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# Carbon Dioxide and Asset Pricing: Evidence from International Stock Markets*

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## Abstract

We use carbon dioxide ($CO_2$) emissions growth to measure consumption risk within a consumption-based capital asset pricing model framework. Given the comprehensive worldwide coverage of $CO_2$ emissions, this measure allows us to use the full history of stock market data in the US, Europe, the world, and fifteen international markets. For the US (Europe/the world), we are able to explain the observed equity market premium with a relative risk aversion of 6 (10/12), which is less than half the size of that estimated using the canonical expenditures-based consumption growth measure. The average estimated relative risk aversion across fifteen other international markets is 5. We also find evidence that the growth of $CO_2$ emissions is a priced risk factor that captures the cross section of stock portfolio returns.

*JEL Classification*: G12, Q43
*Keywords*: International Asset Pricing, Consumption-Based Capital Asset Pricing Model, Carbon Dioxide Emissions

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A fundamental question in asset pricing is how macroeconomic risks, such as consumption risk, relate to the pricing of financial assets in the time series and the cross section. In the consumption-based capital asset pricing model (C-CAPM), introduced by Lucas (1978) and Breeden (1979), investors require risk compensation for holding assets that comove with consumption growth. Therefore, cross-sectional variations in expected returns are driven by the covariances between asset returns and household consumption growth. Despite its theoretical simplicity, many studies have shown that the C-CAPM does not fit the empirical data well. An unreasonably high level of relative risk aversion is required to generate the observed risk premium when household expenditure on nondurable goods and services is used to measure households' consumption (Mehra and Prescott, 1985); the model-implied risk-free rate is too large (Weil, 1989); and, the consumption growth poorly prices the cross section of stock returns.

Researchers have done a tremendous amount of work to validate the C-CAPM and solve the "equity premium puzzle". Theorists have proposed delicately-developed models, including the separation of the elasticity of intertemporal substitution and risk aversion (see., e.g., Epstein and Zin, 1989; Weil, 1989), the slow-moving long-run consumption risk (Bansal and Yaron, 2004), and the habit persistence model (Campbell, 1999b). On the empirical side, researchers have tried to capture consumption variation from many different perspectives, including the ultimate consumption risk over a longer horizon (Parker and Julliard, 2005a), stock market participants' long-run consumption risk (Malloy, Moskowitz, and Vissing-Jørgensen, 2009a), and the year-end effect of household consumption growth (Jagannathan and Wang, 2007a). In addition, some recent studies use creative alternatives to measure consumption, which yields higher volatility for consumption growth and better empirical results (see., e.g., Savov, 2011; Da, Yang, and Yun, 2015; Chen and Lu, 2018). In the context of this research landscape, our study gravitates towards the latter category, exploring innovative measures of consumption growth.

In this study, we employ carbon dioxide ($CO_2$) as a proxy for household consumption to conduct a comprehensive analysis of consumption risk's role in explaining the equity risk premium across diverse international stock markets over extended sample periods. Modern life heavily relies on energy consumption, which is closely tied to $CO_2$ emissions. These emissions stem from a wide array of household consumption activities, encompassing both direct and indirect energy consumption, such as housing operations, transportation, food, and apparel. The robust relationship between $CO_2$ emissions and household consumption is well-established in the ecology and energy economics literature. Our research demonstrates a robust correlation between $CO_2$ emissions growth and stock market returns, with $CO_2$ exhibiting greater volatility compared to the conventional expenditures-based consumption growth measure. Consequently, $CO_2$ emerges an effective instrument for explaining the equity premium and cross-sectional return variation within the traditional C-CAPM framework with CRRA utility function.

In this paper, we significantly expand the scope of C-CAPM analysis by a broader range of markets over sample periods that significantly surpass those attainable through traditional consumption expenditure data or alternative measures. While most existing C-CAPM studies primarily focus on the US market from 1929 due to data constraints, our research extends the US sample by 57 years and includes 15 international markets, 11 of which have samples exceeding 100 years.[1] By investigating C-CAPM across these diverse markets and extended timeframes, our paper offers valuable insights and a deeper understanding of the model's validity and implications, especially given that equity premia are suggested to be higher in earlier sample periods (see., e.g., Heaton and Lucas, 1999; Jagannathan, McGrattan, and Scherbina, 2001).

Our empirical findings offer a valuable contribution to the C-CAPM literature. The $CO_2$-emissions-based consumption risk measure provides key insights into understanding the joint equity risk premium and the implied risk-free rate, especially in historical contexts.

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[1] Table 1 provides a clear comparison between the sample coverage used in this paper and that of existing papers that estimate C-CAPM using consumption expenditure or alternative consumption measures.

Employing annual $CO_2$ emissions as a proxy for household consumption, we observe a relative risk aversion (RRA) coefficient of 6 and a small implied real risk-free rate of 0.63% in the US market within the C-CAPM framework over the extended sample period of 1872 to 2015. Notably, the RRA estimate is less than half of those estimated using the traditional expenditures-based counterparts. In an international context, the $CO_2$-emissions-based measure yields an average RRA of 5.33 across fifteen countries. Yet, the increased focus on greenhouse gas emissions and changes in household consumption patterns in recent decades might impinge on the effectiveness of $CO_2$ as the consumption measure, resulting in reduced explanatory power and even negative implied risk-free rates in certain countries. Furthermore, $CO_2$-emissions-based consumption plays a pivotal role in explaining the cross section of stock return variations. Specifically, in the context of the US 25 Fama-French portfolios, $CO_2$ growth delivers a positive and significant price of risk, coupled with the lower pricing error and root-mean-square error (RMSE) compared to the expenditure-based measure, irrespective of weather the market factor is taken into account.

One potential limitation of our $CO_2$ emission measure is that a portion of $CO_2$ emissions may come from industrial production and/or physical investments. We show that the pricing power of $CO_2$ emission growth does not predominately arise from the part that correlates with industrial production or private non-residential fixed investment. In addition, although higher $CO_2$ emissions may result in rising temperature in the long run and thus cause long-run variation in climate risk (see e.g., Bansal, Kiku, and Ochoa, 2016; Bansal, Ochoa, and Kiku, 2017), the pricing power of $CO_2$ emission growth comes from its ability to capture the relatively short-run variation in household consumption.

The rest of the paper is organized as follows. Section 1 provides a rationale for the choice of using $CO_2$ as a new measure for consumption. Section 2 provides a description of the data. It also details the construction of the annual per capita growth rate of $CO_2$ emissions. We reinvestigate the equity premium puzzles in the US and in international stock markets in Section 3 using the $CO_2$ emissions-based measure. In Section 4, we implement cross-sectional asset pricing tests using $CO_2$ as an alternative measure of consumption. Section 5 provides additional discussions about other potential confounding effects that might contaminate the effect of $CO_2$-measured consumption risk on asset prices. Section 6 concludes.

# 1 $CO_2$ emissions as a proxy for consumption

Literature in ecology and energy economics supports the view that household consumption is the main driver behind $CO_2$ emissions.[2] For example, Bin and Dowlatabadi (2005) show that more than 80% of the $CO_2$ emitted in the US is the consequence of consumer demands and the related economic activities to support these demands. In addition, Pottier (2022) estimates household expenditure elasticities of $CO_2$ emissions to be between 0.81 to 1.14 across various countries/regions, suggesting a stable global relationship between household expenditure and $CO_2$ emissions.

Compared with traditional personal expenditures and other alternative consumption measures, our $CO_2$-emissions-based consumption measure has several advantages in capturing household consumption in a more comprehensive manner. First, $CO_2$ emissions capture a broad range of energy consumption, including but not limited to electricity consumption as in Da et al. (2015). Fossil fuels, the usage of which generates a significant amount of $CO_2$, play an important role in electricity generation in the US. In 2018, around 64% of the electrical energy generated used fossil fuels. The time series of $CO_2$ emissions should not only incorporate movements in electricity consumption but also contain more information about other types of energy consumption caused by household consumption activities.

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[2] Studies indicate that while end-uses of home energy and private transportation contribute to between 13% and 35% of a country's total direct greenhouse gas (GHG) emissions, the number increases to 60% to 80% once indirect household emissions are included (Benders, Kok, Moll, Wiersma, and Noorman, 2006; Kok, Benders, and Moll, 2006; Larsen and Hertwich, 2010; Moll, Noorman, Kok, Engström, Throne-Holst, and Clark, 2005; Nansai, Inaba, Kagawa, and Moriguchi, 2008; Nijdam, Wilting, Goedkoop, and Madsen, 2005; Peters and Hertwich, 2006; Weber and Matthews, 2008). At the global level, 72% of GHG emissions are related to household consumption, 10% to government consumption, and 18% to investments (Hertwich and Peters, 2009).

Second, $CO_2$ emissions capture the transportation component of household consumption. Households have been evolving toward a lifestyle with more travel and leisure activities. Households are also spending more on services that involve an intensive use of transportation. Expenditures related to transportation, however, are difficult to capture by measures like garbage generation (Savov, 2011) or electricity usage (Da et al., 2015). Our data directly include emissions from the consumption of petroleum used in transportation, therefore capturing changes in transportation-related household consumption.

Third, $CO_2$ emissions account for the housing component of household consumption. Households spend a significant portion of income on housing-related consumption. According to the US Department of Labor Statistics' Consumer Expenditures Survey, in 1984, around 16% of household consumption expenditures belong to shelter, which includes property rental expenses and/or mortgage payments. This number has gradually increased to 20% in 2018. $CO_2$ emissions can indirectly address this issue in the way that larger houses typically have more household activities that induce more emissions. Our $CO_2$ emissions also capture consumptions related to housing by including $CO_2$ emissions from cement production and emissions involved with the production and transportation of housing construction. Housing-related expenditures are closely related to the growth of new construction and thus the consumption of cement. Cement manufacturing processes release $CO_2$ when calcium carbonate is heated, generating lime and $CO_2$ in the process. The production of other building materials and the transportation of these materials are petroleum based, meaning that they are made from crude oil, a process that induces $CO_2$ emissions. By including these elements, we can better capture movement in housing-related consumption expenditures using $CO_2$ emissions.

We use $CO_2$ emissions as a proxy for household consumption flow, including both non-durable goods and services, and the service flow of durable goods, under the traditional C-CAPM framework. This approach, not decomposing consumption into its components, enables us to test the C-CAPM model across a broader range of markets and over longer sample periods than possible with traditional consumption data, particularly for durable goods, which are limited mainly to post-1970 US data. Our study contrasts with Chen and Lu (2018), who employ $CO_2$ to extract risk related to time varying durable goods usage in periods characterized by a high proportion of consumption on energy-dependent durable goods, particularly post-1970. While both studies contribute to the C-CAPM literature, they utilize distinct preference functions. Specifically, our paper operates within a traditional C-CAPM with CRRA utility, whereas Chen and Lu (2018) assume a more complex Epstein and Zin (1991) recursive preferences, which allow for the separation between relative risk aversion and elasticity of intertemporal substitution. Recognizing post-1970 complexities in $CO_2$ as a consumption measure, increasing global environmental consciousness, advancements in energy efficiency, and shifts in household consumption patterns including changes in consumption composition and the influx of foreign product, our paper finds $CO_2$ to be a noisier measure of overall consumption in this period, aligning with our results. We build on this understanding by applying $CO_2$ as a proxy of consumption across a broader historical spectrum, thereby offering unique insights into the empirical validity of C-CAPM over an extended timeframe across diverse markets. However, as shown in Chen and Lu (2018), these changes do not prevent $CO_2$ from being effective in extracting key information about the time varying utilization of durable goods, even amidst its increased noisiness as a measure of overall consumption post-1970. Our paper complements the insights of Chen and Lu (2018), contributing to a deeper understanding of $CO_2$'s role in measuring consumption risk.

## 2 Data

In this paper, we use $CO_2$ emissions to proxy for consumption. The $CO_2$ emissions data we use are commonly used in studies of $CO_2$ emissions and have been constructed following the procedures discussed in Marland and Rotty (1984) and Boden and Andres (1995). The data are sourced from the Oak Ridge National Laboratory (ORNL) for the sample prior to 2014 and from the Global Carbon Project for 2015–2016. Emissions data from these two sources are constructed using the same raw data and are based on the same methodology. The change merely reflects a change of its host. These data provide $CO_2$ emissions from aggregate fossil fuel consumption and cement manufacture at an annual frequency over 200 countries worldwide. Quantities of $CO_2$ emissions are measured in the standard unit of 1,000 metric tons of carbon. The time series of $CO_2$ emissions is constructed by applying $CO_2$ emissions conversion coefficients to historical records of energy consumption series.[3] Specifically, $CO_2$ emissions of fuel type $i$ are estimated as the product of three terms: quantity consumed of fuel type $i$, the carbon content of fuel type $i$, and the fraction of the carbon content that is oxidized.[4] Quantities of fuel consumption are controlled for by changes in the form of fuel, fuel imports and exports, and changes in fuel stocks. They provide good estimates for the amount of fuel that generates $CO_2$ emissions as the result of people's consumption.

One key advantage of the $CO_2$ data set, besides it being the commonly used data set in studies of $CO_2$ emissions, is that it provides an exceptionally long record of $CO_2$ emissions for all developed countries and most developing countries tracing back to 1751. The coverage, in both length and breadth, exceeds that of available stock returns data. The long and comprehensive coverage enables us to exploit the full sample of stock market data in a wider range of countries, in addition to looking at some key regions, including the US, Europe, and the world. A consumption measure constructed based on the $CO_2$ emissions thus would allow us to investigate the long-run performance of the C-CAPM and the performance of the C-CAPM in international markets.

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[3] Andres, Fielding, Marland, Boden, Kumar, and Kearney (1999) provide details on the contents and processing of the historical energy statistics from 1800 to 1949. The 1950 to 2016 $CO_2$ emission estimates are derived from energy statistics published by the United Nations. The US Bureau of Mines compiles the cement manufacturing data.

[4] In their estimation methodology, Marland and Rotty (1984) assume the fraction of carbon content and the fraction oxidized to be constant over time. Although the carbon content of fuel has not varied considerably since the nineteenth century, the components of the fraction oxidized do vary because of improvements in combustion efficiencies, as well as nonfuel usage, including appreciable uses in plastics and lubricants. Both nonfuel uses and combustion efficiencies have increased over time. However, the two effects counter one another, and, therefore, we are able to keep the fraction oxidized constant.

We use $CO_2$ emissions net of emissions from gas flaring to measure household consumption. This includes emissions generated from the combustion of solid fuel, liquid fuel, gaseous fuel, and cement production. Solid fuel refers to various types of solid material, such as charcoal and coal, used to produce energy. Liquid fuel includes crude petroleum, natural gas liquid, and liquefied petroleum gas (LPG). Gaseous fuel refers to natural gas. We include emissions from cement production to further capture consumption related to housing. Housing-related expenditures are closely related to new construction and thus the usage of cement. In addition, including emissions from cement production also allows us to extend the sample coverage by up to 55 years, during which a separate account for emissions from cement production is not available.[5] Emissions from gas flaring are generated when natural gas is flared at oil fields because of the lack of markets and infrastructure.

Following Campbell (1999a) and Savov (2011), we adopt the standard approach in the C-CAPM literature to compute $CO_2$ emissions growth and match it with the stock return data using the beginning-of-period convention. Specifically, the growth rate of $CO_2$ emissions in year $t$ is calculated using the $CO_2$ emissions from year $t+1$ and $t$ and then matched with the stock returns of year $t$. We adjust emissions by population whenever possible. The population data for the US is from the US Census Bureau. Population estimates are always reported on the first of July each year, so we use the average of the population in year $t$ and year $t-1$ as the population in year $t$ in the calculation of per capita $CO_2$ emissions. Population data for the rest of the world are only available from the World Bank after 1950. Therefore, we replace the per capita emissions with the raw aggregate emissions data in calculating emissions growth for Europe, the world, and other countries, excluding the US, in the pre-1950 sample. In fact, because of the slow-moving nature of population growth, especially in the list of (mostly developed) countries we consider, the application of population

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[5] Whether including emissions from cement production makes little qualitative difference as cement production only accounts for a small fraction of the total $CO_2$ emissions: 1.29% in the US, 2.52% in Europe, and 2.66% in the world. The rolling window correlation coefficients between $CO_2$ emissions growth computed using emissions with and without cement production are always above 99.9% in our sample. Results using $CO_2$ emission excluding cement production are similar and available upon request.

adjustment has little effect on the movement of the computed emissions growth series: we find that the growth series computed using emissions with and without population adjustment has a correlation of over 99%.

To fully benefit from the length and breadth of $CO_2$ emissions data, we obtain stock returns data from multiple sources to cover a wide range of countries and regions over a long sample. We use a country- and region-level stock market index and portfolios to test the assets in our sample. The US stock market index is based on the value-weighted index available from the Center for Research in Security Prices for the period of 1930–2008 extended backward for the 1872–1929 period using data from Robert Shiller's website. For Europe, the return of the stock market index is constructed by merging the Global Financial Database's Developed World Europe Return index from 1907 to 1985 with the MSCI Europe index post-1986 (both measured in USD). We gauge global stock market returns using the World Index from the DMS database, which underpins the Thomson Reuters Credit Suisse Global Investment Returns Yearbook. This index encompasses countries with well-established equity markets, such as Australia, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, Spain, Sweden, Switzerland, the United Kingdom, and the United States. To analyze global/European stock market returns, we employ the global/European $CO_2$ emissions data, which comprises the combined emissions of the countries listed in the respective index. After matching $CO_2$ emissions growth with the stock market returns data, the longest sample for the US, the Europe, and the world is 1872–2015, 1907–2015, and 1907–2015, respectively. For the stock portfolios, we use the Fama-French 25 size and book-to-market portfolio constructed for the US, the Europe, and the world, respectively. These are downloaded from Kenneth French's website, and we subtract the risk-free rate of each region to calculate the excess returns. The sample period is 1929–2015 for the US and 1991–2015 for the Europe and the world. Annual excess stock market returns for a list of fifteen other countries are obtained from the Global Financial Database. This list of countries includes Australia, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, Spain, Sweden, Switzerland, and United Kingdom. Because the $CO_2$ emissions data are always longer than the length of the stock returns data available, the final samples used in this study are defined by the length of the stock returns sample, which varies by country, with the earliest one starting from 1872.

[Insert Table 2 here]

Table 2 presents the summary statistics of the annual per capital growth of total $CO_2$ emissions. The per capita growth of the total $CO_2$ emissions has a sample mean of 1.40%, 1.42%, and 2.49% per year for the US, Europe, and world, respectively, over the full sample. The corresponding standard deviations in the same period are 7.48%, 6.80%, and 5.29%. The growth of total $CO_2$ emissions has a high correlation with the excess return of the market portfolio for the US (42.05%) and the world (39.73%) but less so for the Europe (18.44%). This comovement can be seen in Figure 1, which plots the time series of $CO_2$ emissions growth and the market real excess return for these three regions. $CO_2$ emissions growth clearly comoves with the stock market returns in all three regions. Comovement is particularly evident in the pre-oil-crisis period, where $CO_2$ picks up most of the large movements in the stock market, especially on the downside, and is stronger in the US and the world but less so in Europe. In Panel A, where we also include periods of US recession, we can see that almost all recessions in the US start with a sharp drop in both the growth rate of $CO_2$ and the associated stock market return. These observations support $CO_2$ emissions growth as a reasonable proxy for consumption risk in explaining the cross-sectional and time-series variation in stock returns. We also observe that the growth of $CO_2$ emissions tends to become smoother over the later part of the sample, particularly in the post-oil-crisis sample: the mean emissions growth is much lower for all three regions and so are the standard deviations. Although the correlation with market returns remains high for the US and the world, $CO_2$ emissions growth no longer responds to the large movement in stock market as sensitively as in the earlier sample, indicating that the ability of our measure of $CO_2$ emissions to proxy for consumption can be regionally and sample dependent.[6]

[Insert Figure 1 here]

# 3 Testing the C-CAPM

Despite its profound theoretical influence, the C-CAPM has encountered problems in empirical testing when using the growth rate of the NIPA personal consumption expenditures for nondurables goods and services. Specifically, there are two commonly well-documented puzzles: first, an extremely high level of risk aversion is required to rationalize the observed equity risk premium, and, second, the model-implied risk-free rate is too large relative to its observed value. Under standard model assumptions, these observations can be interpreted as the result of the NIPA personal consumption expenditures growth being too smooth to capture the true risk associated with consumption growth. $CO_2$ emissions are closely related with households' consumption, and the growth of $CO_2$ emissions is more volatile, while correlated with the market returns, so we believe our $CO_2$-emissions-based consumption measure can capture households' underlying consumption risk more adequately. In this section, we empirically investigate whether our $CO_2$-emissions-based consumption growth measure helps to justify the high risk premia observed in the stock market and yields a more reasonable model-implied risk-free rate.

We conduct our tests under the standard C-CAPM assumptions of Lucas (1978) and Breeden (1979). Key assumptions include (1) a two-period model; (2) a complete market; and (3) a power utility function. Under these standard assumptions, the Euler equation that

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[6] The smoother volatility contributes, at least in part, to the diminished efficacy of $CO_2$ when fitting the C-CAPM to post-1970s data, as evident in our later results. This trend of reduced volatility likely mirrors a combination of factors: increasing global environmental consciousness, advancements in energy efficiency, and shifts in household consumption patterns, including changes in consumption composition and the influx of foreign products. We discuss these aspects in more detail in Section 3.1, following the presentation of our results.

prices any asset is expressed as

$$E_t[\beta(\frac{C_{t+1}}{C_t})^{-\gamma} R^e_{t+1}] = 0, \tag{1}$$

where $\beta$ is the subjective discount factor; $C_t$ and $C_{t+1}$ are the representative agent's consumption in period $t$ and $t+1$; $\gamma$ is the coefficient of relative risk aversion in the representative agent's power utility function; and $R^e$ is the excess return of any asset in the market. We fix the subjective discount factor, $\beta$, to be 0.95 following many studies (see, e.g., Hansen and Singleton, 1983; Savov, 2011; Da et al., 2015; Chen and Lu, 2018, among others).[7] Given the observed market excess return and $CO_2$-emissions-based consumption growth, we estimate the coefficient of relative risk aversion $\gamma$ using the generalized method of moments (GMM), where the Euler equation expressed in Equation (1) is used as the moment condition.

## 3.1 US evidence

We estimate the relative risk aversion coefficient using the US market portfolio as the test asset and US per capita $CO_2$ emissions as a proxy for consumption. The baseline results estimated over the full sample of 1872 to 2015 are presented in the first column of Table 3, Panel A. The annual per capita growth rate of $CO_2$ emissions yields a relative risk aversion coefficient of 6.24, which is realistic from an economic perspective. The model-implied risk-free rate is 0.63% per year, in real terms, which is lower than its empirical counterpart over the same sample period. However, it poses less of a puzzle compared with its counterpart implied by the canonical expenditures-based estimate, which we present in later results. These estimated C-CAPM parameters indicate that the US' total $CO_2$ emissions per capita growth can explain the equity premium in the US market portfolio over a long horizon with an RRA and a model-implied risk-free rate at an economically sensible magnitude.

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[7] The model's performance is not qualitatively affected by the choice of $\beta$. The results are available from the authors on request.

Alleviation of the C-CAPM's associated puzzles is effective when we compare the $CO_2$-emission-based estimates with those estimated using the canonical expenditures-based measure, which is calculated as the growth rate of real per capita personal consumption expenditures on nondurable goods and services. In the first two columns of Table 3, Panel B, we compare these estimates over the sample of 1929–2015, which is the longest sample that can be obtained subject to the availability of expenditure-based measures. Using $CO_2$ emissions growth as a proxy for consumption risk produces a lower estimated RRA of 6.75, which is almost half of the RRA of 16.24 produced using the expenditures-based measure. The model-implied risk-free rate using the $CO_2$ measure is negative at -4.18% over this sample; however, in terms of absolute magnitude, it is still more reasonable compared with the 31.72% implied using the expenditures-based measure.

[Insert Table 3 here]

We further analyze the role of the time series variation of $CO_2$ emissions in explaining the equity risk premium puzzles. We do so by estimating and comparing the estimated relative risk aversion coefficient and the model-implied risk-free rate over two subsamples. We use the oil crisis in 1973–1974 to classify our sample into pre-oil-crisis and post-oil-crisis subsamples. We estimate the Euler equation separately in the pre-oil-crisis period of 1872–1973 and the post-oil-crisis period of 1974–2015. We choose the oil crisis as the subsample classification, because the oil crisis was one of the main driving forces that led to global public awareness of energy conservation and improvements in energy efficiency. In addition, the 1970s mark the beginning of decades of significant increases in trade inflow into many countries, including the US. Therefore, the relation between $CO_2$ emissions growth and the true underlying households' consumption risk could vary between our subsamples because of changes in the quantity of goods (and services) being consumed in one country but manufactured (thus $CO_2$ emissions) in other countries. The C-CAPM parameters estimated using $CO_2$ emissions growth for the two subsample periods are presented in columns 2 to 3 in Table 3, Panel A.

The per capita growth rate of $CO_2$ emissions consistently delivers economically reasonable estimates for the relative risk aversion at zero pricing errors: the estimated RRA is 5.27 over the 1872–1973 period and 14.39 for the 1974–2015 period. The model-implied real risk-free rates are 4.32% and -13.14%, respectively, for the two subsamples. It is true that $CO_2$ emissions growth does a better job matching the C-CAPM to the market excess returns in the pre-oil-crisis period than in the post-crisis period.

$CO_2$ measures' poorer performance in capturing household consumption risk in the post-oil-crisis is consistent with the growing global environmental awareness and improvements in energy efficiency that we observe during this period. In addition, some significant changes have occurred in the realm of household consumer goods. On one hand, there has been an increase in the presence of foreign products within the market. The sourcing of these goods from different countries has had a noticeable impact on consumer choices and options. On the other hand, the composition of household consumption baskets has also evolved over time. There has been a shift from a focus on physical goods to a greater emphasis on services-based goods. This transition can be observed as consumers increasingly prioritize services and experiences over the acquisition of tangible products. Due to these reasons, $CO_2$ loses its effectiveness as a reliable measure of the overall consumption service flow over time, especially in the post-oil-crisis period. While these explanations are intuitive and reasonable, gaining a complete understanding of the situation is not possible without access to detailed supply chain data or comprehensive export-import trade data. We acknowledge such limitations and leave this task for future studies.

[Insert Figure 2 here]

Nevertheless, $CO_2$ emissions growth still outperforms the expenditures-based consumption growth measure by far in terms of delivering more sensible C-CAPM parameter estimates. The expenditures-based consumption growth measure gives a very high RRA estimate of 43.62 and an implied risk-free rate of 92.36% in the post-oil-crisis period. The outperformance of the $CO_2$-emissions-based measure is prevalent in all subsamples. Panel A of Figure 2 graphically illustrates this point by plotting the RRAs estimated using $CO_2$ emissions growth and the expenditures-based consumption growth measure over a rolling window of 50 years. We see that the $CO_2$-emissions-based measure consistently yields a RRA of under 15 right up until the early 1990s. The $CO_2$-emissions-based RRA has never exceeded 40, whereas the expenditures-based RRA reaches almost 100 in the same period. Overall, the $CO_2$-emissions-based RRA is always less than half of that estimated using the expenditures-based measure in terms of magnitude.

In addition, we compare our $CO_2$-emissions-based measure with the growth of per capita garbage (municipal solid waste excluding yard trimmings) by Savov (2011) over the post-oil-crisis sample period.[8] The garbage measure yields an RRA of 10.59, which is slightly lower than the RRA produced by our $CO_2$-emissions-based measure. While the garbage measure demonstrates a relatively better fit to the data compared to our $CO_2$-emissions-based measure, both approaches stand out, exhibiting markedly better performance than traditional expenditure-based measures. Furthermore, our $CO_2$-emissions-based measure offers the significant advantage of applicability to a much broader sample, both in terms of time series and cross-section coverage. Therefore, our study makes a valuable contribution by providing a comprehensive examination of the role of consumption risk in explaining the equity risk premium, incorporating a measure that offers a suitable level of fitting along with broader coverage across various dimensions.

## 3.2 Europe and the world evidence

We then consider whether $CO_2$ emissions growth can act as proxy for consumption growth in the C-CAPM framework using the European and world data. Before testing the data,

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[8] The decision to include the garbage measure in our comparison is motivated by its previous utilization within a traditional C-CAPM framework with CRRA utility, unlike the other alternative measures considered (e.g. the electricity usage growth in Da et al. (2015) was instead utilized as a proxy for service flow from household capital in a household production model).

because there is no predominantly clear prior for how well $CO_2$ emissions growth should perform even with knowledge of its outperformance in addressing the equity premium puzzle in the US market, we aggregate the data at the regional and global levels. Doing so comes with benefits and costs. First, analysis at regional and global levels gives us a macro-view of the ability of the $CO_2$ emissions measure to access consumption in the out-of-US setting. Second, aggregation can partially alleviate the effect caused by trade and outsourcing in the recent period. Third, but very importantly, it offers some insights into the performance of C-CAPM over a long and historical sample, which includes periods that the traditional expenditures-based consumption measures do not cover.

On the other hand, tests using world-level or regional-level data require a strong assumption about financial integration across financial markets that does not always hold in reality.[9] In addition, the emissions-based measure is still prone to the impact of fuel efficiency changes and energy conservation concerns in the later periods. We estimate the Euler equation for Europe and the world separately using their $CO_2$ emissions growth and the market excess returns. Table 4, Panel A (Panel B), presents the estimates for Europe (the world) over three sample periods: the full sample, the pre-oil-crisis sample, and the post-oil-crisis sample. The $CO_2$ emissions growth measure delivers a reasonable estimate for the RRA coefficient in the pre-oil-crisis period and in the full sample period, both in Europe and globally. Specifically, the RRA estimated using $CO_2$ emissions growth in the European market is 6.23 over the pre-oil-crisis period and 9.67 over the full sample; the estimated RRA in the world market is 9.07 over the pre-oil-crisis period and 12.22 over the full sample. Similar to the finding using US data, we find that the ability of $CO_2$ emissions growth to explain the European market risk premium and the world market risk premium weakens over time. This can be seen in Figure 3, where we plot the RRA estimated for the European market and the world market using a rolling window of 50 years starting from 1907: the estimated RRA coefficient clearly

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[9] Countries that constitute the World and European indexes are all developed economies, where concerns regarding integration are relatively minor.

increases in the later part of the sample. Taking the post-oil-crisis sample as an example, the estimated RRA in the European and world markets is 24.40 and 48.22, respectively. That being said, $CO_2$ emissions growth still offers better or at least comparable performance relative to the expenditures-based measure. Over the same sample, the RRA required to match expenditures-based consumption growth to the market excess return is double of that estimated using $CO_2$ emissions growth for Europe in terms of magnitude. However, the estimated RRA coefficient for the world is at a similar level (48.22 vs. 47.21).[10]

$CO_2$ emissions growth leads to more sensible estimates for the model-implied risk-free rate in the pre-oil-crisis period but mixed results in the post-oil-crisis period. In the pre-oil-crisis sample, the model-implied risk-free rate is at 5.18% in the European market and 7.31% in the world market. Over the full sample, the implied risk-free rate is at 5.14% in the world market, but it takes a negative value of -6.20% in the European market. The negative value in Europe is mainly driven by negative growth of $CO_2$ emissions post the oil crisis, probably due to more strict emission standard; the implied risk-free rate is -28.28% over that period. The implied risk-free rate also takes a very high level of 57.66% in the world in the post-oil-crisis period. Nevertheless, using the same sample, we find that the implied risk-free rate using the expenditures-based consumption growth is at an enormous level: 109.82% for Europe and 277.07% for the world. Such unreasonable magnitudes of these estimates indicate a failure in fitting the expenditures-based consumption measure to the C-CAPM framework to explain stock returns in Europe and the world. However, despite the mixed results in the post-oil-crisis period, the $CO_2$-emissions-based consumption measure still partially alleviates the joint equity risk premium and implied risk-free rate puzzle better than the traditional expenditures-based measure.

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[10] The standard error of the RRA coefficient estimated using $CO_2$ emissions growth in the post-oil-crisis is "blown up" and thus denoted as "-." This has to do with the choice of using the efficient variance-covariance matrix in the second stage of GMM, a choice that is intended to maximize the asymptotic information in the sample of the model. The downside of using the efficient matrix is that it may blow up standard errors rather than improve pricing errors as explained by Cochrane (1996).

## 3.3 Other countries and regions

There is less analysis on testing the C-CAPM in international stock markets relative to analysis conducted using US data. This is largely because of the lack of data on both stock returns and consumption at the country level. Most of the analyses rely on the country-level stock indices data from the Morgan Stanley Capital International (MSCI), which starts in 1970 for developed countries and 1990 for most of the emerging countries. Consumption data mainly come from the International Financial Statistics (IFS) of the International Monetary Fund covering the period that goes back to at most 1960 for a small selection of countries. Even using the limited data available, the literature has documented some strong evidence of equity premium puzzles in international stock markets.[11] As a representative example, see Campbell (2003), who finds that, using data from 1970 to 1999 for over eleven international markets, the required levels of risk aversion to justify the high equity market risk premia observed in international stock markets are often with magnitudes of over a hundred and even over a thousand for some countries.

We test $CO_2$ emissions growth as a proxy for consumption risk under the C-CAPM framework for a list of fifteen international markets outside US. These countries include Australia, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, Spain, Sweden, Switzerland, and United Kingdom. The $CO_2$-emissions-based proxy benefits us by extending the coverage in both length and breadth compared with other consumption growth proxies. By matching the $CO_2$ emissions growth data with countries' equity market indices data from the Global Financial Database, we are able to estimate the RRA coefficient for these fifteen countries over samples that span on average over 100 years.

[Insert Table 5 Here]

Table 5 displays the estimated RRA coefficients and the corresponding implied risk-free

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[11] This strand of research includes Wheatley (1988), Braun, Constantinides, and Ferson (1993), Chue (2002), Sarkissian (2003), Li and Zhong (2005), and Darrat, Li, and Park (2011).

rates for fifteen countries. These estimates are derived using each country's $CO_2$ emissions growth as a proxy for household consumption growth covering their longest available sample as well as the pre- and post-oil crisis subsamples. The table also includes cross-country averages for these estimates. Across the full sample of all fifteen countries, the average RRA coefficient is approximately 5.33, with the majority of countries presenting an RRA below 10.[12] Similar findings are documented in the pre-oil-crisis sample, with an average estimated RRA coefficient of 6.58. Notably, the $CO_2$-emissions-based consumption measure yields more favorable estimates in the pre-oil-crisis period, with all countries showing positive RRA estimates and the estimated RRAs generally being lower than those in the full sample. Conversely, the post-oil-crisis sample shows higher RRAs, with an average of 21.55, noticeably lower than the 55.81 average yielded by the expenditure-based measure for the same period. The range is from 2.86 in Switzerland to 75.73 in Australia, with the majority of countries exhibiting RRAs between 10 and 20. Denmark, Finland, and France report negative, albeit small, RRA values. These results corroborate our earlier findings using US and regional data, suggesting that the $CO_2$-emissions-based measure is more effective in earlier samples.

Despite the relatively weaker performance of the $CO_2$-emissions-based consumption measure in the post-oil-crisis sample, it still provides better estimates compared to the expenditure-based consumption data in these international markets.[13] We find that, in general, the growth of expenditure-based consumption requires a higher RRA to explain countries' risk premia compared to the growth of $CO_2$ emissions. Except for Canada, where we observed an improvement in RRA, and three other countries (Ireland, Spain, and the UK), where the levels of $CO_2$ are similar, the percentage increase in RRA is significant, exceeding 30%. Furthermore, we observed that the expenditure-based measure produces some unreasonably high levels of implied risk-free rates, ranging from 31.65% to 851.42%.

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[12] This average is calculated from the absolute values of the estimated RRAs for each country. Notably, only Denmark exhibits a negative estimated RRA.

[13] The expenditure-based consumption risk measure, which is proxied by the annual growth of households and NPISHs' final consumption expenditures (obtained from the World Bank), is only available from 1970 onwards.

A couple of issues remain puzzling. Firstly, a few countries (namely, Denmark, Finland, and France) show negative RRA estimates when using the $CO_2$-emission-based measure, especially in the post-oil-crisis period. Secondly, several countries have negative implied real risk-free rates. The low growth in $CO_2$ emissions for these countries is at least partially responsible for the occurrence of negative implied risk-free rates. Such negative rates typically appear in countries and/or during periods characterized by low logarithmic growth of $CO_2$ emissions, especially in the post-oil-crisis period.[14] However, these puzzling results are in general likely due to the limited ability of $CO_2$ emissions data to capture consumption in these countries, which have a well-known concern for greenhouse gas emissions in recent decades. While our current paper may not fully unravel these puzzles, we perceive them as intriguing avenues for future research. They present an opportunity for scholars to delve deeper into understanding the role of alternative consumption risk measurements in explaining asset returns in European countries.[15]

# 4 Cross-sectional pricing power of $CO_2$ growth

In this section, we investigate whether the growth rate of $CO_2$ emissions can serve as a proxy for consumption risk in explaining the observed cross-sectional differences in stock returns. As described in Jagannathan and Wang (2007a), the linearized version of the Euler equation 1 can be approximated as

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[14] The implied risk-free rate can be linearly approximated by the following expression: $R_f = exp(\delta + \gamma E(\Delta c) - 0.5 * \gamma^2 * Var(\Delta c)) = exp(-log(\beta) + \gamma E(\Delta c) - 0.5 * \gamma^2 * Var(\Delta c))$, where $\beta = exp(-\delta)$ and $\frac{c_{t+1}}{c_t} = exp(log(\frac{c_{t+1}}{c_t})) = exp(\Delta c)$. A low log growth of the consumption measure, $E(\Delta c)$, can result in a negative value for the implied risk-free rate.

[15] One possible explanation that some estimation results for European region and countries are less intuitive, especially for the post-oil-crisis period, is that $CO_2$ emissions could also proxy for climate change risk, which has become a major concern for those countries over the past decades. Several recent studies focus on the relation between climate change risk and asset prices, including Litterman (2011), Giglio, Maggiori, Rao, Stroebel, and Weber (2021), Andersson, Bolton, and Samama (2016), Bansal et al. (2017), Bansal et al. (2016), Karp and Rezai (2018), Daniel, Litterman, and Wagner (2019), Krueger, Sautner, and Starks (2020), and Hong, Li, and Xu (2019).

$$E[R^e_{t+1}] = \gamma \beta R^f Cov(\frac{C_{t+1}}{C_t}, R^e_{t+1}). \tag{2}$$

Equation (2) implies that, under the standard assumptions of C-CAPM, the cross-sectional variation in expected excess returns is determined by the correlation between assets' returns and the consumption risk measured by consumption growth. We perform Fama-MacBeth regressions using $CO_2$ emissions growth. Specifically, we first run time-series regressions for test assets' excess returns on $CO_2$ emissions growth to compute assets' corresponding consumption betas. We then estimate the price of consumption risk at each time $t$ by performing a cross-sectional regression of assets' excess returns on the estimated beta loadings. The unconditional market price of consumption risk is computed as the time-series average of the estimated prices of risk. A constant term is included in both stages of regressions to ensure the first-stage $\beta$ estimate is accurate and to allow for an evaluation of the pricing efficiency in the second stage. We are interested in two things: first, whether the consumption risk proxied using $CO_2$ emissions growth is priced in the stock market with a significant price of risk, and, second, whether it has good pricing power reflected in a small constant term in the second-stage regression and a small RMSE.

[Insert Table 6 Here]

Table 6 presents the results from the Fama-MacBeth two-step regressions using 25 US portfolios sorted by size and book-to-market ratio as test assets. That the US stock portfolio data are available from 1929 offers us a long time series of 86 years to conduct the test. We estimate factor risk premia for five different models: (1) a one-factor model with $CO_2$ emissions growth; (2) a one-factor model with the per capita nondurable goods and services expenditures growth; (3) a two-factor model with $CO_2$ emissions growth and expenditures growth; (4) a two-factor model with $CO_2$ emissions growth and the market factor; and (5) a two-factor model with expenditures growth and the market factor. For each model, we report the price of risk for each factor and its *t*-statistics computed based on Newey and West (1987) three-lagged standard errors. An average of the constant terms in the second-stage regression is also presented as a measure of pricing precision. In a one-factor model without controls, $CO_2$-emissions-based consumption growth yields a positive price of risk, which is statistically significant. This indicates that $CO_2$ emissions growth indeed captures consumption risk, which in turn explains the cross-sectional variation in excess returns of the Fama-French 25 portfolios formed using US stocks. The constant term is small and statistically insignificant. Using the $CO_2$-emissions-based consumption growth measure, the average pricing error for the Fama-French 25 portfolios is -3% per year with a RMSE of 2.21% per year. The NIPA nondurable goods and services expenditures-based measure, on the other hand, has a weak pricing power, because its estimated price of risk is not statistically significant when used in a single-factor model alone or with controls. In addition, estimates for the constant term are statistically nonzero. In general, the $CO_2$ emissions measure yields a lower pricing error compared to the expenditures-based measure, irrespective of whether the market factor is taken into account. The pricing power of this emission-based consumption measure remains significant even after controlling for the market risk factor or the expenditures-based consumption growth measure.

[Insert Table 7 Here]

We then assess whether $CO_2$ emissions growth has cross-sectional pricing power in the international markets. Because of the lack of and/or low-quality portfolio-/stock-level data for international markets in earlier periods, we implement the standard Fama-MacBeth procedure on the twenty-five global portfolios and twenty-five European portfolios from Kenneth French's data library with a sample period of 1991 to 2015. These portfolios are constructed by sorting individual stocks in that market by size and book-to-market ratio. Table 7 presents the second-stage price of risk for both $CO_2$ emissions growth and the expenditures-based consumption growth measures. We find that $CO_2$ emissions growth delivers a positive price of risk in pricing both the European portfolios and the world portfolios. This holds even if we control for the expenditures-based consumption growth and/or the market excess return. In contrast, the expenditures-based consumption growth measure can sometimes yield a negative price of risk when it is used as the single factor in explaining the 25 world portfolios or when it is used together with $CO_2$ emissions growth. In addition, tests using $CO_2$ emissions growth always yield higher adjusted $R^2$ than tests using expenditures-based consumption growth in a one-factor setting and/or a two factors setting controlling for the market factor. However, the statistical significance of the price of risk on both $CO_2$ emissions growth and expenditures-based consumption growth are not significant in all cases. The weak pricing power does not come with too much of a surprise as tests are performed over a very short sample of 25 years of annual data due to its availability and, more importantly, as shown in earlier sections, $CO_2$ emissions growth performs less effectively in the recent period in terms of capturing the consumption risks in Europe and the world. Thus, we would expect the cross-sectional pricing power to improve in earlier sample and when longer data are available.

The $CO_2$ betas are reported in Figure 4. Note that, all $CO_2$ betas are positive across the size and book-to-market sorted portfolios in the US, Europe, and world markets. Specifically, betas of the US portfolios range between 1.49 and 3.06 with a minimum $t$-statistic of 2.75. Small stocks and high book-to-market stocks tend to have higher exposures to the $CO_2$ emission factor. These findings echo with earlier studies of Bansal, Dittmar, and Lundblad (2005) and Hansen, Heaton, and Li (2008) that argue high book-to-market stocks have higher sensitivities to long-run consumption growth risk. We find similar patterns for risk exposures of the global and European portfolios.

# 5 Additional Discussion

One potential limitation of our $CO_2$ consumption measure is that a portion of $CO_2$ emissions may arise from industrial production and/or physical investments. Although, as shown earlier, studies in ecology and energy economics provide evidence of household consumption being the main driver behind $CO_2$ emissions, it remains a concern that the pricing power of the $CO_2$ factor may come from its correlation with industrial production or investment instead of household consumption.

We show that it is not the case. Due to data availability, we restrict this set of analyses to the US market. We proxy industrial production using the seasonally adjusted industrial production index published by the Federal Reserve Board, which measures the level of production and capacity in the manufacturing, mining, electric, and gas industries, relative to a base year. Specifically, we use the Fama-MacBeth two-pass regressions to test: 1) whether $CO_2$ emissions growth still prices the US stock portfolios after controlling for industrial production growth; and, 2) whether the residual from $CO_2$ emissions growth orthogonalized w.r.t. industrial production growth is a priced factor.

Table 8 presents our results, which indicate that $CO_2$ emissions continue to price US stock portfolios even after controlling for industrial production growth. Furthermore, we find that the residual from $CO_2$ emissions orthogonalized w.r.t. industrial production growth is a priced factor. We then repeat the same set of tests controlling for industrial production from electric and gas utilities, which are more directly related to $CO_2$ emissions, and the results are similar. Next, we use private non-residential fixed investment to control for physical investment in factories and machines. We find that the pricing power of $CO_2$ emission growth remains robust. Results from these sets of analyses indicate that it is more likely that the pricing power of $CO_2$ emissions comes from the variation in household consumption activities rather than that in production or physical investment.

[Insert Table 8 Here]

Another possible concern is that the pricing power of $CO_2$ emissions may be subject to the confounding effect of climate change risk: $CO_2$ emissions could lead to rising temperature in the long run; and, Bansal et al. (2017) show that long-run temperature fluctuations carry a positive risk premium in equity market due to their impact on the aggregate economy. While we acknowledge that higher $CO_2$ emissions may be related to global warming and thus the resulting climate change risk in the long run, our paper focuses on the impact of relatively short-run variations in consumption, which is proxied by $CO_2$ emissions, on asset prices.

We empirically demonstrate that our $CO_2$ emission factor differs from the temperature-based climate risk. We reconstruct the long-run climate risk measure proposed by their paper, i.e. the five-year difference in US average temperature, as well as temperature risk over shorter horizons, i.e. i.e. three-year and one-year. The temperature-based long-run climate risk measure has a low correlation coefficient of 6.93% with our $CO_2$ growth measure; and, the numbers are 11.92% and -13.90% for the three-year and one-year temperature risk factors. Moreover, Table 9 shows that both the short-run and the long-run temperature risk measures exhibit different pricing pattern compared to the ones associated with our $CO_2$ factor.In contrast to the positive price of risk for the $CO_2$ growth factor, the temperature-based climate risk measure exhibits a negative price of risk that is consistent with the story in Bansal et al. (2017). More importantly, we show that the significant pricing power of $CO_2$ growth remains unaffected after controlling for the temperature risk. All the findings suggest that the proposed $CO_2$ consumption risk measure is unlikely to be driven by the climate change risk.

[Insert Table 9 Here]

# 6 Conclusion

In this paper, we use $CO_2$ emissions as a proxy for household consumption in testing the C-CAPM with CRRA utility function for the US and fifteen other international markets over 100 years. A broad range of household consumption involves emissions of $CO_2$. Our measure also has favorable features: higher correlation with stock market returns and being more volatile than the canonical expenditures-based consumption measures.

Our empirical results deliver a number of interesting findings. Using the annual growth rate of $CO_2$ emissions as a proxy for the consumption risk, our estimation achieves a very reasonable value for the relative risk aversion coefficient of around 6 and an implied real risk-free rate of 0.63% over the full sample of 1872–2015 in the US. The $CO_2$-emissions-based measure also helps resolve the equity risk premium in international markets: an estimated RRA is 10 in Europe, 12 in the world, and 5 on average in fifteen other international markets. $CO_2$-emissions-based consumption growth also explains the cross section of stock returns. Lastly, the pricing power of our $CO_2$ emission factor is persistent over time, although we do observe a better performance in the pre-oil-crisis period relative to the post-oil-crisis period.

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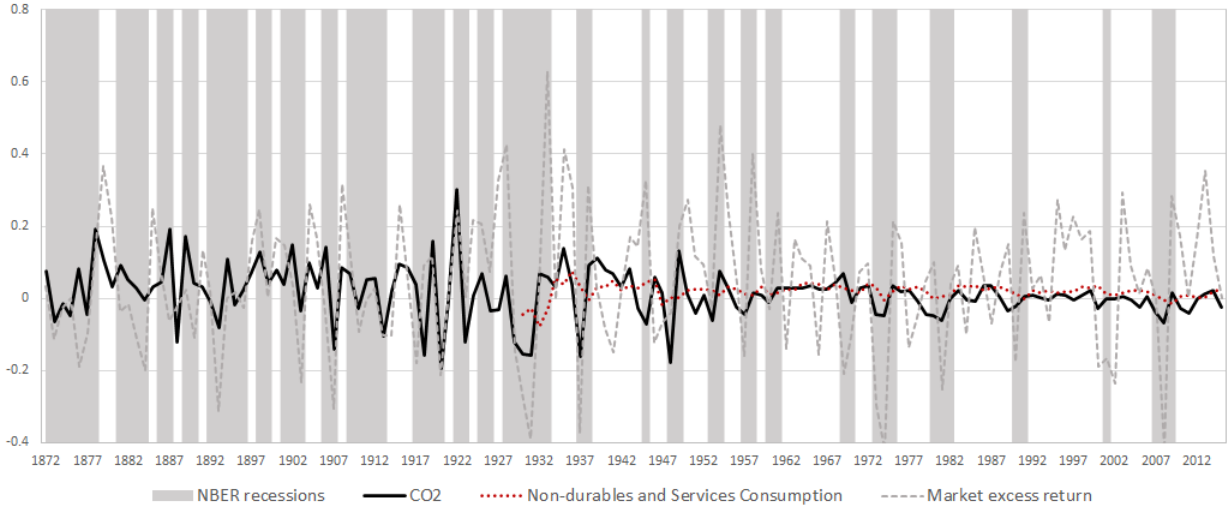
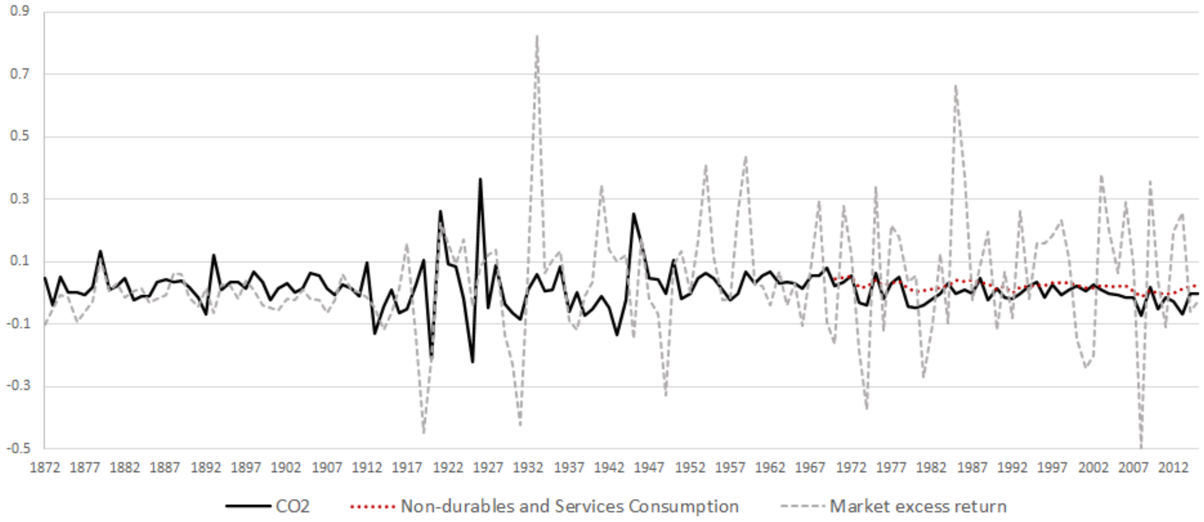
Figure 1: $CO_2$ emissions growth and expenditures growth

This figure compares the time series of the annual $CO_2$ emissions growth to the consumption expenditures growth. Panel A presents the US time series: the solid line represents the annual growth rate of the per capita $CO_2$ emissions; the red dotted line represents the annual per capita growth of nondurables goods and services expenditures from NIPA; the gray dashed line represents the annual real excess returns of the US stock market; and the shaded bands indicate National Bureau of Economic Research (NBER) recessions. Panels B and C present the time series for Europe and the world. Within each figure, the solid line, the dotted line, and the dashed line represent $CO_2$ emissions growth, the households and NPISHs Final consumption expenditures growth, and the excess returns of the corresponding region. Growth rates and returns are demeaned and expressed as percentages. The sample period for $CO_2$ emissions growth and stock returns are 1872–2015 for the US and Europe and 1907–2015 for the world. The sample period for consumption growth is 1929–2015 for the US and 1970–2015 for Europe and the world.

(a) US time series



(b) Europe time series



(c) World time series

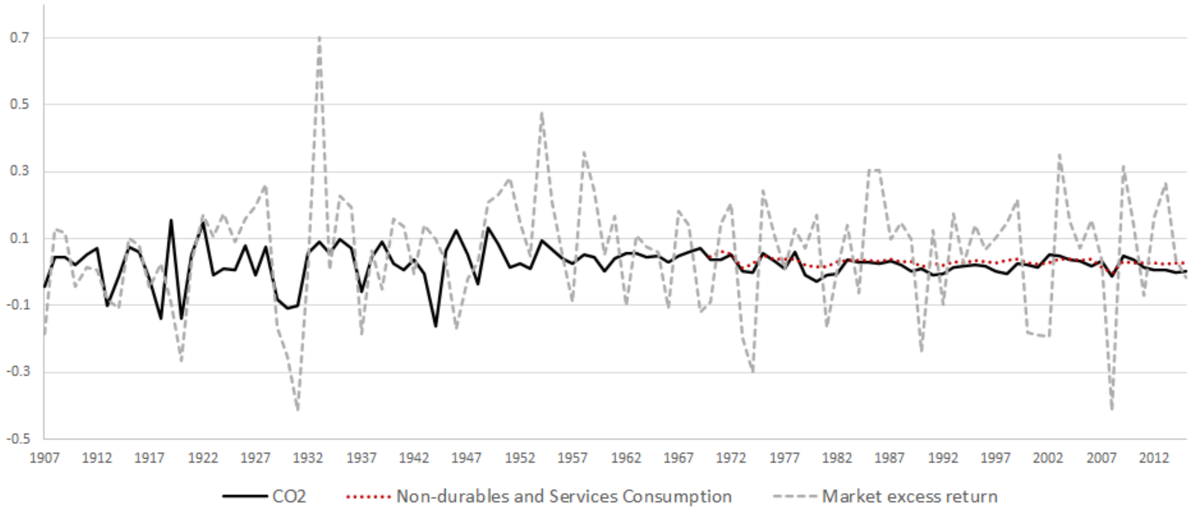


Figure 2: Time-varying RRA estimated using $CO_2$ emissions growth: the US

This figure plots the time series of the relative risk aversion (RRA) coefficients estimated for the US using the growth rate of $CO_2$ emissions as a proxy of consumption growth over a rolling window of 50 years. Specifically, the RRA in year $t$ is estimated using data from year $t$-49 to $t$. The estimated RRAs are represented by the solid line. We also plot the RRAs estimated using nondurable goods and services expenditures growth on the dotted line for comparison purposes. The estimates start in 1921 with an estimation window of 1872–1921 using the $CO_2$ emissions measure, and they start in 1978 for the expenditures-based measure. All estimates end in 2015.

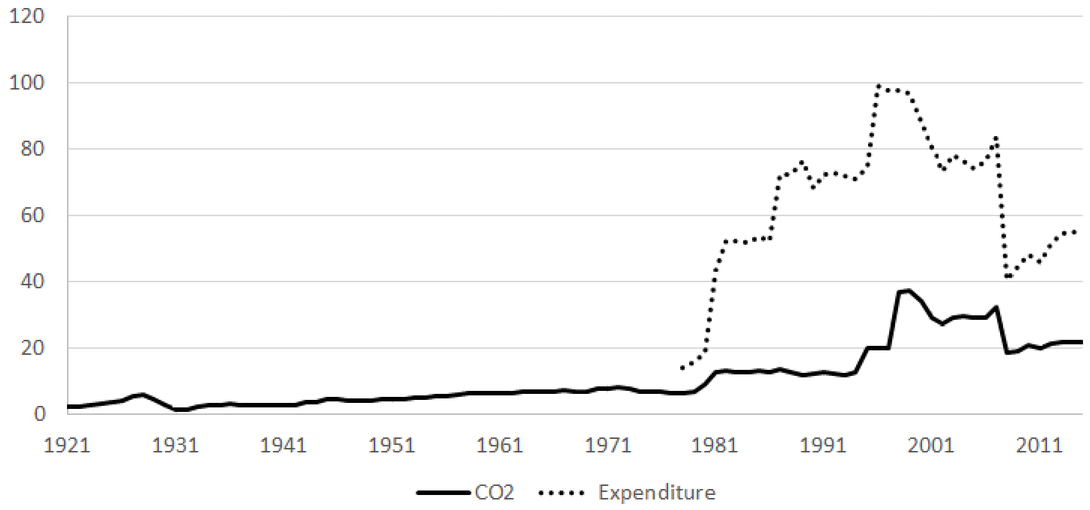


Figure 3: Time-varying RRA estimated using $CO_2$ emissions growth: Europe and the world

This figure plots the time series of the relative risk aversion (RRA) coefficients estimated in Europe and the world using the growth rate of $CO_2$ emissions as a proxy of consumption growth over a rolling window of 50 years. Specifically, the RRA in year $t$ is estimated using data from year $t$-49 to $t$. RRAs estimated in Europe and the world are represented by the dashed and dotted lines, respectively. The estimates start in 1956 with an estimation window of 1907–1956 and end in 2015.

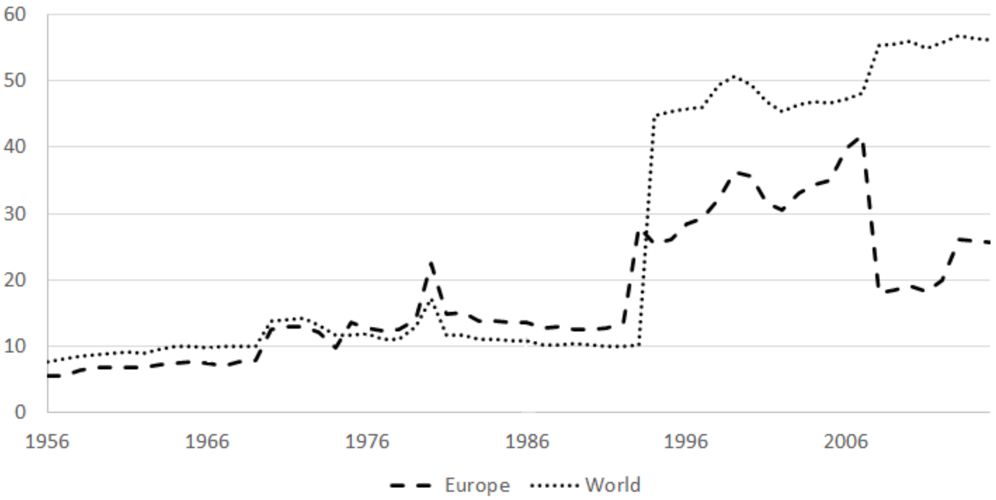
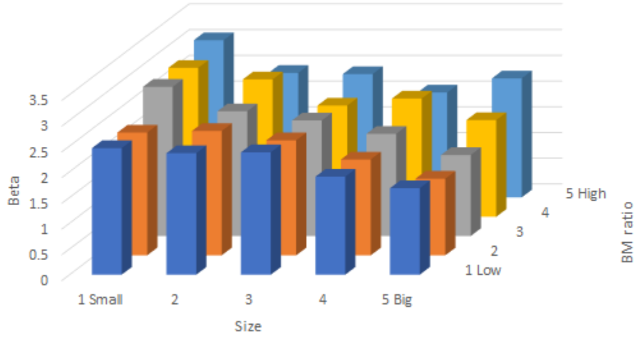


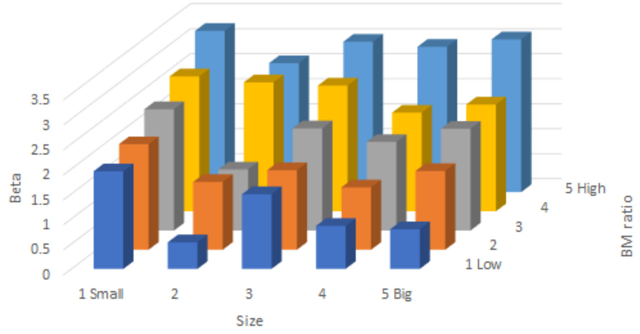
Figure 4: Risk exposures to $CO_2$ emissions growth

This figure plots portfolios' risk exposures to the $CO_2$ emissions growth from the first stage of the Fama-MacBeth two-pass regressions that estimate a linear one-factor model using twenty-five portfolios sorted by size and book-to-market ratio as test assets. Panel (a)/(b)/(c) presents results using the US/European/global stock portfolios respectively. The sample period is 1927–2015 for the US portfolios and 1991–2015 for European and global portfolios.

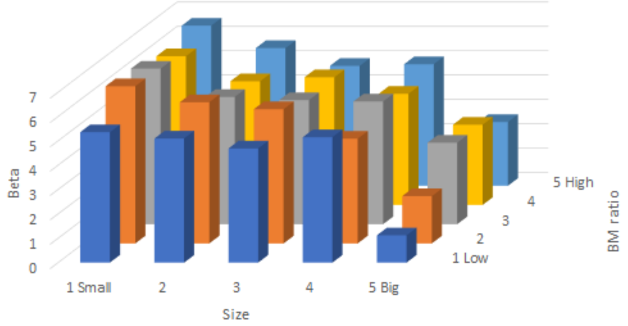
(a) Risk exposures of US portfolios



(b) Risk exposures of European portfolios



(c) Risk exposures of global portfolios



# Table 1: Sample coverage comparison

This table compares the sample period used in this paper to sample periods used in existing papers that estimate C-CAPM using consumption expenditure data or alternative consumption measures.

| Country | Our sample period | Sample periods from existing papers |
|---|---|---|
| Australia | 1872 – 2015 | 1970 – 2007 (Li, 2010b), 1974 – 1992 (Faff, 1998), 1974 – 2006 (Li, 2010a), |
| Belgium | 1897 – 2015 | |
| Canada | 1872 – 2015 | 1970 – 1988 (Braun et al., 1993) |
| Denmark | 1874 – 2015 | 1922 – 1990 (Lund and Engsted, 1996) |
| Finland | 1913 – 2015 | 1970 – 2007 (Li, 2010b), 1990 – 2009 (Virk, 2012) |
| France | 1872 – 2015 | 1970 – 1988 (Braun et al., 1993), 1970 – 1988 (Braun et al., 1993), 1970 – 2007 (Li, 2010b) |
| Germany | 1872 – 2015 | 1885 – 1913 and 1951 – 1990 (Lund and Engsted, 1996), 1970 – 2007 (Li, 2010b) |
| Ireland | 1935 – 2015 | |
| Italy | 1925 – 2015 | 1970 – 2007 (Li, 2010b), 1970 – 2007 (Li, 2010b) |
| Japan | 1886 – 2015 | 1970 – 1988 (Braun et al., 1993), 1980 – 1988 (Hamori, 1992) |
| Netherlands | 1951 – 2015 | |
| Sweden | 1872 – 2015 | 1918 – 1990 (Lund and Engsted, 1996), 1970 – 2007 (Li, 2010b) |
| Spain | 1941 – 2015 | 1970 – 2007 (Li, 2010b) |
| Switzerland | 1914 – 2015 | 1970 – 2007 (Li, 2010b) |
| UK | 1872 – 2015 | 1919 – 1987 (Lund and Engsted, 1996) |
| US | 1872 – 2015 | 1939 – 1982 (Breeden, Gibbons, and Litzenberger, 1989), 1947 – 1998 (Campbell, 2003b), 1950 – 2009 (Liu, Luo, and Zhao, 2016), 1951 – 2001 (Yogo, 2006), 1954 – 2003 (Jagannathan and Wang, 2007b), 1954 – 2003 (Parker and Julliard, 2005b), 1955 – 2012 (Da et al., 2015), 1960 – 2007 (Savov, 2011), 1963 – 1998 (Lettau and Ludvigson, 2001), 1969 – 2000 (Li and Zhong, 2005), 1970 – 1988 (Braun et al., 1993), 1970 – 2010 (Chen and Lu, 2018), 1982 – 2004 (Malloy, Moskowitz, and Vissing-Jørgensen, 2009b), 1986 – 2000 (Ait-Sahalia, Parker, and Yogo, 2004) |

Table 2: Summary statistics for the $CO_2$ emissions growth measure

This table presents summary statistics for $CO_2$ emissions growth. Statistics include the mean, standard deviation, and AR(1) autocorrelation coefficient of $CO_2$ emissions growth for the US, Europe, and the world and their correlations with the corresponding stock market returns $R^M$. We present the summary statistics for three sample periods: the full sample of 1872–2015, the pre-oil-crisis sample of 1872–1973, and the post-oil-crisis period of 1974–2015, with the exception for the world, where its first observation starts in 1907. All statistics are expressed as percentages.

| | Full sample | | | Pre-oil-crisis | | | Post-oil-crisis | | |
|---|---|---|---|---|---|---|---|---|---|
| | US | Europe | World | US | Europe | World | US | Europe | World |
| Mean | 1.40 | 1.42 | 2.49 | 2.29 | 2.15 | 2.90 | -0.76 | -0.35 | 1.85 |
| SD | 7.48 | 6.80 | 5.29 | 8.58 | 7.73 | 6.53 | 2.65 | 3.05 | 2.09 |
| AR(1) coeff. | -14.52 | -10.68 | 8.94 | -20.52 | -15.26 | -3.45 | 22.71 | 3.55 | 27.81 |
| Corr. with $R^M$ | 42.05 | 18.44 | 39.73 | 44.66 | 17.91 | 45.24 | 49.44 | 28.75 | 34.89 |

Table 3: Relative risk aversion estimation: Evidence from the US

This table presents the C-CAPM parameters estimated using $CO_2$ emissions growth and market real excess returns in the US. $CO_2$ emissions growth acts as a proxy for consumption risk in the C-CAPM. Estimates are obtained by estimating the following Euler equation using the GMM:

$$E[\beta(\frac{C_{t+1}}{C_t})^{-\gamma}R^e_{t+1}] = 0.$$

The subject discount factor, $\beta$, is set to be 0.95. The relative risk aversion (RRA) coefficient, $\gamma$, is presented with Newey-West three-lagged adjusted GMM standard errors displayed in parentheses. The model-implied risk-free rates ($R^f$) are computed based on the estimated RRA and expressed as percentages. Pricing errors are defined as $\sqrt{g'_T g_T/N}$, where $N$ is the number of test assets. Panel A presents the estimates using $CO_2$ emissions growth. Panel B compares the estimates obtained using $CO_2$ emissions growth and ones obtained using nondurable goods and services (ND&S) expenditures growth and the growth of garbage (1974 – 2015). Estimates are presented for three sample periods: the full sample (1872–2015 in Panel A and 1929–2015 in Panel B), the pre-oil-crisis period (1872–1973 in Panel A and 1929–2015 in Panel B), and the post-oil-crisis period (1974–2015).

A. Estimates using the $CO_2$ emissions growth

| | Full sample<br>1872–2015 | Pre-oil-crisis<br>1872–1973 | Post-oil-crisis<br>1974–2015 |
|---|---|---|---|
| RRA($\gamma$) | 6.24 | 5.27 | 14.39 |
| (SE) | (2.22) | (2.12) | (9.74) |
| Implied $R^f$ (%) | 0.63 | 4.32 | -13.14 |
| Pricing error | 0.0000 | 0.0000 | 0.0000 |

B: Comparison between estimates using different consumption growth proxies

| | Full sample | | Pre-oil-crisis | | Post-oil-crisis | | |
|---|---|---|---|---|---|---|---|
| | 1929–2015 | | 1929–1973 | | 1974–2015 | | |
| | $CO_2$ | Expenditures | $CO_2$ | Expenditures | $CO_2$ | Expenditures | Garbage |
| RRA($\gamma$) | 6.75 | 16.24 | 5.91 | 11.99 | 14.39 | 43.62 | 10.59 |
| (SE) | (3.30) | (8.13) | (3.30) | (7.76) | (9.74) | (27.47) | (7.46) |
| Implied $R^f$ (%) | -4.18 | 31.72 | -0.61 | 24.77 | -13.14 | 92.36 | 8.67 |
| Pricing error | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.000 |

Table 4: Relative risk aversion estimation: Evidence from Europe and the world

This table presents the C-CAPM parameters estimated using $CO_2$ emissions growth and market real excess returns in Europe and the world. $CO_2$ emissions growth acts as a proxy for consumption risk in the C-CAPM. Estimates are obtained by estimating the following Euler equation using the GMM:

$$E[\beta(\frac{C_{t+1}}{C_t})^{-\gamma}R^e_{t+1}] = 0.$$

The subject discount factor, $\beta$, is set to be 0.95. The relative risk aversion (RRA) coefficient, $\gamma$, is presented with Newey-West three-lagged adjusted GMM standard errors displayed in parentheses. The model-implied risk-free rates ($R^f$) are computed based on the estimated RRA and expressed as percentages. Pricing errors are defined as $\sqrt{g'_T g_T/N}$, where $N$ is the number of test assets. Panel A presents results for Europe, and Panel B presents results for the world. The test asset for these two markets is the real market excess returns, obtained from Thomson Reuters Credit Suisse Yearbook and the Global Financial Database. We present estimates for three sample periods: the full sample (1907–2015), the pre-oil-crisis period (1907–1973) and the post-oil-crisis period (1974–2015). For comparison purposes, we also present the estimates obtained using households' final consumption expenditures growth as a proxy for consumption growth for the post-oil-crisis period.

A. The European market

| | Full sample 1907–2015 | Pre-oil-crisis 1907–1973 | Post-oil-crisis 1974–2015 | |
|---|---|---|---|---|
| | $CO_2$ | $CO_2$ | $CO_2$ | Expenditures |
| RRA($\gamma$) | 9.67 | 6.23 | 24.40 | 42.87 |
| (SE) | (3.83) | (4.27) | (17.44) | (21.77) |
| Implied $R^f$ (%) | -6.20 | 5.18 | -28.28 | 109.82 |
| Pricing error | 0.0000 | 0.0000 | 0.0000 | 0.0000 |

B. The world market

| | Full sample 1907–2015 | Pre-oil-crisis 1907–1973 | Post-oil-crisis 1974–2015 | |
|---|---|---|---|---|
| | $CO_2$ | $CO_2$ | $CO_2$ | Expenditures |
| RRA($\gamma$) | 12.22 | 9.07 | 48.22 | 47.21 |
| (SE) | (5.28) | (4.68) | - | (22.63) |
| Implied $R^f$ (%) | 5.14 | 7.31 | 57.66 | 277.07 |
| Pricing error | 0.0000 | 0.0000 | 0.0000 | 0.0000 |

Table 5: Relative risk aversion estimation: International markets

This table reports the estimated relative risk aversion (RRA) coefficient, $\gamma$, and implied risk-free rate, $R^f$, for fifteen international markets using GMM estimation. The moment condition for each country is

$$E[\beta(\frac{C^i_{t+1}}{C^i_t})^{-\gamma}R^{e,i}_{t+1}] = 0,$$

where $\frac{C^i_{t+1}}{C^i_t}$ is the growth rate consumption in country $i$, and $R^{e,i}_{t+1}$ is country $i$'s stock market excess return. The subject discount factor, $\beta$, is set to be 0.95. The relative risk aversion (RRA) coefficient, $\gamma$, is presented with Newey-West three-lagged adjusted GMM standard errors displayed in parentheses. The model-implied risk-free rates ($R^f$) are computed based on the estimated RRA and expressed as percentages. The growth rate consumption in country $i$ is proxied by country $i$'s $CO_2$ emissions growth ($CO_2$), and the annual growth of households and NPISHs final consumption expenditures (Expenditures). The parameters for each country are estimated using each country's longest available sample, the pre-oil-crisis sample (before 1974), and the post-oil-crisis sample (1974–2015). The average relative risk aversion coefficient and the model-implied risk-free rates across the country-level estimates are also reported.

| | Available | $CO_2$ Full sample | | $CO_2$ Pre-oil-crisis | | $CO_2$ Post-oil-crisis | | Expenditures Post-oil-crisis | |
|---|---|---|---|---|---|---|---|---|---|
| Country | sample | RRA | Implied $R^f$ | RRA | Implied $R^f$ | RRA | Implied $R^f$ | RRA | Implied $R^f$ |
| Australia | 1872–2015 | 10.04 | 13.90 | 8.91 | 15.49 | 75.73 | -65.43 | 101.49 | 851.42 |
| | | (-) | | (-) | | (136.81) | | (73.69) | |
| Belgium | 1897–2015 | 5.86 | -26.07 | 4.12 | -11.60 | 16.12 | -34.92 | 192.31 | 267.41 |
| | | (1.90) | | (2.41) | | (8.98) | | (206.81) | |
| Canada | 1872–2015 | 10.70 | 0.10 | 8.92 | 6.80 | 48.00 | -43.91 | 30.47 | 112.24 |
| | | (4.26) | | (4.2) | | (-) | | (11.88) | |
| Denmark | 1874–2015 | -0.83 | 2.79 | 2.77 | 11.17 | -8.11 | -16.71 | 37.95 | 37.27 |
| | | (-) | | (-) | | (-) | | (16.85) | |
| Finland | 1913–2015 | 2.24 | -45.76 | 1.84 | -39.49 | -0.59 | 5.19 | 24.03 | 51.16 |
| | | (0.48) | | (0.46) | | (-) | | (11.1) | |
| France | 1872–2015 | 9.61 | 3.60 | 8.65 | 1.85 | -0.34 | 5.48 | 91.97 | 273.22 |
| | | (3.12) | | (3.87) | | (-) | | (52.74) | |
| Germany | 1872–2015 | 3.48 | -65.37 | 3.46 | -72.23 | 31.83 | -54.90 | 71.13 | 116.51 |
| | | (-) | | (-) | | (17.2) | | (34.59) | |
| Ireland | 1935–2015 | 8.01 | -7.46 | 1.70 | 7.39 | 19.89 | -11.88 | 17.48 | 50.01 |
| | | (-) | | (-) | | (10.4) | | (8.13) | |
| Italy | 1925–2015 | 0.87 | 3.79 | 0.82 | 3.23 | 16.21 | -15.54 | 21.12 | 32.99 |
| | | (-) | | (-) | | (12.94) | | (15.69) | |
| Japan | 1886–2015 | 5.86 | -5.36 | 5.09 | 1.81 | 15.42 | 0.27 | 39.98 | 89.90 |
| | | (2.79) | | (2.67) | | (12.19) | | (38.76) | |
| Netherlands | 1951–2015 | 14.10 | -10.79 | 44.64 | 13.31 | 12.89 | -23.11 | 55.94 | 53.09 |
| | | (4.81) | | (23.75) | | (5.12) | | (25.15) | |
| Spain | 1941–2015 | 0.17 | 5.45 | 0.06 | 5.59 | 15.68 | -46.85 | 15.69 | 31.65 |
| | | (-) | | (-) | | (11.76) | | (12.18) | |
| Sweden | 1872–2015 | 0.97 | 7.37 | 0.63 | 7.34 | 21.69 | -42.28 | 44.29 | 58.77 |
| | | (-) | | (-) | | (11.45) | | (15.09) | |
| Switzerland | 1914–2015 | 3.01 | -14.90 | 3.01 | -24.16 | 2.86 | 3.10 | 55.74 | 91.51 |
| | | (-) | | (-) | | (-) | | (20.43) | |
| United Kingdom | 1872–2015 | 4.20 | -3.87 | 4.12 | -4.48 | 37.88 | -81.09 | 37.5 | 88.49 |
| | | (-) | | (-) | | (45.12) | | (18.01) | |
| Average | | 5.33 | -9.51 | 6.58 | -5.20 | 21.55 | -26.77 | 55.81 | 147.04 |

Table 6: Cross-sectional pricing of US stock portfolios

This table reports results from the Fama-MacBeth two-pass regressions of the linear factor models with twenty-five US portfolios sorted by size and book-to-market ratio as test assets. A cross-sectional constant is included in the estimation. $CO_2$ is annual $CO_2$ emissions per capita growth. Expenditures are the annual growth of seasonally adjusted per capita expenditures on nondurable goods and services (ND&S) from NIPA. Market is the market excess return. We estimate factor risk premia for five different models: (1) a one-factor model with $CO_2$ emissions growth; (2) a one-factor model with ND&S growth; (3) a two-factor model with $CO_2$ emissions growth and ND&S growth; (4) a two-factor model with $CO_2$ emissions growth and the market factor; and (5) a two-factor model with ND&S growth and the market factor. A constant is included in the second-stage regression. Regression coefficients (factor risk premia) are reported, with $t$-statistics adjusted using Newey–West three-lagged corrections in parentheses. Root-mean-square errors (RMSEs) and adjusted $R^2$ are measured as percentages. The sample period is 1929–2015.

| | $CO_2$ | Expenditures | Market | Constant | RMSE | Adj. $R^2$ |
|---|---|---|---|---|---|---|
| (1) | 5.023 | | | -0.25 | 2.21 | 19.86 |
| | (2.21) | | | (-0.07) | | |
| (2) | | 0.94 | | 3.52 | 2.53 | 22.69 |
| | | (1.65) | | (1.10) | | |
| (3) | 5.90 | 1.05 | | -0.34 | 2.18 | 25.45 |
| | (3.35) | (1.80) | | (-0.09) | | |
| (4) | 6.77 | | 6.29 | 2.10 | 2.10 | 27.88 |
| | (3.65) | | (1.28) | (0.46) | | |
| (5) | | 1.55 | 0.63 | 8.57 | 2.40 | 28.93 |
| | | (2.66) | (0.14) | (2.06) | | |

Table 7: Cross-sectional pricing of European and world stock portfolios

This table reports results from the Fama-MacBeth two-pass regressions of the linear factor models with twenty-five size and book-to-market portfolios constructed using European stocks and stocks from developed markets as test assets. $CO_2$ is the annual $CO_2$ emissions growth. Expenditures are the annual growth of households and NPISHs final consumption expenditures from the World Bank. Market is the market excess return. We estimate factor risk premia for five different models: (1) a one-factor model with $CO_2$ emissions growth; (2) a one-factor model with ND&S growth; (3) a two-factor model with $CO_2$ emissions growth and ND&S growth; (4) a two-factor model with $CO_2$ emissions growth and the market factor; and (5) a two-factor model with ND&S growth and the market factor. A constant is included in the second-stage regression. Regression coefficients (factor risk premia) are reported, with $t$-statistics adjusted using Newey–West three-lagged corrections in parentheses. Root-mean-square errors (RMSEs) and adjusted $R^2$ are expressed as percentages. We present results for European stock portfolios in Panel A and results for world portfolios in Panel B. The sample period is 1991–2015.

A. European portfolios

| | $CO_2$ | Expenditures | Market | Constant | RMSE | Adj. $R^2$ |
|---|---|---|---|---|---|---|
| (1) | 2.43 | | | 4.18 | 1.97 | 19.48 |
| | (1.62) | | | (0.77) | | |
| (2) | | 0.82 | | 2.88 | 2.37 | 13.26 |
| | | (1.35) | | (0.47) | | |
| (3) | 2.64 | 0.30 | | 7.85 | 1.86 | 22.83 |
| | (1.68) | (0.65) | | (1.56) | | |
| (4) | 2.36 | | -0.51 | 8.89 | 1.91 | 30.80 |
| | (1.62) | | (-0.10) | (2.88) | | |
| (5) | | 0.80 | 0.19 | 9.85 | 2.24 | 25.37 |
| | | (1.35) | (0.04) | (3.02) | | |

B. World portfolios

| | $CO_2$ | Expenditures | Market | Constant | RMSE | Adj. $R^2$ |
|---|---|---|---|---|---|---|
| (1) | 0.64 | | | 4.40 | 1.93 | 20.21 |
| | (0.94) | | | (1.18) | | |
| (2) | | -0.06 | | 8.38 | 2.10 | 19.90 |
| | | (-0.16) | | (1.95) | | |
| (3) | 0.66 | -0.15 | | 8.04 | 1.80 | 41.15 |
| | (0.95) | ( -0.39) | | (1.89) | | |
| (4) | 0.76 | | -3.07 | 9.23 | 1.58 | 50.10 |
| | (1.07) | | (-0.60) | (2.22) | | |
| (5) | | 0.74 | 0.68 | 6.51 | 1.50 | 40.55 |
| | | (1.64) | (0.14) | (1.68) | | |

Table 8: The effect of industrial production and investment

This table reports the Fama-MacBeth two-pass regressions of linear factors models that assess the the cross-sectional pricing power of $CO_2$ emissions growth controlling for the effect of industrial production and private non-residential fixed investment. Test assets are the twenty-five US portfolios sorted by size and book-to-market ratio. $CO_2$ is annual $CO_2$ emissions per capita growth. Control variables include growth of industrial production index, growth of industrial electric and gas utilities, and growth of private non-residential fixed investment. For each control variable, we estimate factor risk premia for a two-factor model including $CO_2$ emissions growth and the control variable (Regression (1)/(3)/(5)), and we estimate a one-factor model that includes residuals from regressing $CO_2$ emissions growth on the control variable (Regression (2)/(4)/(6)); and, we repeat the estimations of the two models in Regression (7) and (8) by controlling for all three variables at once. A cross-sectional constant is included in the estimation. Factor risk premia are reported with $t$-statistics adjusted using Newey–West three-lagged corrections in parentheses. Root-mean-square errors (RMSEs) and adjusted $R^2$ are measured as percentages. The sample period is 1929–2015 for regression (1), (2), (5), (6), and 1939–2015 for regressions (3), (4), (7) and (8).

| | Control variable | $CO_2$ | Control | $CO_2^{\perp}$ *control* | Constant | RMSE | Adj. $R^2$ | Sample |
|---|---|---|---|---|---|---|---|---|
| (1) | Industrial production | 6.59 | 1.31 | | 4.76 | 2.03 | 14.32 | 1929 – 2015 |
| | | (2.41) | (0.65) | | (2.12) | | | |
| (2) | Industrial production | | | 4.62 | -0.66 | 1.99 | 20.40 | 1929 – 2015 |
| | | | | (2.52) | (-0.20) | | | |
| (3) | Electric and gas utilities | 8.13 | -2.17 | | 2.68 | 1.58 | 26.40 | 1939 – 2015 |
| | | (2.66) | (-0.86) | | (1.04) | | | |
| (4) | Electric and gas utilities | | | 9.08 | 1.62 | 1.66 | 12.59 | 1939 – 2015 |
| | | | | (3.73) | (0.68) | | | |
| (5) | Private non-residential fixed investment | 7.11 | 4.70 | | -0.86 | 2.09 | 27.98 | 1929 – 2015 |
| | | (3.99) | (1.21) | | (-0.25) | | | |
| (6) | Private non-residential fixed investment | | | 6.82 | 3.69 | 2.37 | 13.41 | 1929 – 2015 |
| | | | | (4.12) | (1.38) | | | |
| (7) | All | 5.80 | - | | 1.82 | 1.43 | 38.24 | 1939 – 2015 |
| | | (4.34) | (-) | | (0.64) | | | |
| (8) | All | | | 5.84 | 9.36 | 2.21 | 12.55 | 1939 – 2015 |
| | | | | (4.68) | (4.20) | | | |

Table 9: $CO_2$ emissions growth and climate risk

This table reports the cross-sectional pricing power of the temperature risk measured over 1-year, 3-year and 5-year horizon and the cross-sectional pricing power of $CO_2$ emissions growth controlling for each of the three temperature risk measures. Test assets are the twenty-five US portfolios sorted by size and book-to-market ratio. $CO_2$ is annual $CO_2$ emissions per capita growth. Temperature risk over 1-year/3-year/5-year horizon is measured as the 1-year/3-year/5-year difference in average US temperature. In regression (1) to (3), we estimate the factor risk premium for a one-factor model with each of the temperature risk factors one at a time; and, in regression (4) to (5), we estimate the factor risk premia for a two-factor model with $CO_2$ emissions growth controlling for the temperature risk factor. A cross-sectional constant is included in the estimation. Factor risk premia are reported with $t$-statistics adjusted using Newey–West three-lagged corrections in parentheses. Root-mean-square errors (RMSEs) and adjusted $R^2$ are measured as percentages. The sample period is 1896–2015 for regression (1) and (4), 1898–2015 for regressions (2) and (5), and 1900–2015 for regressions (3) and (6).

| | Temperature risk 1-year | Temperature risk 3-year | Temperature risk 5-year | $CO_2$ | Constant | RMSE | Adj. $R^2$ | Sample |
|---|---|---|---|---|---|---|---|---|
| (1) | -41.92 | | | | 8.89 | 2.67 | 9.00 | 1896–2015 |
| | (-2.80) | | | | (3.31) | | | |
| (2) | | 6.23 | | | 10.91 | 2.96 | 1.30 | 1898–2015 |
| | | (0.31) | | | (4.05) | | | |
| (3) | | | -40.10 | | 6.57 | 2.63 | 14.99 | 1900–2015 |
| | | | (-1.99) | | (2.74) | | | |
| (4) | -21.43 | | | 4.60 | 0.13 | 2.15 | 27.03 | 1896–2015 |
| | (-1.54) | | | (1.93) | (0.03) | | | |
| (5) | | -11.51 | | 4.89 | -0.13 | 2.18 | 22.28 | 1898–2015 |
| | | (-0.59) | | (2.17) | (-0.04) | | | |
| (6) | | | -10.75 | 4.58 | -0.25 | 2.20 | 24.69 | 1900–2015 |
| | | | (-0.69) | (2.05) | (-0.07) | | | |

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