.03	0.06*	0.06*	0.05*
	(0.28)	(0.05)	(0.04)	(0.02)	(0.33)	(0.03)	(0.07)	(0.04)	(0.26)	(0.02)	(0.03)	(0.04)

without transaction costs (respectively 2.69%, 2.46%, and 3.33%) and much larger than the political premia of the home indices, reaffirming the importance of political risk.

B.3 Short positions

We finally consider short positions in developed markets (De Roon, Nijman, and Werker, 2001) and report the results in Table B5.

We observe that the unrestricted MtC portfolio political beta is not statistically significant for the US and Japanese investors. For Eurozone investors, it is half the beta of the NSS unrestricted portfolios from Table 3. The optimal portfolios have long-short positions that hedge political risk. This is an outcome of optimizing tail risk and is not achieved with Sharpe ratio maximization, as shown in section 4.2. The asset allocations of the unconstrained portfolio for the US go long in developed countries with negative political beta (Denmark, Switzerland) and short in positive betas (Finland, Greece). The resulting portfolio beta is not statistically significant. In contrast, the SR portfolio from Table 2 has a statistically significant political beta of 0.12. The MtC portfolio of the Japanese investor has a non-significant beta, whereas the SR portfolio has a beta of 0.15. Examining the optimal portfolio weights, we observe diversified,

Table B3 – Randomized test of the hedging model

This table reports statistics for solving 100 instances of the MtC model with returns generated using block bootstrapping. Reported are the percentage of portfolios with significant political risk exposure β_P with its average and the percentage of runs when the hedged portfolios outperform the benchmarks. MtC_{H}, MtC_{I}, MtC_{EW} are the MtC ratios for the hedged and benchmark portfolios, respectively, and SR_{H}, SR_{I}, SR_{EW} are the corresponding Sharpe ratios. The sample includes 22 developed economies and 20 emerging markets, spanning 1999–2019.

Proportion non-zero		Proportion positive	
		US	
β_P	90%	$\text{MtC}_{\text{H}} - \text{MtC}_{\text{I}}$	99%
(Average	0.26)	$\text{SR}_{\text{H}} - \text{SR}_{\text{I}}$	97%
		$\text{MtC}_{\text{H}} - \text{MtC}_{\text{EW}}$	95%
		$\text{SR}_{\text{H}} - \text{SR}_{\text{EW}}$	95%
		Eurozone	
β_P	86%	$\text{MtC}_{\text{H}} - \text{MtC}_{\text{I}}$	100%
(Average	0.23)	$\text{SR}_{\text{H}} - \text{SR}_{\text{I}}$	100%
		$\text{MtC}_{\text{H}} - \text{MtC}_{\text{EW}}$	96%
		$\text{SR}_{\text{H}} - \text{SR}_{\text{EW}}$	94%
		Japan	
β_P	92%	$\text{MtC}_{\text{H}} - \text{MtC}_{\text{I}}$	100%
(Average	0.28)	$\text{SR}_{\text{H}} - \text{SR}_{\text{I}}$	100%
		$\text{MtC}_{\text{H}} - \text{MtC}_{\text{EW}}$	89%
		$\text{SR}_{\text{H}} - \text{SR}_{\text{EW}}$	92%

balanced portfolios in both the unrestricted and hedged cases, with about nine assets in the long and six in the short legs and holdings in both developed and emerging markets.

We observe performance gains again when diversifying internationally without any restrictions on political risk. MtC gains over I are in the range of 0.135–0.221 for the three investors, with Sharpe increases in the range of 0.18–0.27, and all are strongly statistically significant. Significant MtC gains are also achieved over EW. The gains persist when hedging political risk, with MtC ratios increasing in the range of 0.132–0.207 over I and 0.122–0.174 over EW.

In conclusion, political risk can be diversified away if short sales are allowed in the case of the US and Japan and are significantly reduced from the no-short-sales case for the Eurozone. Consistently with our main result, the benefits from international diversification persist with politically neutral portfolios.

Table B4 – Hedging political risk with transaction costs

This table reports performance statistics on portfolio performance with varying levels of transaction costs in developed (c_d) and emerging (c_e) markets. Results are reported for the MSCI home market index I and international portfolios using equally weighted EW and mean-to-CVaR unrestricted optimal portfolios (U) and with political risk hedging (H). Reported are the exposures to a global political risk factor β_P, and the monthly performance ratios mean-to-CVaR (MtC) and Sharpe ratio. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

Transaction costs			I	EW	U	H	U-I	H-I	U-EW	H-EW
					(a) US					
$c_d = 0.0\%$	$c_e = 0.0\%$	MtC	0.053	0.059	0.106	0.091	0.053*	0.038*	0.047*	0.032*
							(0.06)	(0.10)	(0.01)	(0.01)
		Sharpe	0.12	0.13	0.21	0.19	0.09*	0.07	0.08*	0.06*
							(0.10)	(0.15)	(0.03)	(0.04)
$c_d = 0.2\%$	$c_e = 0.2\%$	MtC	0.032	0.042	0.089	0.072	0.057*	0.040*	0.048*	0.030*
							(0.05)	(0.08)	(0.02)	(0.02)
		Sharpe	0.07	0.10	0.18	0.16	0.10*	0.08	0.08*	0.06*
							(0.08)	(0.11)	(0.04)	(0.04)
$c_d = 0.2\%$	$c_e = 0.4\%$	MtC	0.032	0.034	0.079	0.065	0.047*	0.033*	0.045*	0.031*
							(0.07)	(0.10)	(0.01)	(0.00)
		Sharpe	0.07	0.08	0.16	0.15	0.08*	0.07	0.08*	0.07*
							(0.10)	(0.11)	(0.02)	(0.01)
$c_d = 0.2\%$	$c_e = 0.5\%$	MtC	0.032	0.030	0.074	0.062	0.042*	0.030*	0.044*	0.033*
							(0.08)	(0.10)	(0.01)	(0.00)
		Sharpe	0.07	0.07	0.15	0.14	0.08	0.07	0.08*	0.07*
							(0.14)	(0.11)	(0.02)	(0.00)
					(b) Eurozone					
$c_d = 0.0\%$	$c_e = 0.0\%$	MtC	0.030	0.063	0.119	0.090	0.089*	0.060*	0.056*	0.026*
							(0.00)	(0.00)	(0.00)	(0.04)
		Sharpe	0.07	0.15	0.24	0.22	0.16*	0.14*	0.09*	0.07*
							(0.00)	(0.00)	(0.01)	(0.06)
$c_d = 0.2\%$	$c_e = 0.2\%$	MtC	0.013	0.044	0.101	0.070	0.088*	0.057*	0.057*	0.026*
							(0.00)	(0.00)	(0.01)	(0.04)
		Sharpe	0.03	0.11	0.19	0.17	0.16*	0.14*	0.08*	0.07*
							(0.01)	(0.00)	(0.04)	(0.07)
$c_d = 0.2\%$	$c_e = 0.4\%$	MtC	0.013	0.035	0.087	0.064	0.074*	0.051*	0.052*	0.029*
							(0.01)	(0.00)	(0.01)	(0.02)
		Sharpe	0.03	0.09	0.17	0.16	0.14*	0.13*	0.08*	0.08*
							(0.02)	(0.00)	(0.04)	(0.03)
$c_d = 0.2\%$	$c_e = 0.5\%$	MtC	0.013	0.031	0.081	0.063	0.068*	0.050*	0.051*	0.033*
							(0.01)	(0.00)	(0.01)	(0.02)
		Sharpe	0.03	0.08	0.16	0.16	0.13*	0.12*	0.09*	0.08*
							(0.02)	(0.00)	(0.03)	(0.02)
					(c) Japan					
$c_d = 0.0\%$	$c_e = 0.0\%$	MtC	0.044	0.065	0.109	0.090	0.065*	0.045*	0.044*	0.024*
							(0.03)	(0.09)	(0.01)	(0.03)
		Sharpe	0.09	0.15	0.22	0.20	0.12*	0.11*	0.07*	0.06*
							(0.03)	(0.05)	(0.05)	(0.03)
$c_d = 0.2\%$	$c_e = 0.2\%$	MtC	0.025	0.050	0.094	0.073	0.070*	0.048*	0.045*	0.023*
							(0.02)	(0.07)	(0.01)	(0.03)
		Sharpe	0.05	0.11	0.19	0.17	0.14*	0.12*	0.08*	0.06*
							(0.02)	(0.04)	(0.03)	(0.05)
$c_d = 0.2\%$	$c_e = 0.4\%$	MtC	0.025	0.042	0.082	0.068	0.058*	0.043*	0.040*	0.026*
							(0.05)	(0.08)	(0.01)	(0.01)
		Sharpe	0.05	0.10	0.17	0.16	0.12*	0.11*	0.07*	0.06*
							(0.04)	(0.04)	(0.05)	(0.00)
$c_d = 0.2\%$	$c_e = 0.5\%$	MtC	0.025	0.039	0.077	0.066	0.053*	0.042*	0.039*	0.028*
							(0.06)	(0.08)	(0.02)	(0.01)
		Sharpe	0.05	0.09	0.16	0.16	0.11*	0.11*	0.07*	0.07*
							(0.06)	(0.04)	(0.05)	(0.00)

Table B5 – Diversifying political risk with short positions

This table reports performance statistics of international portfolios when we allow for short sales in developed but not emerging markets. Statistics are reported for the MSCI home market index (I) and international portfolios using equally weighted (EW) and mean-to-CVaR unrestricted optimal portfolios (U) and with political risk hedging (H). Reported are the exposures to a global political risk factor β_P, and the monthly performance ratios mean-to-CVaR (MtC) and Sharpe ratio. The sample includes 22 developed economies and 20 emerging markets, spanning 1999 to 2019. We use * to denote rejection of the null at the 10% level or less, as shown by the p-values in parentheses.

	I	EW	U	H	U-I	H-I	U-EW	H-EW
(a) US								
β_P	-0.01	0.12*	0.09	0.00	0.10*	0.01	-0.03	-0.12*
	(0.57)	(0.00)	(0.13)	(1.00)	(0.05)	(0.83)	(0.52)	(0.02)
Av. excess return	0.52	0.71	1.28	1.23	0.76	0.71	0.57	0.52
CVaR	9.90	12.16	6.82	6.68	-3.08	-3.22	-5.34	-5.48
MtC	0.053	0.059	0.188	0.185	0.135*	0.132*	0.129*	0.126*
					(0.00)	(0.00)	(0.00)	(0.00)
Sharpe	0.12	0.13	0.30	0.30	0.18*	0.18*	0.17*	0.17*
					(0.01)	(0.01)	(0.01)	(0.01)
(b) Eurozone								
β_P	0.04	0.16*	0.15*	0.00	0.11*	-0.04	-0.01	-0.16*
	(0.22)	(0.00)	(0.02)	(1.00)	(0.09)	(0.59)	(0.81)	(0.02)
Av. excess return	0.36	0.68	1.55	1.77	1.19	1.40	0.87	1.08
CVaR	12.16	10.80	6.19	7.45	-5.97	-4.71	-4.60	-3.35
MtC	0.030	0.063	0.251	0.237	0.221*	0.207*	0.187*	0.174*
					(0.00)	(0.00)	(0.00)	(0.00)
Sharpe	0.07	0.15	0.34	0.34	0.27*	0.27*	0.19*	0.19*
					(0.00)	(0.00)	(0.00)	(0.01)
(c) Japan								
β_P	0.04	0.14*	0.08	0.00	0.04	-0.04	-0.06	-0.14*
	(0.41)	(0.00)	(0.29)	(1.00)	(0.59)	(0.61)	(0.44)	(0.06)
Av. excess return	0.47	0.88	1.69	1.69	1.22	1.22	0.81	0.81
CVaR	10.54	13.50	8.82	9.04	-1.72	-1.50	-4.69	-4.46
MtC	0.044	0.065	0.192	0.187	0.147*	0.143*	0.126*	0.122*
					(0.00)	(0.00)	(0.00)	(0.00)
Sharpe	0.09	0.15	0.30	0.30	0.21*	0.21*	0.15*	0.16*
					(0.00)	(0.00)	(0.01)	(0.01)