Chapter 3

Capital Share Risk in International Asset Pricing

This chapter introduces the second dissertation study which analyzes whether capital share growth is a priced risk factor across international equity markets.

Zusatzmaterial online

Zusätzliche Informationen sind in der Online-Version dieses Kapitel (https://doi.org/10.1007/978-3-658-35479-4_3) enthalten.

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3.1 Introduction

In a seminal study, [Lettau et al. \(2019\)](#) demonstrate that a single macroeconomic factor can explain a wide range of equity and nonequity portfolio returns within the U.S. market. This factor, which is based on the growth in the capital share of aggregate income, is able to outperform, yet even subsume information in well-established factor models as for instance the Fama-French three factor model. The aim of this paper is to study whether the explanatory power of this factor maintains across international equity markets.

It is well acknowledged that international markets are important ([Jacobs and Müller, 2020](#); [Karolyi, 2016](#)). By the end of 2017, U.S. firms accounted for approximately 35% of world-wide market capitalization, implying that 65% of the entire global equity market worth is attributable to international markets.¹ Beyond their importance in terms of absolute economic figures, international markets are explicitly relevant from a financial research perspective for at least two reasons.

First, they offer scientists the possibility to conduct broad international replications and out-of-sample tests. In doing so, they are an important tool for evaluating the relative importance and robustness of newly discovered asset pricing factors and anomalies. This seems particularly necessary given that throughout the past decades, a vast amount of new asset pricing factors and anomalies has inundated top finance journals. A trend that is increasingly associated with negative terms such as *p-hacking* ([Harvey, 2017](#)), *data mining* ([Harvey et al., 2016](#)), and *zoo of new factors* ([Cochrane, 2011](#)). In the anomaly literature, [Hou et al. \(2020a\)](#) for instance illustrate that the economic magnitude of the majority of 452 replicated anomalies is significantly smaller than reported in the corresponding reference studies. A result strikingly emphasizing the need for broad replication studies and broad out-of-sample tests.

Second, analyses of international markets enhance our understanding of what segments or integrates international equity markets ([Bekaert et al., 2011](#); [Bekaert and Harvey, 2017](#)). Thus, they are inevitable for obtaining a holistic understanding of global capital markets

¹Calculations are based upon our data set described in Section 3.3.

and asset price formation processes. Detecting differences in factor and anomaly estimates for international markets compared to those obtained for the U.S. market is the first crucial step in identifying drivers of global equity market segmentations. Indeed, a wide range of asset pricing studies finds that many patterns obtained within the U.S. deviate across international markets (Fama and French, 2012, 2017; Jacobs and Müller, 2018, 2020).

Given these recent developments, the primary aim of our study thus is to employ an extensive global data set to analyze the relative importance and robustness of the newly discovered capital share growth factor. In particular, we strive to answer the following research questions:

1. Is capital share growth a persistent and consistent driver of equity portfolio returns across global equity markets?
2. Are there differences and commonalities across countries and regions worldwide? What explains potential cross-country differences?
3. Is it possible to exploit the information contained in the capital share growth factor by a tradable (out-of-sample) investment strategy?

To answer these questions, we implement an analysis of 45 developed and emerging equity markets across the globe. Specifically, to ensure comprehensive and thorough, yet simultaneously solid and robust analyses, we restrict our sample to countries listed in the following regional indices: MSCI North America, MSCI Europe, MSCI Pacific, and MSCI Emerging Markets.

Our paper reports the following key results. First, we are able to replicate the results reported by Lettau et al. (2019) for the U.S. market in both, size and significance, thus dispersing potential p-hacking and data mining concerns. Second, we illustrate that capital share growth is a priced risk factor across international markets. There exists, however, strong geographic heterogeneity in the explanatory power of this factor. Pooled estimates reveal that capital share growth is particularly significant for Emerging Markets, while being less relevant for G7 + Australia (excl. USA). Unlike the U.S. market, though, the information contained in the KS risk factor of international markets does not subsume information contained in

alternative factor models as for instance the Fama-French three factor or q-factor models, but rather adds additional explanatory content to these model specifications. Pooled estimates for G7 + Australia (excl. USA), however, reveal that the capital share factor is in part even entirely subsumed by alternative factor model specifications.

Third, capital share growth is a local rather than global risk factor. Once applying global capital share estimates, the explanatory power of the factor is substantially reduced and oftentimes disappears entirely. These results are in line with studies for instance conducted by [Hollstein \(2020\)](#) or [Hou et al. \(2011\)](#), implying that local factors exhibit greater explanatory power for international equity markets compared to global factor constructions.

Fourth, cross-sectional regression analysis reveals that cross-country differences in the explanatory power of local KS risk estimates can best be explained by variations in private wealth inequality as well as further macroeconomic variables as for instance variations in public wealth. These findings underpin the theoretical explanations given by [Lettau et al. \(2019\)](#) in an international setting: The smaller the amount of capital market risk beard by a given subset of investors, the smaller the expected compensation in terms of equity portfolio returns.

Lastly, we are able to illustrate that - while exhibiting great in-sample explanatory power and significance - the information contained in the capital share growth factor cannot be exploited out-of-sample. An investment strategy that goes long the decile of past capital share growth beta winners and short the decile of past capital share growth beta losers, yields insignificant results across international markets as well as even for the U.S. market only.

Overall, our findings are consistent with the fact that wealth inequality tends to increase whenever stock market returns are high ([Gomez, 2017](#)). Additionally, our line of reasoning supports empirical evidence recently reported by [Zhang \(2019\)](#) who shows that there exists a negative link between a firm's (option to replace) routine-based labor tasks and its exposure to systematic risk as well as its expected equity returns. Most importantly, though, our results confirm the significance of capital share growth as a source of priced risk within international equity markets.

The remainder of our study is organized as follows. Section 3.2 places our work within the current state of research and formally derives our hypotheses. Section 3.3 outlines the data set underlying our analysis and describes the methodological approach. Section 3.4 presents baseline findings and discusses the empirical evidence. Section 3.5 summarizes our findings, discusses implications, and concludes.

3.2 Related Literature and Hypothesis Development

Our study builds upon Lettau et al. (2019) who examine whether growth in the capital share of national income has explanatory power for equity characteristic portfolio returns as well as nonequity asset classes within the U.S. market. In line with Danthine and Donaldson (2002), Lettau et al. (2019) argue that wealthy people finance consumption out of asset ownership, whereas workers finance consumption out of salaries and wages. Aggregate shocks that shift capital income towards labor income are thus considered as a source of systematic risk for security holders. This tension between labor and capital has long been recognized within academia as for instance within studies by Lustig and Van Nieuwerburgh (2008) or Julliard (2007). Given this theoretical backdrop, our work adds to a broader literature stream analyzing potential interdependencies between unequal asset ownership and stock market fluctuations (Heaton and Lucas, 2000; Gomez, 2017; Guvenen, 2009; Malloy et al., 2009; Mankiw and Zeldes, 1991). More specifically, our study relates to prior academic works examining the role of redistributive shocks as a potential source of systematic risk (Danthine and Donaldson, 2002; Greenwald et al., 2014; Lettau et al., 2019).

In this regard, Heaton and Lucas (2000) present empirical evidence that income derived out of entrepreneurial ventures constitutes a source of undiversifiable risk, which in turn tends to be more highly correlated with common equity returns. Likewise, Danthine and Donaldson (2002) argue that equity returns are solely determined by the consumption and preferences of firm owners. Guvenen (2009) argues that the interaction of both, limited participation in the stock market as well as heterogeneity in the elasticity of intertemporal substitution in consumption, is relevant for macroeconomic asset pricing. Gomez (2017) finds that wealth inequality increases whenever stock market returns are high, while simultaneously arguing

that higher wealth inequality predicts lower future stock returns. [Greenwald et al. \(2014\)](#) find that long-term stock market fluctuations are driven by shocks that redistribute production rewards between firm owners and workers.

Besides this macroeconomic asset pricing literature stream, our paper relates to a growing research field analyzing the validity and persistence of various asset pricing factors and anomalies observed within the U.S. for international markets. In this regard, prior academic work has uncovered that many patterns obtained within the U.S. deviate for international markets ([Fama and French, 2012](#); [Hou et al., 2011](#); [Jacobs and Müller, 2018, 2020](#)). [Fama and French \(2012\)](#) for instance find that Japan is an exception when it comes to the existence of the size or momentum premium. Studying 240 cross-sectional predictors, [Jacobs and Müller \(2018\)](#) show that there exists a large geographic heterogeneity in the significance of asset pricing factors. Relatedly, [Jacobs and Müller \(2020\)](#) show that the post-publication decline of asset pricing anomalies detected by [McLean and Pontif \(2016\)](#) for the U.S. market is not observable across a wide range of international markets. Given that prior studies to a great extent find that empirical asset pricing factors and anomalies observed within the U.S. market are not (fully) transferable to international markets, we thus hypothesize for our study as follows:

Hypothesis 1: There exists (strong) geographic heterogeneity in the explanatory power of local capital share growth for expected equity characteristic portfolio returns.

Similarly, there exists an ongoing academic debate on whether global, regional, or rather local factors are more suitable in explaining international asset prices ([Griffin, 2002](#); [Hau, 2011](#); [Hou et al., 2011](#); [Hollstein, 2020](#); [Karolyi and Wu, 2014](#); [Lewis, 2011](#)). Throughout the past decades, national boundaries have become increasingly blurred and financial markets have become increasingly global. Thus, it is a logical endeavor to ask whether international equity returns are (increasingly) driven by global rather than local risk factors ([Petzev et al., 2016](#)). Indeed, studies conducted by [Brooks and Del Negro \(2005\)](#) as well as [Hau \(2011\)](#) underline the importance of regional and global rather than local risk factors. [Karolyi and Wu \(2014\)](#), however, emphasize that a partial-segmentation approach, that is a combination of global and local risk factors, is superior to purely-global and purely-local factor models.

For our study, this implies that a potential absence of explanatory power of local capital share growth for international markets might be attributable to the fact that a global capital share growth factor might be more useful in explaining international portfolio returns. Conversely, however, the bulk of most recent empirical evidence suggests that local factors are more useful in explaining international equity returns (Fama and French, 2012; Hollstein, 2020; Hou et al., 2011). For this reason, we formulate our second hypothesis as follows:

Hypothesis 2: Capital share growth is a local rather than global risk factor.

Provided that we indeed observe geographic heterogeneity in the explanatory power of capital share growth, we strive to understand potential causes of cross-country differences. Within their study, Lettau et al. (2019) argue that positive KS risk estimates are driven by imperfect risk-sharing due to unequal equity ownership. Aggregate shocks that shift capital income towards labor income are thus considered as a source of systematic risk for security holders. In this regard, it seems intuitive to assume that good news on the future growth of labor income constitutes adverse information to security holders. A security's price is determined by its future payouts (dividends). The higher a company's expenses, the smaller the amount left for potential payouts to its security holders. This is especially plausible since it is a stylized fact that labor costs represent one of the biggest cost drivers within the firm-universe. A logical conjecture thus might be that cross-country differences in unequal asset ownership lead to variations in empirical results of reported KS risk estimates. This conjecture relates to the broader literature stream arguing that limited stock market participation has an impact upon asset prices (Basak and Cuoco, 1998; Favilukis, 2013; Gomez, 2017; Guvenen, 2009; Guo, 2004). Following this line of argumentation, we further hypothesize as follows:

Hypothesis 3: Cross-country differences in the explanatory power of capital share growth are driven by variations in local wealth inequality.

However, we argue that there exist further drivers of potential cross-country differences in local capital share risk estimates. In this regard, an evident driver of cross-country variations lies in potential misreportings of a country's labor and capital share values. These figures might be distorted either via a country's aggregate income derived out of wages and salaries or a country's income obtained from the ownership of firms and financial investments. In

this regard, [Burq and Chancel \(2020\)](#) for instance emphasize that despite improvements throughout the past decades, there still exist opaque financial systems which lack available information on income and wealth distributions across the globe. Relatedly, we argue that one of the biggest factors distorting national labor share values lies in the local shadow economy labor force. Following this line of reasoning, differences in KS risk estimates are driven by differences in national shadow economies. According to [Schneider et al. \(2010\)](#), the weighted average size of the shadow economy between 1999 and 2007 amounts to 36.4% within Europe and Central Asia, whereas it equals only 13.4% for high income OECD countries. These differences, in turn, impact the accuracy of estimated labor and capital shares, thus potentially explaining cross-country differences in reported KS risk estimates.

Also, and as for instance argued by [Chen et al. \(2011\)](#), workers' wages and salaries are particularly influenced by labor unions which simultaneously impact a firm's cost of equity. This rationale is in line with further studies in the field of labor economics as for instance the ones conducted by [Abowd \(1989\)](#) and [Hirsch \(1991\)](#). Apart from a potential increase in labor shares, these studies have uncovered that powerful labor unions reduce firms' operating flexibility, profitability, and in doing so they inevitably impact fundamental firm values. Beyond, labor unions impose a crucial burden to employment contracts in the form of wage inflexibility and workforce layoffs. Therefore, we argue that collective bargainings and trade union density increases the risk of shocks which redistribute the income between shareholders and employees. This implies that we expect capital share risk estimates to be higher within countries which exhibit higher numbers of trade unions as well as collective bargainings.

Lastly, assuming that potential shocks which redistribute the income between workers and security holders are at least partially offset by local governments, we presume a limiting effect upon KS risk estimates. As for instance argued by [Alvaredo et al. \(2018\)](#), a decline in public wealth limits a government's ability to redistribute income or to offset rising inequality. Also, and as for instance reported by [Dominguez et al. \(2012\)](#), countries which exhibit higher levels of total reserves are hypothesized to simultaneously exhibit stronger economic recovery after potential crises, thus partially offsetting adverse effects for security holders. We therefore hypothesize a negative link between a country's level of public wealth as well as a country's

total reserves and reported KS risk estimates.

Overall, we lastly hypothesize as follows:

Hypothesis 4: Cross-country differences in the explanatory power of capital share growth can further be explained by differences in economic transparency, differences in the bargaining power of local workforce as well as variations in public wealth or public reserves.

3.3 Data and Methodology

3.3.1 Data Sample and Summary Statistics

To construct our final sample, we combine data from multiple sources and screen as follows. First, we derive stock market data from Datastream/Worldscope. Monthly stock returns and further stock-related measures are gathered from Datastream. Accounting data is derived from Worldscope. Equity returns are measured in U.S. dollar and winsorized at the 0.1% and 99.9% levels to account for the effect of outliers.

Second, and in order to calculate a country's capital share, we gather data on national labor shares (*labsh*) from the PENN World Table (PWT). The PWT is compiled by the University of Groningen and has served as an established source of data on real GDP across countries for over four decades. In this regard, to make local GDPs at national prices comparable across countries, international GDPs are converted to U.S. dollars (Feenstra et al., 2015). *Labsh* indicates a country's share of labor compensation in GDP at current national prices. We then calculate a country's capital share (KS) as follows: $KS = 1 - Labsh$. While Lettau et al. (2019) build their analysis on quarterly capital share growth measures, we apply annual capital share growth values due to international data availability issues. Applying annual KS data does not alter the spirit and interpretation of empirical results reported. Once applying annual KS data for the same period investigated by Lettau et al. (2019) within the U.S., we obtain values for R^2 and coefficients within the 95% confidence interval reported by Lettau et al. (2019).

We restrict our sample to countries listed in the following regional indices: MSCI North

America, MSCI Europe, MSCI Pacific, and MSCI Emerging Markets. However, if there is no data on labor share available for a certain nation, we exclude the corresponding country from our analyses. These screening criteria lead to a final sample of 45 countries within our study, including the U.S. market. Our sample period runs from January 1989 to December 2017. The start date is a tradeoff between maximizing the length of the time-series and ensuring a maximum number of countries within the study. Our end date is set by availability of data on national labor shares within the PENN World Table. Please note that while our overall sample period runs from January 1989 to December 2017, data coverage in Datastream starts later for some emerging markets. Table 3.1 exemplifies corresponding start and end dates as well as remaining summary statistics on a country-basis.

At the country-level, the average number of firms per month ranges from highest 4,444.93 (United States) to lowest 32.11 (Hungary). Average KS is highest in Qatar (0.76), Malaysia (0.67), and Egypt (0.62). Lowest average KS values are reported for Taiwan (0.29), Canada (0.33), and Switzerland (0.35). Average KS growth over one-year horizons is highest for India (1.66%), Ireland (1.52%), and Taiwan (1.12%) and lowest for Argentina (-0.61%), Brazil (-0.46%), and Malaysia (-0.39%). Average KS growth over two-year horizons is also highest for India (3.40%), Ireland (3.05%), and Taiwan (2.26%). Lowest average KS growth values over two-year horizons accordingly are observed within Argentina (-1.16%), Brazil (-0.91%), and Malaysia (-0.78%).

For reasons of clarity and simplicity, we cluster our chosen set of countries for the main analyses described below as follows: U.S. market, G7 + Australia (excl. U.S.)², and Emerging Markets³. Additionally, we run analyses for a World Sample, containing all applied countries with the exception of the U.S. market.

²Australia, Canada, Germany, France, Italy, Japan, and United Kingdom.

³Argentina, Brazil, Chile, China, Columbia, Czech Republic, Egypt, Greece, Hungary, India, Indonesia, Malaysia, Mexico, Peru, Philippines, Poland, Qatar, Russia, South Africa, South Korea, Taiwan, Thailand, and Turkey.

Table 3.1: Summary Statistics: Data Sample and Capital Share Values

This table exemplifies summary statistics for all countries included within the analyses. Our sample period runs from M1:1989 to M12:2017, however, data coverage in Datastream starts later for some emerging markets. Columns 3 and 4 show start and end dates for each equity market correspondingly. Column 5 reports the aggregate number of firms throughout our sample period, column 6 the average number of firms per month. Column 7 states a country's worldwide percental market value as of December 2017. Column 8 shows a country's average annual KS (i.e. the percentage of national GDP at current national prices that is non-attributable to the national share of labor compensation). Columns 9 and 10 illustrate time-series averages of growth in local KS over horizons of one (H4) and two (H8) years, respectively.

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Abbrev	Country	Start	End	Total # Firms	Avg # Firms/ Month	% of MV	Avg KS	Avg KS Growth (H4)	Avg KS Growth (H8)
arg	Argentina	1/1989	12/2017	112	53.07	0.10%	0.5754	-0.61%	-1.16%
atl	Australia	1/1989	12/2017	3043	1005.39	1.94%	0.4012	0.22%	0.47%
aus	Austria	1/1989	12/2017	182	79.01	0.20%	0.4040	0.57%	1.13%
bel	Belgium	1/1989	12/2017	254	114.26	0.56%	0.3754	0.12%	0.40%
bra	Brazil	6/1994	12/2017	285	123.83	0.83%	0.4559	-0.46%	-0.91%
can	Canada	1/1989	12/2017	4743	1444.01	2.82%	0.3324	0.42%	0.68%
chi	Chile	7/1989	12/2017	253	130.27	0.35%	0.5394	0.29%	0.62%
chn	China	1/1991	12/2017	3711	1312.62	11.76%	0.4230	0.10%	0.22%
col	Columbia	1/1992	12/2017	81	38.33	0.15%	0.5204	-0.09%	-0.19%
cze	Czech Republic	6/1996	12/2017	92	29.98	0.04%	0.4870	-0.12%	-0.17%
den	Denmark	1/1989	12/2017	338	163.77	0.54%	0.3648	0.33%	0.64%
egy	Egypt	12/1997	12/2017	221	112.17	0.06%	0.6230	0.21%	0.40%
fin	Finland	1/1989	12/2017	248	107.27	0.40%	0.3987	0.81%	1.53%
fra	France	1/1989	12/2017	1737	697.22	3.49%	0.3750	-0.03%	0.12%
ger	Germany	1/1989	12/2017	1534	560.60	2.94%	0.3646	0.55%	1.15%
gre	Greece	1/1989	12/2017	396	191.45	0.07%	0.4903	-0.06%	-0.10%
hkg	Hong Kong	1/1989	12/2017	2033	779.22	4.58%	0.5027	-0.32%	-0.59%
hun	Hungary	6/1993	12/2017	75	32.11	0.04%	0.3816	0.70%	1.34%
ind	India	1/1990	12/2017	3546	1203.22	3.16%	0.4312	1.66%	3.40%
ido	Indonesia	1/1989	12/2017	641	268.68	0.68%	0.5488	-0.12%	-0.23%
ire	Ireland	1/1989	12/2017	133	54.11	0.17%	0.5157	1.52%	3.05%
ita	Italy	1/1989	12/2017	597	236.85	0.97%	0.4729	0.32%	0.69%
jap	Japan	1/1989	12/2017	5402	3045.57	8.48%	0.4265	0.17%	0.41%
mal	Malaysia	1/1989	12/2017	1311	626.35	0.61%	0.6727	-0.39%	-0.78%
mex	Mexico	1/1989	12/2017	225	93.79	0.51%	0.6102	0.37%	0.76%
net	Netherlands	1/1989	12/2017	321	142.07	0.85%	0.3918	1.05%	2.23%
nzl	New Zealand	1/1989	12/2017	254	85.97	0.13%	0.4502	0.31%	0.74%
nor	Norway	1/1989	12/2017	536	162.54	0.41%	0.4877	0.55%	0.96%
per	Peru	1/1992	12/2017	190	82.99	0.11%	0.5615	-0.12%	-0.24%
phi	Philippines	6/1989	12/2017	309	163.31	0.38%	0.5992	0.48%	0.97%
pol	Poland	6/1992	12/2017	732	240.29	0.26%	0.3983	0.59%	1.18%
por	Portugal	1/1989	12/2017	138	57.73	0.09%	0.3703	0.57%	1.18%
qat	Qatar	6/2004	12/2017	49	37.41	0.17%	0.7613	0.74%	1.49%
rus	Russia	6/1997	12/2017	547	195.74	0.84%	0.4106	0.09%	0.13%
sin	Singapore	1/1989	12/2017	1065	431.04	0.85%	0.5569	-0.11%	-0.18%
soa	South Africa	1/1989	12/2017	850	268.30	0.80%	0.4424	0.30%	0.64%
sok	South Korea	1/1989	12/2017	2672	936.82	2.24%	0.4703	0.34%	0.61%
spa	Spain	1/1989	12/2017	351	141.11	1.09%	0.3786	0.58%	1.15%
swe	Sweden	1/1989	12/2017	994	296.72	0.86%	0.4539	0.11%	0.27%
swi	Switzerland	1/1989	12/2017	407	212.19	2.12%	0.3477	0.14%	0.29%
tai	Taiwan	1/1989	12/2017	2285	870.92	1.60%	0.2882	1.12%	2.26%
tha	Thailand	1/1989	12/2017	877	371.99	0.73%	0.6107	0.18%	0.39%
tur	Turkey	1/1989	12/2017	451	190.77	0.30%	0.5913	-0.16%	-0.33%
uni	United Kingdom	1/1989	12/2017	4462	1500.34	4.51%	0.4168	-0.29%	-0.55%
usa	USA	1/1989	12/2017	11684	4444.93	35.27%	0.3909	0.22%	0.43%

3.3.2 Econometric Approach

We follow the two-step regression approach proposed by [Lettau et al. \(2019\)](#). It is assumed that workers do not participate in the stock market, but entirely consume their wages and salaries. Thus, worker consumption is to be neglected in the following. As in equilibrium stockholders' share of consumption is assumed to equal the capital share, stockholders' consumption is to be neglected as well ([Lettau et al., 2019](#)).

First, we run univariate time-series regressions for each of the corresponding 5x5 equity portfolios on local KS growth, indicating the growth of local capital share over horizons of one (H4) and two (H8) years:

$$R_{j,t+H,t}^e = \alpha_j + \beta_{jKS,H} \left(\frac{KS_{t+H}}{KS_t} \right) + u_{j,t+H,t} \quad (3.1)$$

Within equation 3.1, we use the cumulative H -period excess return on asset j from the end of t to the end of $t+H$ ($R_{j,t+H,t}^e$). In this regard, a one period portfolio return equals the quarterly portfolio return on asset j from the end of t to the end of $t+1$.

To this end, we apply the following equity characteristic portfolio sorts: 25 portfolios sorted on Size/Book-to-Market (Size/BM), Size/Investment (Size/INV), Size/Operating Profitability (Size/OP), Size/Reversal (Size/REV), Size/Momentum (Size/Mom) as well as 10 portfolios sorted on Long-Term Reversal (REV). Calculation of equity portfolio returns follow the methodology described on Kenneth French's Dartmouth website. For each of our 5x5 equity portfolios, we calculate quarterly portfolio returns in the following way:

$$1 + R_{Q,Y} = \left(1 + \frac{R_{Q,Y}^{m1}}{100} \right) \left(1 + \frac{R_{Q,Y}^{m2}}{100} \right) \left(1 + \frac{R_{Q,Y}^{m3}}{100} \right) \quad (3.2)$$

Equation 3.2 implies that a portfolio's return in quarter Q of year Y ($R_{Q,Y}$) is the compounded monthly return over the three months within the respective quarter. We use the quarterly 3-month T-bill rate as a proxy for the risk-free rate.

Second, we analyze whether KS risk exposures estimated within equation 3.1 ($\beta_{jKS,H}$) exhibit

statistically significant explanatory power in the cross-section of average one-quarter equity portfolio returns:

$$\left(R_{j,t}^e\right) = \lambda_0 + \hat{\beta}_{j,KS,H} \lambda_{KS,H} + \epsilon_j \quad (3.3)$$

Where $R_{j,t}^e$ is the average quarterly 5x5 portfolio return and $\lambda_{KS,H}$ represents the KS risk price to be estimated within the cross-section.

The rationale beyond this two-step regression approach is to analyze whether longer-horizon (i.e. one-year and two-year) capital share risk exposures are able to explain average quarterly 5x5 equity portfolio returns. [Lettau et al. \(2019\)](#) argue that a more accurate estimate of KS exposure is obtained once applying longer-horizon capital share values (see [Lettau et al. \(2019\)](#) for detailed model specifications and justifications).

3.4 Empirical Results

3.4.1 Local Capital Share Growth in International Equity Markets

We begin the analysis by testing our first hypothesis, implying that there exists (strong) geographic heterogeneity in the explanatory power of local capital share growth for expected equity portfolio returns. In [Table 3.2](#), we summarize baseline findings obtained from applying local KS factors to local equity portfolio returns as described in [Section 3.3.2](#).

For comparison purposes, we show results for the U.S. market within Panel A. Panel B shows findings for G7 + Australia (excl. U.S.), Panel C for Emerging Markets, and Panel D for our World Sample (excl. U.S.).

Results for the U.S. market shown in Panel A are consistent with findings reported by Lettau et al. (2019). Across all equity portfolios, over one-year and two-year horizons, capital share growth is a highly significant priced risk factor. Even for the Size/Mom portfolios (not applied by Lettau et al. (2019)), we obtain estimates which are significant at the 1% level. Average $\lambda_{KS,H}$ values across all equity portfolio sorts within the U.S. amount to 0.6059 (t -statistics: 6.48) over one-year horizons and to 0.5563 (t -statistics: 7.42) over two-year horizons. The corresponding R^2 equals 0.64 and 0.71, respectively.

Within Panels B to D, we estimate first-step regression betas for each of our 5x5 equity portfolios using country-fixed effects. This means we calculate corresponding 5x5 equity portfolio returns as well as capital share growth values at the country level. We then append country-level portfolio returns for each of the corresponding 5x5 portfolio sorts. For instance, we append all returns obtained from lowest Size/BM portfolios (P1 of P25) for corresponding countries of each panel, accordingly. We check for robustness by applying country fixed-effects within the second, instead of first regression step following the methodology proposed by Lettau et al. (2019). Reported patterns are left unaffected, while coefficients and R^2 appear to be smaller. Corresponding standard errors are clustered at the country-level.

Within Panel B, i.e. for G7 + Australia (excl. U.S.), results indicate that capital share growth is again a significant factor in explaining equity portfolio returns. Yet, reported values are not as clear as for the U.S. market and call for a more differentiated perspective. Whereas risk estimates for Size/BM, REV, Size/INV, and Size/OP portfolios provide empirical support for capital share growth to be a priced risk factor, we obtain negative risk estimates for the Size/REV and Size/Mom portfolios (t -statistics smaller than two). Across all equity portfolios, we obtain average $\lambda_{KS,H}$ values of 1.3837 (t -statistics: 3.89) over one-year investment horizons, whereas we observe a decline to 0.7881 (t -statistics: 1.45) once applying two-year periods.

For Emerging Markets (Panel C), results imply a high explanatory power of local capital share growth for local equity portfolio returns. The only exception remains the Size/INV sort for which we observe insignificant estimates. Over one-year horizons, the average $\lambda_{KS,H}$ equals 1.2174 (t -statistics: 4.57) with an R^2 of 0.49. Unlike Panel B, the explanatory power

of local capital share growth maintains for Emerging Markets over two-year investment horizons. Here, we observe an average $\lambda_{KS,H}$ of 1.0249 (t -statistics: 5.50) with a corresponding R^2 of 0.55.

Results reported for our World Sample (excluding the U.S. market) within Panel D once more corroborate the strong explanatory power of local capital share growth for international markets at an aggregate level. The pooled international results are comparable to the ones obtained for the U.S. market in statistical significance. In size, they even exceed estimates for the U.S. sample. With the exception of the Size/INV portfolio sorts, all findings are statistically significant at the 1% level. Mean $\lambda_{KS,H}$ values across all portfolio sorts amount to 2.1205 (t -statistics: 5.28) with a corresponding R^2 of 0.53 for one-year investment horizons. Considering two-year investment horizons, we observe average $\lambda_{KS,H}$ values across all portfolio sorts of 1.7822 (t -statistics: 5.20) with a corresponding R^2 of 0.52.

We interpret these results as a first affirmation of our first hypothesis, implying that capital share growth is a priced risk factor which exhibits geographic heterogeneity across international equity markets. At an aggregate level, this finding strongly supports the theory proposed by Lettau et al. (2019) that positive exposure to capital share risk yields a positive risk premium. Simultaneously, however, the empirical evidence provided in Table 3.2 implies that the capital share risk factor is unable to explain the full range of equity portfolio returns tested.

3.4.2 Local Capital Share Growth in Multi-Factor Models

We proceed by further testing our first hypothesis, implying that there exists (strong) geographic heterogeneity in the explanatory power of local capital share growth. Within their study, Lettau et al. (2019) find that $\lambda_{KS,H}$ is a strong risk factor which is able to outperform the prominent Fama-French three factor model in explaining expected returns of portfolios sorted on Size/BM within the U.S. market. Within this section, we test whether this central finding deviates for international markets. This implies that we strive to understand whether local capital share growth is an independent and robust risk factor, which contains information that is not yet covered by existing multi-factor asset pricing models within international

equity markets. Otherwise, it would be reasonable to conclude that capital share growth is a simple proxy for the exposure to other, already well-known risk factors (Letttau et al., 2019).

To this end, in line with Lettau et al. (2019), we add estimated betas from alternative factor models within the cross-sectional regressions. For instance, we ordinarily estimate Fama-French three factor betas locally and then add estimated local KS risk exposure as follows:

$$\left(R_{j,t}^e\right) = \lambda_0 + \hat{\beta}_{j,KS,H}\lambda_{KS,H} + \hat{\beta}_{j,MKT}\lambda_{MKT} + \hat{\beta}_{j,SMB}\lambda_{SMB} + \hat{\beta}_{j,HML}\lambda_{HML} + \epsilon_{j,t} \quad (3.4)$$

Beyond the Fama-French three-factor model, we test for the traditional CAPM by Sharpe (1964) and Lintner (1965, 1969), the four-factor model by Carhart (1997), the five-factor model by Fama and French (2015) as well as the q-factor model proposed by Hou et al. (2015). Estimated capital share betas are calculated over two-year (H8) horizons.

In accordance with Lettau et al. (2019), we show results obtained for Size/BM portfolios for each of our chosen regions. Another reason for displaying results for the Size/BM portfolios only is that these portfolios together with portfolios sorted on long-term reversal (REV) exhibit highest statistical explanatory power of KS across all panels and time horizons as shown within Table 3.2.

Table 3.3 shows baseline findings from capital share growth as an alternative asset pricing factor in international equity markets. We start by summarizing findings for G7+Australia (excl. the U.S. market). Subsequently, we show aggregate results for Emerging Markets. Lastly, we summarize findings obtained for our World Sample (excluding the U.S. market). Our results obtained for the U.S. market imply that $\lambda_{KS,H}$ is a highly significant risk factor that almost entirely subsumes information contained in other factor model specifications. These results are consistent with findings reported by Lettau et al. (2019). For brevity reasons, we thus do not display U.S. results within Table 3.3.

Table 3.3: Capital Share Growth in Alternative Factor Models – Size/BM

This table reports cross-sectional regressions of average equity characteristic portfolio returns for 5x5 portfolios sorted on Size/BM on factor betas. Local capital share betas are estimated over horizons of two years (H8). All coefficients are multiplied by 100. Corresponding t-statistics are indicated within parentheses. Monthly returns are winsorized at the 0.01% and 99.9% levels. Our sample runs from M1:1989 to M12:2017. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% level. We show results for Size/BM sorted equity portfolios only due to reasons of brevity.

G7 + Australia (Excl. USA)						
Panel A: CAPM						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	R^2		
2.14*** (2.77)				0.25		
	-20.59*** (-4.24)			0.4		
1.21* (1.73)	-17.13*** (-3.38)			0.51		
Panel B: FF 3 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	R^2		
	-23.05** (-2.10)	1.71 (1.59)	0.34 (0.16)	0.45		
1.50* (1.64)	-20.19* (-1.88)	1.93* (1.85)	-1.60 (-0.68)	0.51		
Panel C: Carhart 4 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	WML	R^2	
	17.10 (0.95)	2.88* (1.86)	-1.51 (-0.46)	-0.49 (-0.04)	0.21	
1.90** (2.44)	13.45 (0.83)	2.22 (1.57)	-1.57 (-0.53)	-1.11 (-0.09)	0.40	
Panel D: FF 5 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	RMW	CMA	R^2
	-25.37** (-1.99)	0.29 (0.20)	0.58 (0.23)	-6.56 (-0.99)	7.72 (1.14)	0.53
1.55* (1.71)	-20.24 (-1.62)	0.78 (0.54)	-1.91 (-0.68)	-6.14 (-0.98)	8.63 (1.33)	0.59
Panel E: q-factors						
$\lambda_{KS,H}$	$MKTRF$	ME	IA	ROE	R^2	
	-22.52* (-1.80)	-0.43 (-0.10)	5.43 (0.99)	-3.84 (-0.41)	0.47	
1.53* (1.78)	-20.21* (-1.69)	-0.43 (-0.10)	1.09 (0.19)	-3.83 (-0.43)	0.54	

Table 3 (Cont'd): Capital Share Growth in Alternative Factor Models – Size/BM

This table reports cross-sectional regressions of average equity characteristic portfolio returns for 5x5 portfolios sorted on Size/BM on factor betas. Local capital share betas are estimated over horizons of two years (H8). All coefficients are multiplied by 100. Corresponding t-statistics are indicated within parentheses. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% level.

Emerging Markets						
Panel A: CAPM						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	R^2		
0.90*** (6.22)				0.63		
	-15.79*** (-4.43)			0.46		
0.68*** (4.41)	-8.17** (-2.58)			0.71		
Panel B: FF 3 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	R^2		
	-21.80*** (-3.66)	0.00 (0.00)	4.99*** (3.35)	0.67		
0.56*** (3.39)	-17.31*** (-3.43)	-0.41 (-0.54)	2.39* (1.66)	0.79		
Panel C: Carhart 4 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	WML	R^2	
	-15.64** (-2.47)	-0.75 (-0.89)	8.11*** (5.52)	-12.01*** (-3.01)	0.77	
0.37** (1.99)	-14.86** (-2.51)	-0.74 (-0.95)	5.52*** (2.92)	-7.43* (-1.69)	0.81	
Panel D: FF 5 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	RMW	CMA	R^2
	-15.69** (-2.29)	0.92 (1.16)	2.89** (1.96)	-2.43 (-0.59)	9.51*** (2.98)	0.77
0.39** (2.04)	-14.77** (-2.32)	0.47 (0.61)	1.80 (1.23)	0.34 (0.09)	7.25** (2.30)	0.81
Panel E: q-factors						
$\lambda_{KS,H}$	$MKTRF$	ME	IA	ROE	R^2	
	-8.26 (-1.54)	-0.42 (-0.33)	1.65 (0.51)	-17.28*** (-4.55)	0.75	
0.3836 (1.58)	-9.62* (-1.84)	-0.80 (-0.64)	2.05 (0.65)	-11.05** (-2.05)	0.78	

Table 3 (Cont'd): Capital Share Growth in Alternative Factor Models – Size/BM

This table reports cross-sectional regressions of average equity characteristic portfolio returns for 5x5 portfolios sorted on Size/BM on factor betas. Local capital share betas are estimated over horizons of two years (H8). All coefficients are multiplied by 100. Corresponding t-statistics are indicated within parentheses. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% level.

World (Excl. USA)						
Panel A: CAPM						
$\lambda_{KS,H}$	$MKTRF$					R^2
1.56*** (6.04)						0.61
	-13.44*** (-4.19)					0.43
1.20*** (4.18)	-6.57** (-2.13)					0.68
Panel B: Fama/French 3 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML			R^2
	-24.29*** (-4.04)	-0.30 (-0.34)	3.38** (2.49)			0.64
1.12*** (4.08)	-20.09*** (-4.30)	-1.03 (-1.50)	1.01 (0.86)			0.80
Panel C: Carhart 4 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	WML		
	-22.76*** (-3.06)	-0.51 (-0.53)	6.04*** (2.98)	-8.78 (-1.20)		
1.13*** (3.84)	-20.73*** (-3.60)	-1.19 (-1.55)	2.72 (1.53)	-4.26 (-0.74)		
Panel D: Fama/French 5 factors						
$\lambda_{KS,H}$	$MKTRF$	SMB	HML	RMW	CMA	R^2
	-14.39** (-2.46)	0.92 (1.38)	0.40 (0.31)	0.46 (0.11)	18.58*** (4.44)	0.82
0.66** (2.16)	-15.49*** (-2.88)	0.30 (0.44)	-0.28 (-0.23)	2.50 (0.60)	13.85*** (3.14)	0.85
Panel E: q-factors						
$\lambda_{KS,H}$	$MKTRF$	ME	IA	ROE		
	-13.49** (-2.26)	-2.81** (-2.03)	5.17 (1.51)	-20.52*** (-3.88)		
1.03*** (3.58)	-16.34*** (-3.41)	-3.22*** (-2.92)	4.98* (1.84)	-0.14* (-1.74)		

As shown within Table 3.3, the explanatory power of $\lambda_{KS,H}$ for G7+Australia (excl. U.S.) is substantially reduced once adding alternative factor betas. While being initially significant at the 1% level (t -statistics: 2.77) over one-year horizons, the KS risk estimate is left only marginally relevant (t -statistics smaller than two) within four out of five alternative model specifications. The only exception remains the Carhart four-factor model (Panel C). Here, $\lambda_{KS,H}$ remains significant at the 5% level (t -statistics: 2.44). This finding implies that, unlike the U.S. market, the information contained within $\lambda_{KS,H}$ is almost entirely subsumed by alternative factors. This pattern also holds for portfolios sorted on long-term reversal (REV) as well as all other equity portfolio sorts. Even if having statistically significant explanatory power for certain equity characteristic portfolio groups, the information contained within the capital share factor seems to be almost entirely subsumed by existing factor specifications.

For Emerging Markets, our results shown in Table 3.3 imply that the information of KS risk estimates is not subsumed by alternative factor betas but rather adds additional explanatory power to these model specifications. Once adding the capital share factor, $\lambda_{KS,H}$ remains statistically significant at the 5%-level across four out of five alternative factor model specifications. For instance, cross-sectional R^2 increases from 0.46 to 0.71 once adding $\lambda_{KS,H}$ to the CAPM, whereas the KS factor alone explains roughly 63% in the cross-section. This pattern also holds for other equity characteristic portfolio returns within Emerging Markets (not shown for brevity issues).

Similar to Emerging Markets, we find that adding $\lambda_{KS,H}$ as a risk factor increases cross-sectional R^2 across all equities for our World Sample. Within four out of five factor model specifications, $\lambda_{KS,H}$ remains even relevant at the 1% level in explaining cross-sectional portfolio returns. Unlike the U.S. market, though, the information contained within capital share growth does not subsume the information contained within other factors in international equity markets.

Taken together, we interpret our results reported in Tables 3.2 and 3.3 as a clear affirmation of our first hypothesis implying that there exists (strong) geographic heterogeneity in the explanatory power of local capital share growth for expected equity characteristic portfolio returns.

3.4.3 Global Capital Share Growth in International Equity Markets

We continue the analysis by testing our second hypothesis, implying that capital share growth is a local rather than global risk factor in a quest to further analyze the existence of and potential causes for international market segmentations. To construct a global KS factor, we apply market capitalization-weighted annual capital share values of our full set of applied countries (including the U.S.). For analyses described below, we again follow the two-step regression approach by [Lettau et al. \(2019\)](#). In line with Section 3.4.1, we estimate first-step regression betas for each of our 5x5 equity portfolios using country-fixed effects. Reported results are, however, robust to applying country fixed-effects within the second rather than first regression step following the methodology proposed by [Lettau et al. \(2019\)](#).

Table 3.4 summarizes corresponding findings. Panel A reports findings for the U.S. market. Panel B summarizes the empirical evidence for G7 + Australia (excluding the U.S. market). Panel C shows results for Emerging Markets. Finally, Panel D presents insights for our World Sample (excluding the U.S. market).

Results within Panel A indicate that our global KS factor is statistically significant in explaining U.S. equity portfolio returns. While results for one-year investment horizons remain statistically significant, they are, however, in part substantially reduced. Over two-year investment horizons, U.S. results seem to change only slightly. An exception remains the Size/Mom portfolio which increases in both, size and significance, over one and two-year investment perspectives.

Similar to the U.S. market, we observe for G7 + Australia (excl. U.S.), shown in Panel B, that our global risk factor is able to explain a large fraction of expected equity portfolio returns. Consistent with results reported in Table 3.2, however, the global factor is unable to explain expected returns of portfolios sorted on Size/REV as well as of portfolios sorted on Size/Mom. Across all equity portfolio sorts the global KS risk estimate amounts to 0.5039 (t -statistics: 2.18) with a corresponding R^2 of 0.24 over one-year horizons. Accordingly, over two-year horizons we report an average $\lambda_{KS,H}$ of 0.5975 (t -statistics: 2.63) and a corresponding R^2 of 0.26.

Conversely, as further shown within Table 3.4, the explanatory power of capital share growth decreases considerably once applying a global rather than local risk factor for Panels C and D. At an aggregate level (average values across all equity portfolio sorts), results for Emerging Markets (Panel C) as well as our World Sample (Panel D) become insignificant. These findings underpin our hypothesis that capital share growth is a local risk factor. Therefore, our findings are in line with studies by Fama and French (2012), Hollstein (2020) as well as Hou et al. (2011) who emphasize the importance of local factors in explaining expected returns across international markets.

3.4.4 Cross-Country Analyses: Determinants of Capital Share Risk Estimates

3.4.4.1 Country Characteristics

An important aspect in empirical asset pricing is to understand why risk factors matter within certain countries while they are irrelevant within other countries. In this section, we thus strive to understand and test potential drivers of global differences in reported capital share risk estimates. In accordance with our third and forth hypotheses derived and outlined in Section 3.2 as well as prior cross-country studies, we identify the following set of variables described below. Detailed variable definitions and data sources are given in Appendix B of the Electronic Supplementary Material.

First, and in order to test for the theoretical explanation given by Lettau et al. (2019), we

apply multiple variables to account for cross-country differences in private wealth inequality. Our first proxy is a country's domestic investor participation rate. This figure stands for the fraction of domestic households holding directly stocks in the domestic stock market and is taken from [Barniv et al. \(2010\)](#) in accordance with [Giannetti and Koskinen \(2010\)](#). Further, we account for a country's top 1% and bottom 50% income shares. The logic behind inclusion of these two variables is that [Lettau et al. \(2019\)](#) hypothesize a positive relationship between the top 1% income share and corresponding capital share values. Relatedly, we test for a potential negative link between a country's bottom 50% income share and corresponding capital share values. The fourth proxy illustrates the average of a country's private financial assets. Lastly, we include the portion of local population that is covered by at least one social protection benefit as we consider social protection benefits to partially offset implications of private wealth inequality.

Second, we apply a set of four variables to account for a country's overall economic opaqueness: An inequality transparency index, the average size of local shadow economies, an opacity index, and the number of news articles per country. The wealth inequality transparency index is gathered from the World Inequality Database accessible online at <https://wid.world> and measures the availability as well as quality of information on wealth inequality across countries. The average size of local shadow economies is a measure taken from [Schneider et al. \(2010\)](#), indicating the average size of national shadow economies in percent of official GDP. The opacity index is a proxy for local information opaqueness which we gather from [Kurtzman et al. \(2004\)](#). It measures the extent to which a country lacks clear and accurate as well as widely accepted business practices among firms, investors and governments. Lastly, the number of news articles is taken from [Griffin et al. \(2011\)](#) and stands for the number of news articles scaled by the number of firms per country.

Third, to account for the bargaining power of workers and employees, we include the following proxies which we obtain from the International Labour Organisation (ILO): Collective bargaining, trade union density as well as the number of days not worked by 1,000 workers. Collective bargaining stands for the fraction of employees covered by one or more collective agreement. Trade union density indicates the number of union members who are employees

as a percentage of the total number of employees. Finally, we apply the number of days not worked because of strike per 1,000 workers. As outlined within Section 3.2, the rationale beyond inclusion of these three proxies is that we hypothesize that collective bargaining as well as trade unions increase the risk of potential shocks which redistribute the income between shareholders and employees.

Lastly, we include two proxies for differences in public wealth and public reserves: Total Nonfinancial Government Assets as well as Total Reserves by GDP. The first variable comprises produced assets as for instance buildings and inventories as well as nonproduced assets such as natural resources (e.g. oil or gas). The latter comprises the holdings of monetary gold, special drawing rights, reserves of IMF members held by the IMF as well as holdings of foreign exchange which are controlled by monetary authorities.

3.4.4.2 Cross-Sectional Regressions

To analyze the relationship between obtained capital share risk estimates and our chosen set of country variables, we perform various cross-sectional regression specifications. In the regression analyses below, our dependent variable constitutes the average capital share risk estimate $\lambda_{KS,H}$ across applied equity characteristic portfolio sorts as reported within the last column of Table 3.2.

Our chosen set of country characteristics reported in Section 3.4.4.1 primarily exhibits a panel data structure. Occasionally, however, they only exhibit the country-dimension, implying that they are time-invariant. Thus, we apply averages of time-series country variables and exclusively apply cross-sectional regressions below. In doing so, we are able to analyze drivers of differences in cross-country capital share growth risk estimates.

Table 3.5 reports empirical results for various cross-sectional regression specifications.

Table 3.5: Cross-Country Analyses of Average KS Risk Estimates

This table shows results of cross-sectional regressions examining the significance of our chosen set of country variables. The dependent variable is the within-country time-series average estimate of two-year capital share growth risk estimates as reported within Table 2. Reported estimates are multiplied by 100. t -statistics are shown within parentheses. *, **, and *** indicate statistical significance at the 10%, 5% as well as 1% level.

	M1	M2	M3	M4	M5	M6	M7	M8	M9	M10	M11	M12	M13	M14
	Dependent Variable: Average two-year KS risk estimates (H8)													
Investor Participation Rate	-0.00834 (-0.11)													
Top 1% Income Share		0.39343 (0.51)												
Bottom 50% Income Share			-0.42383 (-0.56)											
Net Private Financial Wealth				-0.97800* (-1.75)										
Social Security Coverage					-0.00431* (-1.88)									
Wealth Transparency Index						0.01223 (0.65)								
Shadow Economy							-0.00264 (-0.72)							
Opacity Index								-0.00418 (-0.91)						
Number of News Articles									0.00269* (1.79)					
Collective Bargaining										-0.00048 (-0.34)				
Trade Union Density											-0.00053 (-0.22)			
Days not Worked by 1,000 Workers												0.00028 (1.23)		
Nonfinancial Government Assets													-0.00123* (-1.79)	
Total Reserves by GDP														-0.51884** (-2.10)
R^2	0.00	0.01	0.01	0.15	0.10	0.01	0.01	0.03	0.10	0.00	0.00	0.07	0.14	0.10

As illustrated within Table 3.5, our chosen set of proxies for economic opaqueness as well as relative bargaining strength of local workforce are insignificant in explaining cross-country differences in average capital share growth risk estimates, with the exception of a country's number of news articles. This proxy, which is significant at the 10% level, implies that there exists a positive link between the number of news articles and a country's average capital share risk estimates.

Conversely, though, two of our chosen variables to proxy for wealth inequality are significant at the 10% level: Net private financial wealth as well as social security coverage. These variables are however inversely related to a country's average KS risk estimates. This finding underpins the theoretical backdrop given by Lettau et al. (2019) in an international setting: The smaller the level of private wealth inequality, the lower the amount of capital market risk beard by a small subset of investors, thus the lower the expected compensation in terms of equity portfolio returns. The insignificance of the proxy for (domestic) investor participation rate does not contradict this rationale. For instance, while the investor participation rate within the U.S. is among the highest worldwide, most households who participate in the stock market own very little stocks (Lettau et al., 2019). That is, the participation rate itself does not give information about the level of risk-sharing prevalent in the equity market.

Further, reported results shown in Table 3.5 illustrate the significance of our proxies for differences in public wealth and public reserves. The variable *Total Nonfinancial Government Assets* is significant at the 10% level, whereas the proxy *Total Reserves by GDP* is significant at the 5% level in explaining cross-country differences in average KS risk estimates. This finding is in accordance with the theoretical motivation taken from Alvaredo et al. (2018), implying that a decline in public wealth is associated with a decline in the government's ability to redistribute income or to offset rising inequality. Beyond, this finding is in line with empirical evidence reported by Dominguez et al. (2012) who illustrate that countries which exhibit higher levels of total reserves are found to simultaneously exhibit stronger economic recovery after the global financial crisis, thus partially offsetting adverse effects for security holders.

Overall, we interpret results shown within Table 3.5 as an affirmation of our third hypothesis,

implying that cross-country differences in the explanatory power of capital share growth are driven by variations in local wealth inequality. Simultaneously, however, we discard our fourth hypothesis as the only significant variables remain the number of news articles scaled by the number of firms within a country as well as those variables proxying for differences in public wealth as well as public reserves. Conversely, proxies applied for the relative bargaining strength of local workforce as well as the majority of variables included to account for economic opaqueness are not able to explain cross-country differences in KS risk estimates.

3.4.5 Cross-Country Trading Strategy

Can we exploit the information contained within capital share growth by a simple trading strategy? To this point, reported capital share risk estimates are entirely based upon in-sample calculations. To assess the relative importance and robustness of the capital share growth factor, we thus proceed by constructing an out-of-sample investment strategy. Specifically, we perform rolling ten-year-window regressions of cumulative quarterly equity characteristic portfolio returns on local capital share growth for each portfolio style (5x5 Size/BM portfolios, 10 REV portfolios, 5x5 Size/INV portfolios, 5x5 Size/OP portfolios, 5x5 Size/REV portfolios, 5x5 Size/Mom portfolios), on a country-basis. This procedure is consistent with the methodological approach described in Section 3.3.2, with the exception that we now conduct country estimates. In a next step, we go long the decile of past capital share growth beta winners and short the decile of past capital share growth beta losers for the following investment quarter.

Table 3.6 summarizes average quarterly long-short returns. Additionally, we report corresponding sharpe ratios, quarterly minimum returns, and standard deviations. Investment strategies are built for the U.S. market only, G7 + Australia (excl. the U.S.), Emerging Markets as well as for a World Sample (excl. the U.S.) and across the full range of applied countries. Panel A shows findings when applying one-year horizons for cumulative portfolio returns as well as capital share growth estimates. Panel B summarizes findings for two-year horizons.

Table 3.6: Out-of-Sample Long-Short Portfolio Returns

This table reports portfolio returns obtained from long-short investment strategies based upon local capital share growth betas (first step regression beta as described by Lettau et al. (2019)). We perform rolling ten-year-window regressions of cumulative quarterly equity characteristic portfolio returns on local capital share growth for each portfolio style (5x5 Size/BM portfolios, 10 REV portfolios, 5x5 Size/INV portfolios, 5x5 Size/OP portfolios, 5x5 Size/REV portfolios, 5x5 Size/Mom portfolios), for each country, respectively. In a next step, we go long the decile of past capital share growth beta winners and short the decile of past capital share growth beta losers for the following investment quarter. Returns are winsorized at the 0.01% and 99.9% levels. The sample runs from M1:1989 to M12:2017.

Panel A: One-Year Capital Share Growth (H4)				
	ret quarterly	sharpe ratio	min ret	sd
USA	3.13% (1.22)	0.27	-7.72%	11.48%
G7 + Australia (Excl. USA)	0.96% (0.35)	0.08	-19.51%	12.31%
Emerging Markets	-1.10% (-0.29)	-0.06	-45.34%	17.09%
World Excl. USA	-0.19% (-0.10)	-0.02	-17.94%	8.43%
World Incl. USA	-0.16% (-0.09)	-0.02	-18.03%	8.41%
Panel B: Two-Year Capital Share Growth (H8)				
	ret quarterly	sharpe ratio	min ret	sd
USA	1.06% (0.36)	0.08	-37.58%	13.30%
G7 + Australia (Excl. USA)	-1.53% (-0.53)	-0.12	-19.29%	12.96%
Emerging Markets	-2.39% (-0.54)	-0.12	-67.76%	19.81%
World Excl. USA	0.44% (0.26)	0.06	-12.83%	7.48%
World Incl. USA	0.51% (0.30)	0.07	-13.99%	7.48%

As shown within Table 3.6, returns obtained from our out-of-sample investment strategy illustrate that investors are not able to exploit information from this single capital share growth factor in real time. Across all regions, even within the U.S. market, our strategy yields insignificant quarterly portfolio return estimates. This finding holds for both, one-year and two-year capital share growth estimates. Our results simultaneously show that sharpe ratios are below the critical value of one while minimum returns and standard deviations are considerably high.

At an aggregate level, this finding exemplifies that - while exhibiting great in-sample explanatory power and significance - the information contained in the capital share growth factor seems difficult to be exploited out-of-sample.

3.5 Conclusion

Our study analyzes whether growth in the capital share (KS) of aggregate income (GDP) can explain equity portfolio returns in international stock markets as proposed by Lettau et al. (2019) for the U.S. market.

Reported results imply that local capital share growth is a priced risk factor across international markets. There exists, however, strong geographic heterogeneity in the explanatory power of this risk factor. Applying country fixed-effects regressions, we find that capital share growth has particularly high explanatory power for equity returns obtained within the United States as well as within Emerging Markets. Unlike the U.S. market, though, the information contained in the KS risk factor of Emerging Markets does not subsume information contained in alternative factor models as for instance the Fama-French three factor or q-factor models, but rather adds additional explanatory content to these model specifications. For G7 + Australia (excl. U.S.), we find that even if having statistically significant explanatory power for certain equity characteristic portfolio groups, the information contained within the capital share factor seems to be almost entirely subsumed by existing factor model specifications.

Furthermore, our results imply that capital share growth is a local, rather than global risk

factor. Once applying global capital share growth estimates, the explanatory power of this factor either disappears entirely or is at least substantially reduced. Beyond, cross-sectional regression analysis reveals that country differences in KS risk estimates can best be explained by cross-country variations in private wealth inequality as well as country-differences in public wealth/reserves. Simultaneously, we report that proxies for economic opaqueness as well as relative bargaining strength of local workforce are primarily insignificant in explaining cross-country differences in average capital share growth risk estimates.

Finally, we are able to illustrate that - while exhibiting great in-sample explanatory power and significance - the information contained in the capital share growth factor seems difficult to be exploited out-of-sample. An investment strategy that goes long the decile of past capital share growth beta winners and short the decile of past capital share growth beta losers, yields insignificant results across international markets as well as even for the U.S. market only.