# Essays on the Pricing of Financial and Human Wealth




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*To Professor Favero,*
*for suggesting the way.*

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# Abstract

This thesis presents three empirical analyses on the systematic risk exposure that global and domestic asset holders face. Each paper investigates a distinct source of macroeconomic risk, but they all stem from the premise that holding human capital and financial assets is risky. This ultimately affects agents' optimal consumption choices. The first paper, *The International CAPM Redux*, proposes a novel empirical model to price international assets. Building on recent advances in asset pricing research on currency markets, it documents that investors are compensated for bearing exposure to currency risk when investing in foreign equity, either directly or via delegated portfolios of international assets. The second paper, *One Central Bank To Rule Them All*, shows that while global investors demand a premium to bear risks associated with Federal Reserve decisions, there is no comparable result for other major central banks. This puzzling finding points to the uniqueness of the Federal Reserve for global investors, that does not simply stem from the size and importance of the U.S. economy. The third paper, *Human Capital, Unemployment Risk and Asset Prices*, relates the riskiness of human capital to uncertainty in the labour market and documents a role for unemployment as a determinant of human wealth. Sorting U.S. industry-level portfolios by differential exposure to unemployment risk yields a novel cross-section of average excess returns. High risk-premia are consistent with less income-constrained highly educated workers willing to take financial risk beyond hedging their labour income risk. Taken together, these three studies contribute to the empirical asset pricing literature and open the door to further theoretical and empirical research in the field.

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# 1 | Introduction

A major goal in the field of asset pricing is to understand differences in the prices or values of assets, namely claims on a future stream of payments. The empirical observation is that some assets pay historically higher average returns than others. The classic example is the *U.S. equity premium*: over the last 50 years, U.S. investors have earned on average about 8% investing in domestic stocks (in real terms), but have locked in only 1% when investing in the domestic short-term interest rate. This phenomenon is well known as the *equity premium puzzle* (Mehra and Prescott, 1985).

A large body of asset pricing work aims to explain this evidence. More generally, it seeks to document and understand return differentials arising in equity as well as other asset classes (e.g., bond, foreign exchange, commodities), both within and across countries. A prevailing argument for this pervasive phenomenon is compensation for bearing risk exposure.[1] Assets are classified as *safe*, *risky*, or *insurance providers*.

*Safe* assets are assets that carry no uncertainty about their future payments: their returns are known one period in advance. Short-term interest rates are safe assets. Any asset yielding a return above the risk-free rate yields an excess return or a *premium*. Under a risk-based explanation, the magnitude of this premium corresponds to the product of the quantity of risk taken by the asset holder and the compensation per unit of risk as determined by the capital market.

*Risky* assets are assets that pay off poorly in bad states of the world for the investor. Put simply, they pay off badly when asset holders need money. Since rational investors anticipate this possibility, they can be persuaded to buy these assets only by being adequately compensated ex ante. Under this view, the expected

[1] See Campbell (2000) for a broad survey of the asset pricing field.

return on an asset is proportional to the quantity of risk taken by the investor. The higher the risk, the higher the expected return. In turn, risky assets earn historically high returns on average, because they award a premium to their holders for bearing the risk of making a loss in *bad times*.

On the contrary, *insurance providers* are assets that pay off well in *bad times* for the investor. Asset holders who are concerned about particular states of the world are willing to accept a lower return ex ante to ensure their investment will not pay off poorly when adverse states will materialise. Under a risk-based explanation for financial premia, the lower the risk borne by the investor, the lower the expected return on the asset. Assets in this category earn historically low average returns, because they provide insurance to their holders against the risk of making a loss in *bad times*.

One of the major challenges in this research area is to correctly identify the sources of risk that drive asset prices. Only sources of undiversifiable risk - risk that cannot be mitigated through diversification - should be *priced* in capital markets. Equivalently, only assets that are exposed to these risks should command a *risk premium*. Plausible risk drivers, namely *factors*, are variables that are both theoretically defensible and empirically supported. The requirement is therefore twofold: (i) economic significance and (ii) statistical significance. In the 1960s these ideas were fairly new. Nowadays, "we have a zoo of factors" (Cochrane, 2011). As documented by Harvey, Liu and Zhu (2015), at least 316 different factors have been proposed over the last decades raising concerns about the reliability of the empirical analyses and challenging the achievements of the profession.

This thesis contributes to the asset pricing field by proposing three empirical analyses on the systematic risk exposure that global and domestic asset holders face. Each paper investigates a distinct source of macroeconomic risk, but they all stem from the premise that holding human capital and financial assets is risky. This ultimately affects agents' optimal consumption choices. Papers differ with respect to the research question, the asset class and the source of risk under investigation.

Joint work with Tarun Ramadorai and Adrien Verdelhan, *the International CAPM Redux*, examines whether U.S. investors are compensated for bearing exposure to currency risk when investing in foreign equity, either directly or via delegated portfolios of international assets. This paper proposes a novel empirical three-factor model, *the International CAPM Redux*, and documents that (i) it successfully captures foreign exchange risk in a wide cross-section of foreign equity returns and (ii) it also helps explain the risk exposure of mutual funds and hedge funds which are well diversified internationally. A simple complete-markets model replicates these empirical findings and delivers a novel perspective on the optimal currency hedging problem.

Joint work with Pavel Savor and Mungo Wilson, *One Central Bank To Rule Them All*, investigates whether international investors demand a premium for holding stocks on days when major interest rate announcements are scheduled for release. This paper finds that global investors demand a premium to bear risks associated with Federal Reserve decisions, but there is no comparable result for other major central banks, namely the Bank of England, the Bank of Japan, and the European Central Bank. Employment and inflation announcements have impact on local stock markets and, to some extent, even on the U.S. stock market. We interpret these findings as evidence that the Federal Reserve exerts a unique impact on global equity prices that does not simply stem from the size and importance of the U.S. economy.

In the third paper, *Human Capital, Unemployment Risk and Asset Prices*, the focus of the analysis shifts from the riskiness of financial assets to the riskiness of human capital, and from global financial markets to the United States. This paper relates the riskiness of human capital to uncertainty in the labour market and documents a role for unemployment as a determinant of human wealth. I derive two unemployment-adjusted measures of labour income from the labour market equilibrium outcome and use micro-level survey data to document their historical dynamics at aggregate level. I work with representative agents who differ in their endowment of human capital (i.e. education/work experience). My findings suggest that inequality in human wealth – the discounted value of future labour

income – is worse than the well documented inequality in labour income. Sorting U.S. industry-level portfolios by differential exposure to unemployment risk yields a novel cross-section of average excess returns. I interpret the high risk-premia as compensation for less income-constrained highly educated workers willing to take financial risk beyond hedging their labour income risk.

Given the distinct nature of each analysis, I review the major insights of each paper separately, highlighting the relevance, contribution to the field and the key findings. A detailed survey of the existing body of research related to these topics is presented in Chapter 2.

**Paper 1: The International CAPM Redux.** Investigating whether international investors are compensated for bearing exposure to currency risk has a long tradition in international asset pricing studies. In a world of real rigidities, economic theory predicts that exchange rate risk matters to the pricing of risky assets. Since international investors invest abroad, but consume at home, deviations from purchasing power parity affect their consumption paths.[2] This implies an equilibrium compensation for the risk of low returns on foreign investments once these returns are expressed in real domestic terms.

This question has become increasingly relevant over time due to the steady increase in international financial integration. Aggregate foreign equity holdings as a percentage of global gross domestic product, for example, have increased steadily from roughly 3% in the 1980s to approximately 30% in 2011, alongside a steady dismantling of de-jure cross-border capital flow restrictions. In the United States, foreign equity holdings are currently worth roughly US$ 6 trillion.

Currency risk is incorporated in most plausible theoretical asset pricing models (Solnik, 1974; Sercu, 1980; Adler and Dumas, 1983), but empirical work has hardly provided convincing evidence of its pricing in global equity markets. Dumas and Solnik (1995) is a notable exception. The standard approach to address our research question is to empirically test the *International Capital Asset Pricing Model (CAPM)*

[2]Purchasing power parity postulates that national price levels should be equal when converted to a common currency.

of Adler and Dumas (1983). Taken literally, this model requires excess returns on all currencies in the world to be risk factors in addition to a world equity factor. This is empirically cumbersome.

The novelty of our paper is to depart from the traditional approach of restricting the full set of currencies in the world to a small set of noisy bilateral exchange rates to overcome this empirical issue.[3] Instead, our strategy is to fully approximate the foreign exchange risk component postulated by the theory via two currency risk factors, only. These two factors, dollar and carry, effectively summarise variation in a broad cross-section of bilateral exchange rates (Verdelhan, 2015) and act as a noise-reduction scheme.[4] This twist yields a novel three-factor model, the *International CAPM Redux*, that successfully captures risks in international portfolios of assets. The three factors embedded in this new model are the excess return on a world market portfolio denominated in local currency terms, and the two aforementioned currency factors.

First, we document that our model delivers an important role for currency risk in a wide cross-section of foreign equity returns. We compile a comprehensive dataset of equity returns spanning value, size, and country aggregate market portfolio returns from 25 developed markets and 21 emerging markets between February 1976 and April 2013. In conditional asset pricing tests, we show that our model outperforms the single-global factor World CAPM of Sharpe (1964) and Lintner (1965) as well as the "Classic" International CAPM of Adler and Dumas (1983) tested by Dumas and Solnik (1995), and also delivers comparable performance to the global versions of the popular Fama-French factors (Fama and French, 2012).

Second, we confirm that our approach is also relevant to those interested in delegated portfolios of international assets. Our three factors offer a simple metric to infer the amount of currency risk taken on by international investment managers, a

[3]This small set usually collapses to the four most important currencies in the world, namely the U.S. dollar, the British pound, the Japanese yen and the German mark or, more recently, the euro.

[4]The dollar factor is the average excess return earned by an investor that borrows in the U.S. and invests in a broad portfolio of foreign currencies. The carry factor is the average excess return earned by an investor going short (long) a portfolio of low (high) interest rate currencies.

notably difficult task. We assemble a cross-section of returns earned by mutual funds and hedge funds invested internationally over the last 25 years and find that our model explains a significant fraction of these returns. This evidence suggests that the exposure to currency risk taken by these managers can potentially be reproduced at a much lower cost than by incurring the substantial fees actually charged.

Third, to better understand the dynamics captured by our empirical risk factors, we build a simple theoretical model in which exchange rates, currency risk factors, and equity market returns are all precisely defined. In leading international equity asset pricing models, exchange rate variation arises exogenously. In contrast, we assume that exchange rates are driven by the same shocks that drive the pricing kernels. We assume that financial markets are complete and posit the law of motion of the lognormal stochastic discount factors (SDFs) in all countries.

In our setup, both country SDFs and country aggregate dividend growth rates depend on country-specific shocks, as well as three global shocks. To create a role for a pure equity risk factor (which affects all countries in the same way), we allow one of the global shocks to affect all SDFs similarly. Since changes in exchange rates are differences in log SDFs, this global shock does not show up in currency markets, resulting in a degree of segmentation between equity and currency markets. This set up yields a very tractable framework in which (realized and expected) equity and currency returns can be written down in closed form.

Solving this model reveals that the world aggregate equity return expressed in a common currency does contain all relevant information necessary for pricing international assets, meaning that there should be no role for bilateral exchange rates. The twist is that time-variation in the prices of global shocks – necessary to account for time-varying currency risk premia – confounds empirical estimation using a single-factor model, especially when the relevant state variables are unknown to the econometrician. Including the dollar and carry factor arises as a natural solution to the challenges of empirical estimation in this framework.

We use the model to shed light on our main empirical results, calibrating it to match a large set of equity and currency moments, and generating simulated data.

We apply exactly the same estimation procedure as we do on real data to these simulated data, and replicate our empirical finding that the International CAPM Redux outperforms both the World CAPM and the "Classic" International CAPM. We note that time-variation in the prices of risk appears key to this result.

Finally, we use our model to revisit a longstanding and important issue in international finance, namely optimal currency hedging in international portfolios. The prevailing wisdom arises from the mean-variance efficient framework of Black (1989) in which the optimal amount of currency hedging depends on the mean and standard deviation of world market returns, as well as on exchange rate volatility. This implies that the optimal amount of currency hedging is constant through time and across investor location. In contrast, in our model exchange rates and equities are driven by many of the same risks, thus altering the inference. We show that the optimal amount of investment in currency portfolios such as carry and dollar is time-varying and investor-location-specific in our set up, and also depends on the prices of risk of both country-specific and global shocks.

**Paper 2: One Central Bank To Rule Them All.** The role and importance of central banks is documented by a large literature, spanning both economic and financial fields. Central banks have large impact on security markets and are key players during crisis periods. Their decisions are also covered by the popular press and have attracted a great deal of attention especially in the recent years, due to extraordinary macroeconomic circumstances (e.g., zero interest rates and unconventional monetary policy).

Historically, researchers have relied on pre-scheduled macroeconomic and monetary policy announcements to infer a casual relationship between asset prices and major economic news. According to the *efficient market hypothesis*, asset prices should immediately incorporate any publicly available information into prices. Early studies found weak support for this hypothesis in U.S. asset markets, except for monetary news, but the availability of high-frequency data has enabled major progress in the field. Bernanke and Kuttner (2005), for example, show

that U.S. equity prices respond strongly to Federal Open Market Committee (FOMC) interest rate decisions.[5]

A more recent body of asset pricing work has shifted the focus of the analysis from the sensitivity of asset prices to the impact of pre-scheduled announcements on expected returns, thereby investigating whether investors demand a premium for bearing macroeconomic risk. Since the timing of pre-scheduled announcements is publicly available, if asset prices respond to news and investors care about those news, asset returns should be predictably higher on announcement days, as investors should be compensated in equilibrium to hold securities around releases. Jones, Lamont and Lumsdaine (1998) find that both average long-term Treasury bond returns and their volatilities are higher on U.S. inflation and employment announcement days relative to all other days. Savor and Wilson (2013) document that in the 1958-2009 period the average excess daily return on a broad index of U.S. stocks is 11.4 basis points on announcement days (i.e., FOMC, employment, and inflation releases), but only 1.1 basis points on non-announcement days, and the effect is larger for monetary policy announcements than for macroeconomic announcements. Lucca and Moench (2015) uncover that the high average U.S. equity returns earned on FOMC days are mainly collected ahead of FOMC announcements and international equity exhibits similar patterns as well. Mueller, Tahbaz-Saleni and Vedolin (2015) establish both a pre- and a post- FOMC announcement premium in currency markets. Cieslak, Morse and Vissing-Jorgensen (2015) document that the FOMC premium does not simply pertain FOMC announcement days, but it is a broader phenomenon. Since 1994, the U.S. equity premium is earned entirely in even weeks (i.e., 0, 2,4, and 6) in FOMC cycle time, that is the window between a FOMC meeting and the following one. Furthermore, on FOMC announcement days, the Capital Asset Pricing Model (CAPM) explains the cross-section of stock returns, with a strong positive relation between market beta and returns (Savor and Wilson (2014)).

In this paper, we investigate whether there exists an equivalent announcement-day-premium effect for non-U.S. stocks, on days when the Federal Reserve (Fed)

[5]FOMC meetings are pre-scheduled monetary policy meetings in the United States.

as well as major non-U.S. central banks release interest rate decisions. We find a very strong FOMC effect (often significantly stronger than in the U.S.) for almost all international stock markets. The overall global average market excess return is 27.6 bps on FOMC days compared to 2.7 bps on non-announcement days, and the difference is positive (and significant) in 37 (27) out of 38 countries. The effect is substantially stronger in more recent years, with FOMC days accounting for the bulk of the excess return earned by equity investors over the 1998-2013 period. At the same time, and despite remarkable analogies, we find virtually no effect for the announcements by non-U.S. central banks, either in the U.S. stock market or, more interestingly, in their home markets.

During the 1999-2013 period, the FOMC, the Bank of Japan (BoJ), the European Central Bank (ECB), and the Bank of England (BoE) all held regular, pre-scheduled meetings after which they announced their decisions about current monetary policy, including changes to target interest rates, monetary aggregates (in the case of the ECB), and planned asset purchases. All four central banks pursued an active monetary policy over this period, and all four currency zones possess large, liquid, and important equity markets, both by turnover and market capitalization. However, investors in Europe, the United Kingdom, and Japan seem to demand a higher risk premium for risks associated with FOMC decisions than those by their own central banks.

These results raise an important question: Why is the Fed so powerful compared to other central banks, even those in charge of monetary policy for comparable economies? We discuss a number of potential explanations.

One straightforward possibility is that the domestic economy associated with some central banks is too small for macroeconomic announcements to affect risk premia, even in their own markets. This argument potentially holds for countries such as Switzerland or even the U.K., whose stock market indices contain a large share of multinational firms that are not overly exposed to the local economy. However, it is a less plausible hypothesis for a country like Japan or a whole region like the euro-zone. To address this hypothesis, we examine other non-U.S.

macroeconomic announcements, such as inflation and employment, finding in many cases a significant effect on average domestic stock market returns, as well as a non-trivial effect on U.S. market returns. We interpret these results as evidence that economy size does not explain our results.

A second possibility is that non-U.S. central banks potentially matter to investors, but they choose not to pursue an active monetary policy (or at least have not done so in the past), thereby greatly reducing investor uncertainty regarding their decisions. Contrary to this hypothesis, in the period we study all three of the world's non-U.S. major central banks at times exhibit an avowedly activist policy, with large variations in interest rates as well as use of unconventional monetary policy tools.

A third related explanation is that non-U.S. central banks actually pursue an active policy, but that their decisions are widely anticipated in advance, whether by design or inadvertently, so that their is little scope for the markets to be surprised by their scheduled announcements.[6] To address this hypothesis, we study futures-implied interest rate changes for the euro, yen, and sterling. We find that ECB and BoE announcement days are associated with much larger absolute changes in interest rate expectations than non-announcement days, which is consistent with the hypothesis that these central banks at least occasionally surprise investors. For the BoJ, we do not find a comparable result, which is not surprising given that BoJ rates did not move much during our sample period (the BoJ activism manifested itself primarily through its use of unconventional monetary policy tools).[7]

A further explanation is that the Federal Reserve is simply unique among the world's central banks. One interesting, though at this stage speculative, possibility is that the Federal Reserve sets the global price of money, at least money used to finance securities trading, and that other central banks are just a sideshow.[8] However, even in this case our results sound puzzling, considering that some of

[6]Cieslak, Morse and Vissing-Jorgensen (2015) find evidence consistent with the hypothesis that the Fed often releases important information outside its regular meetings.

[7]More informally, we also identify through Google searches multiple surprise decisions by all three central banks, which resulted in significant moves in equity, fixed income, and currency markets.

[8]This may occur because the U.S. dollar's role as the dominant currency for trade (see, e.g., Gopinath (2015)) gives the Fed special freedom of action.

these other economies are large and important global players. The combined GDP of the eurozone countries, for example, was actually larger at the end of 2013 than that of the U.S., and both economic regions have open economies and active capital markets. Under this explanation, our results highlight the role of the dollar investor as the marginal investor in global stocks, U.S.- and foreign-traded alike. Finally, the Fed may be special due its capabilities. Its researchers and proprietary data sources (see, e.g., Cieslak, Morse and Vissing-Jorgensen (2015)) could provide it with better insight into global economic prospects than any of its peers, making its pronouncements much more important to investors locally and globally.

At this stage, we conclude that the Federal Reserve is unique in its importance to global investors, and that this uniqueness does not simply stem from the size of the U.S. economy and its securities markets.

**Paper 3: Human Capital, Unemployment Risk and Asset Prices.** The importance of embedding human wealth, the present discounted value of future labour income, in both empirical and theoretical asset pricing work has been recognised since the 1970's. The first contributions date back to Mayers (1972) and Roll (1977). Despite the profession has made substantial progress towards the understanding of the determinants of financial wealth since then, attention to human wealth has been intermittent and low in volume. Four decades later, the subject still matters. A recent empirical study finds that human wealth accounts for about 90% of the total wealth of U.S. households (Lustig, Van Nieuwerburgh and Verdelhan, 2013).

A major challenge to this subject is posed by human capital being a non-tradable asset. This feature makes the treatment of human wealth cumbersome, as cash-flows are observable (wages), but discount rates are not. In the asset pricing field, researchers have mainly taken two approaches to address this issue over time. Early empirical asset pricing work imposes ad-hoc restrictions on the human capital discount rate (Campbell, 1996; Shiller, 1995; Jagannathan and Wang, 1996), while recent work adopts VAR techniques to recover it from aggregate consumption data conditioning on a set of standard predictors for consumption growth and stock returns (Lustig and Van Nieuwerburgh, 2007; Bansal, Kiku, Shaliastovich

and Yaron, 2014). This paper proposes an alternative and novel perspective: it sheds light on the riskiness of human capital by empirically pinning down its determinants in the labour market.

Human capital is a risky asset, owned by workers and invested in the labour market. Under these premises, human capital risk can be mapped into labour market risk and empirically measured. My starting point is a straightforward insight from labour macroeconomics: At each point in time the labour market equilibrium outcome is jointly determined by wage and unemployment dynamics. I propose a simple measure of human capital that accounts for both channels and quantify it at aggregate level using U.S. micro-level data. I work with representative agents who differ in their endowment of human capital. Building on the wide labour literature on the returns to education, I define heterogeneity with respect to the educational investment these agents made prior to entering the labour market (i.e., diploma, college, or post-college education).

Abstracting from governmental transfers, this human capital measure adjusts the growth rate of wages by the unemployment risk borne by employed workers over the same horizon. In other words, it is growth in unemployment-(U-) adjusted labour income. To the best of my knowledge, existing work in asset pricing accounts for wages (human capital cash-flows), but does not directly incorporate unemployment dynamics.[9] This paper contributes to the literature on human wealth by documenting a key role for the unemployment channel beyond the more traditional wage channel. I interpret unemployment as the bad state of the world faced by any employed worker and unemployment risk as the downside risk he bears. I take the probability of being unemployed as a proxy of the worker's exposure to unemployment risk, and propose the unemployment rate as an empirical proxy for

[9]Standard asset pricing measures allow unemployment to affect human cash-flows, as national labour income data include unemployment benefits. Human capital measures derived from life-cycle models allow unemployment to affect the human discount rate (e.g. Brown, Fang and Gomes (2012)). However, since these models concern the choices of individual households over the life-cycle, a substantial share of the risks they capture is idiosyncratic rather than systematic. These measures are therefore not useful in asset pricing work.

this probability at aggregate level. I measure the unemployment rate conditioning on the endowment of human capital. I draw inference in two steps.

First, I present an historical comparison of the cumulated quarterly growth in U-adjusted labour income experienced by a representative diploma, college and post-college holder. I examine these time-series from 1979 to 2014 alongside simple wage growth measures compiled at the same level of aggregation. All series are constructed by assembling Current Population Survey data (Outgoing Rotation Group) – the major source of labour market statistics in the United States – and generalised to the U.S. labour force by means of population sampling weights. To improve precision of estimates, I estimate the wage-related component of U-adjusted labour income via quasi-experimental methods (propensity score weighting).

I show that, in *bad times*, highly educated workers historically experience a less severe drop in U-adjusted labour income than low educated workers.[10] This finding points to human capital being historically riskier for low educated workers than highly educated ones. By definition, risky assets are assets that pay-off poorly in bad times for the investor. Since labour income is the pay-off of the human capital asset, human capital is risky if labour income drops in bad times. Accounting for unemployment dynamics is key to draw this conclusion: wages are sticky, while unemployment captures a deterioration of labour market conditions in real time. As revealed by the data, the unemployment component of U-adjusted labour income visibly crashes in recessions, while there is no comparable evidence for its wage component. These findings, coupled with the fact that average wages are historically higher for highly educated agents than low educated ones, have implications for human wealth. Inequality in human wealth has to be worse than the well established inequality in labour income. Overall, this evidence points to unemployment being a key driver of human capital discount rates.

As a second step, I investigate whether these findings imply heterogeneous financial investment behaviour. Intuitively, agents who face high exposure to changes in labour market conditions (low educated) should invest in financial

[10] *Bad times* are defined as quarters when firms search for new workers by posting vacancies at a low rate, as revealed by the job posting index of Barnichon (2010).

markets mainly to hedge their labour income risk – investing in assets that perform well when they are more likely to lose their jobs – while agents who face mild exposure (highly educated) should be willing to bear the risk of a poor performance of their financial assets in bad labour market states – investing primarily in risky financial assets – if compensated in expectation. I follow a standard long-short approach to build a macroeconomic factor that captures differential exposure to unemployment risk in the United States. This variable is defined as the difference between the growth in quarterly employment rates of highly educated workers (post college holders) and low educated ones (diploma holders).[11]

Sorting industry-level portfolios by their time-varying loadings on this factor − controlling for aggregate market risk − yields a novel cross-section of average excess returns. A standard long-short strategy, long in high-$\hat{\beta}^{Emp}$ industries and short in low-$\hat{\beta}^{Emp}$ industries, yields a statistically significant average premium of about 4% over the last 35 years. These risk premia are consistent with heterogeneous financial investment behavior as outlined above, and driven (i) by the demand of highly educated workers, and (ii) by bad labour market states (6.5% in bad times vs. 2% in good times). Finally, I provide preliminary evidence about the relationship between this novel risk premium and the value premium.[12] These two sources of risk are closely related, but do not fully overlap.

The thesis is organised as follows. Chapter 2 presents an extensive literature review spanning the three topics above. Chapter 3, Chapter 4 and Chapter 5 correspond, respectively, to my first (*The International CAPM Redux*), second (*One Central Bank To Rule Them All*) and third (*Human Capital, Unemployment Risk and Asset Prices*) papers. Chapter 6 concludes highlighting the general implications of these three-paper thesis and directions for future work.

[11]This difference corresponds to the spread in the unemployment component of U-adjusted labour income of the two extreme representative agents.

[12]The value-premium is the average excess return difference between high book-to-market (value) and low book-to-market (growth) stocks.

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# 2 | Literature Review

Standard asset pricing theory rationalises large differences between the average returns on equity and short-term interest rates as a *premium* for bearing non-diversifiable aggregate risk. Contrary to this prediction, the empirical magnitude of this premium largely exceeds the spread postulated by the theory. Since Mehra and Prescott (1985) a large literature has investigated this enduring phenomenon – the *equity premium puzzle* – with both risk-based explanations and alternate explanations put forward.[1]

More broadly, finance theory explains the expected excess return on any risky asset over the risk free rate as the product of the quantity of risk and the price of risk. Theoretically, the quantity of risk corresponds to the covariance of the excess return on the risky asset and the source of risk driving the price of that asset (i.e., the *factor*). Empirically, it is the regression coefficient of the former on the latter (i.e., the *beta*).

In the Consumption Capital Asset Pricing Model (CAPM), for example, the covariance between the asset return and consumption growth explains the premium on the asset, but the smoothness of consumption in the data fails to validate the theory. In the Capital Asset Pricing Model (CAPM), one of the most widely used models to price assets in the financial industry, the covariance between the asset return and the excess return on the market portfolio explains the premium on the asset, but a wide literature has established its poor performance in the data.

These findings have spurred a large literature attempting to identify sources of risk that can explain patterns in excess returns in time-series, and account for a

[1] See Mehra and Prescott (2003), Donaldson and Mehra (2008) and Mehra and Prescott (2008) for a review of the literature on the *equity premium puzzle* and evidence on its various explanations.

large share of variation in the cross-section of expected returns. More broadly, the goal is to pin down risk exposure that is compensated in capital markets.

This three-paper thesis broadly fits this research area. This chapter provides an extensive review of the existing work mostly related to the topics discussed in the next three chapters. I start reviewing studies in international asset pricing, next move to the relation between macroeconomic/monetary policy announcements and asset prices, and conclude with the contributions on the value and risk of human capital.

# 1 International Asset Pricing

The issue of pricing foreign assets lies at the hearth of the international asset pricing literature. This question has been examined by a flourishing body work over the last four decades, both theoretically and empirically.

Early theoretical work in international asset pricing generally assumed that consumption and investment opportunity sets do not differ across countries. This restrictive assumption of investor indifference between domestic and foreign markets was removed in later work.[2] With country-specific investor heterogeneity in preferences, or homogenous preferences but differences in relative prices faced by investors in different countries, currency risk matters for asset pricing. In a world of deviations from purchasing power parity (PPP), bilateral exchange rates do not fully adjust to equalise national price levels expressed in a common currency.[3] In such a world, investing in foreign currencies is risky, as adverse exchange rate movements imply low income from foreign investment when expressed in domestic currency terms. In this context, Solnik (1974), Sercu (1980), Stulz (1981), and Adler and Dumas (1983) incorporated currency risk into theoretical asset pricing models, allowing for differences across countries in consumption opportunity sets.

[2]See Glen and Jorion (1993) and Stulz (1995) for a comprehensive survey of the theoretical and empirical literature on international portfolio choice and asset pricing. More recently, Campbell, Serfaty-De Medeiros and Viceira (2010) consider optimal currency hedging in international equity investment, while Kroencke, Schindler and Schrimpf (2013) study the optimal portfolio allocation across equity and exchange rate investment styles.

[3]Empirical work suggests that PPP holds only in the long run. See Froot and Ramadorai (2005) for a comprehensive survey of the early literature on PPP.

Specifically, Solnik (1974) modeled exchange rates as cross-country relative prices of consumption baskets, and Sercu (1980) generalized the approach, allowing stock returns expressed in local currencies to be correlated with exchange rates. Stulz (1981) and Adler and Dumas (1983) assumed stochastic country-specific inflation in addition to deviations from PPP.[4]

Early empirical work in international asset pricing generally extended the standard Capital Asset Pricing Model (CAPM) (Sharpe, 1964; Lintner, 1965) to global markets, or simply augmented the domestic CAPM with the addition of a few international factors. In this early literature, unconditional tests of these models generally proved inconclusive (see, for example, Stehle Richard (1977), Solnik (1974) and Korajczyk and Viallet (1989)).

Later conditional studies yielded more promising results: for example, Bekaert and Harvey (1995) provided evidence that countries' capital markets become increasingly globally integrated over time, Chan, Karolyi and Stulz (1992) identified a time-varying global market premium in U.S. equity markets, and Karolyi and Stulz (1996) studied co-movement in Japanese and U.S. stock markets.[5] Ferson and Harvey (1993) showed that most of the predictability of foreign equity returns is related to time-variation in global risk premia, while Bekaert and Hodrick (1992), and Bekaert (1995, 1996) studied the predictability of equity and currency returns.

[4]In the model of Adler and Dumas (1983), there are $L+1$ countries, and a set of $m = n+L+1$ assets – other than the base-currency deposit – comprised of $n$ portfolios of equities, $L$ foreign currency deposits, and the world market portfolio. In that model, the expected return on asset $j$ is:

$$E_t\left(r_{j,t+1} - r_{f,t}\right) = \sum_{i=1}^{L} \lambda_{i,t} cov_t(r_{j,t+1}, r_{n+i,t+1}) + \lambda_{m,t} cov_t(r_{j,t+1}, r_{m,t+1}) \tag{1}$$

where $r_{j,t+1} - r_{f,t}$ is the nominal return on the equity portfolio $j$ in excess of the risk-free rate (denoted $r_{f,t}$) of the currency in which returns are measured, and $r_{m,t+1}$ is the excess return on the world market portfolio. The covariance terms $cov_t(r_{j,t+1}, r_{n+i,t+1})$ measure the quantity of exchange rate risk. The time-varying coefficients $\lambda_{i,t}$ are the world prices of exchange rate risk. The time-varying coefficient $\lambda_{m,t}$ is the world price of market risk. The unconditional version of this model is presented later in this section.

[5]A large related literature studies the integration of emerging and developed equity markets (see, e.g., Bekaert and Harvey (2000); Bekaert, Harvey and Lundblad (2003); Bekaert, Harvey, Lundblad and Siegel (2007b); Bekaert, Hodrick and Zhang (2009); Bekaert, Harvey, Lundblad and Siegel (2011); Bekaert and Harvey (2014)).

Harvey (1991) introduced a novel approach to the literature, modelling time-variation in both the exposure and the price of risk by conditioning on common and country-specific set of instruments. This study found that time-variation reveals differences across countries, but fails to fully predict conditional expected returns. Implementing a variation on this methodology, and incorporating currency risk explicitly, Dumas and Solnik (1995) found that an exchange rate risk model outperforms a simple "World CAPM", that is a model with global market risk but no explicit role for currency risk. In their empirical work, Dumas and Solnik (1995) assumed that currency risk for a U.S. investor is well captured by the excess returns on three major currencies: the Japanese Yen, the British Pound, and the German Mark. Providing evidence that currency risk is priced in global capital markets, they attributed the widespread failure of the World CAPM to a misspecification problem. These findings were subsequently corroborated by De Santis and Gerard (1998), who extended the analysis to account for volatility dynamics. Harvey, Solnik and Zhou (2002) also provided support for currency risk in international equity returns. Using latent factors, they found that their first factor premium resembles the expected return on the world market portfolio, while their second factor premium is related to foreign exchange risk. These studies provided strong empirical support for the predictions of international asset pricing theory, but were generally implemented on a small sample of assets from developed countries, over relatively short sample periods.[6]

More recently, a variety of alternate explanations have been proposed for differences in international average equity returns (see Lewis (2011) for the most complete and recent survey of international asset pricing). These explanations include global economic risks (Ferson and Harvey, 1994); inflation risk (Chaieb and Errunza, 2007); liquidity risk (Karolyi, Lee and van Dijk, 2012), Bekaert et al. (2007a), Malkhozov et al. (2014) factors including momentum and a global cash-flow-to-price factor (Hou, Karolyi and Kho, 2011); and investability restrictions

[6]Dumas and Solnik (1995) considered a set of four countries (i.e., United Kingdom, Germany, Japan and the United States) from March 1970 to December 1991. De Santis and Gerard (1998) took into account the same countries from June 1973 to December 1994.

(Karolyi and Wu, 2014). These specifications do not account for currency risk; nevertheless they significantly outperform the World CAPM. This is also a feature of the work of Fama and French (2012), who present international versions of both the Fama and French (1993) three-factor model, and Carhart (1997) four-factor model. Fama and French (2012) find that their global models perform reasonably well in unconditional time-series tests on global size and book-to-market sorted portfolios as well as size and momentum sorted portfolios, but perform worse when the same portfolios are constructed at a regional level (i.e., North America, Japan, Asia Pacific, and Europe).

A separate strand of asset pricing work has focused specifically on the pricing of currency returns. Relative to research on equity markets, this literature has a short tradition – contributions are limited to the recent years - but is currently a very active and flourishing research area.

It is well-known (Meese and Rogoff, 1983) that economic models of exchange rate determination generally lack empirical support in the short-run, with few identifiable links between nominal exchange rates and economic fundamentals.[7] A recent asset pricing literature has attempted to explain returns on portfolios of currencies as compensation for risk.

Carry trades are speculative strategies in foreign exchange markets based on exploiting interest rate differentials across countries. Investors go long a basket of high interest rate currencies and short a basket of low interest rate currencies seeking to profit on average. The historical success of this strategy provides evidence for the failure of the *uncovered interest parity condition*, one of the major puzzles in International Finance. Under this condition, carry trades are not profitable ex-ante, as any potential gain yielded by the interest rate differential is expected to be exactly offset by an exchange rate change during the horizon of the investment. Applying a risk-based arbitrage pricing view (Ross, 1976) to currency markets,

[7]This phenomenon is known as the *exchange rate disconnect puzzle*. At high frequencies, a number of papers, including Evans and Lyons (2002), and Froot and Ramadorai (2005), show that order flow in exchange rate markets is helpful at predicting and explaining exchange rate movements, but there is considerable debate about the source of this explanatory power, with both rational and behavioral explanations rationalizing these findings.

Lustig, Roussanov and Verdelhan (2011) show that the cross-section of currency portfolios sorted by interest rates can be well-explained by a "slope" factor – the "carry" factor – with high interest rate currencies loading more on this factor than low interest rate currencies.[8] Backus, Foresi and Telmer (2001) show that heterogeneity in exposure to country-specific risk can generate negative UIP slope coefficients for individual currency pairs, but it cannot explain the cross-section of carry trade returns. On the contrary, Lustig, Roussanov and Verdelhan (2011) show that heterogeneity in exposure to common risk – captured by the carry factor – can both explain carry trade returns and deliver negative UIP coefficients. In this view, carry trade investors earn high excess returns on average for taking exposure to global risk. Currency factors have been linked to several observable variables, such as measures of global volatility in equity markets (Lustig et al., 2011); global volatility in currency markets (Menkhoff, Sarno, Schmeling and Schrimpf, 2012); downside risk in equity market (Lettau, Maggiori and Weber, 2014); and proxies of financial stress (Maggiori, 2012).

Explanations for the profitability of the currency carry trade includes consumption growth risk (Lustig and Verdelhan, 2007) and liquidity risk (Brunnermeier, Nagel and Pedersen, 2008; Brunnermeier and Pedersen, 2009). Specifically, (Lustig and Verdelhan, 2007) show that U.S. consumption growth risk is priced in currency markets, while Brunnermeier et al. (2008) and Brunnermeier and Pedersen (2009) argue that carry traders are subject to funding liquidity constraints. On the contrary, a strand of the literature strongly rejects the risk-based view (e.g., see Burnside, Eichenbaum and Rebelo (2011a); Burnside (2011)), and proposes alternate explanations. In this vein, Burnside, Eichenbaum and Rebelo (2009) investigate the role of market frictions, and Burnside, Eichenbaum, Kleshchelski and Rebelo

[8]This cross-sectional approach was introduced by Lustig and Verdelhan (2005), who departed from the previous literature used to investigate the relation between foreign currency investments and interest rates in time-series. Sorting currencies by their interest rate differentials is helpful, as changes in exchange rates that are purely idiosyncratic – and thus not priced in currency markets – average out.

(2011b) link carry trades to *peso problems*.[9] The latter explanation has been recently ruled out by Jurek (2014) using data on out-of-the-money foreign exchange options.

Most recently, Verdelhan (2015) shows that the cross-section of currency portfolios sorted by exposure to the dollar factor – the average change in the exchange rate between the U.S. dollar and all other currencies – can be well explained by a different "slope" factor, long in high dollar beta currencies and short in low dollar beta currencies. Being nearly orthogonal to the carry factor, this novel factor captures a distinct source of risk in currency markets. These two factors together are shown to account for a large share of the variation in bilateral exchange rates, indicating that at least two global risk factors are necessary to describe dynamics in bilateral exchange rates.

Finally, with respect to further advances in this field, Della Corte, Riddiough and Sarno (2016) uncover a "global imbalance" risk factor that captures the spread in countries' external imbalances, and Della Corte, Ramadorai and Sarno (2014) present a new currency strategy – the *volatility risk premium* – that reflects the costs of insuring currency volatility fluctuations. This strategy, that corresponds to the difference between expected realised volatility and model-free implied volatility, sells high-insurance-cost currencies and buys low-insurance cost currencies, and generates returns mainly attributable to movements in spot exchange rates rather than interest rate differentials.

# 2 Macroeconomic Announcements, Asset Prices, and Central Banks

In both Economics and Finance, the relationship between macroeconomic announcements and asset prices has attracted an increasing deal of attention over time. Asset pricing work on this subject has been particularly prolific in the last decade.

Financial markets are constantly hit by news, mostly at random, challenging researchers seeking to infer a relationship between asset prices and news from the

[9]The *peso problem* relates the failure of the UIP to the presence of expectational errors, that is the possibility that some infrequent or unprecedented event may occur affecting currency prices.

data. Pre-scheduled announcements help address this issue, as their content is unknown ex-ante, but investors face no uncertainty about the day (and often the timing) such information will become publicly available. Major macroeconomic and monetary policy announcements are pre-scheduled and their calendar releases are made available to the public well in advance. For this reason, they have been the object of an extensive literature.

A large body of empirical work in Finance has examined the reaction, or *sensitivity*, of realised asset returns to the unexpected component of scheduled macroeconomic announcements, mainly in the United States. According to the *efficient market hypothesis*, asset prices fully reflect publicly available information. This implies that asset prices should respond to economic releases, as any surprise – the difference between the actual and the expected component of the announcement – should be immediately incorporated into prices.

Empirically, a major issue in addressing this question is posed by the so-called *identification problem*, that is to correctly establish, or *identify*, a causal relationship between the asset price response and the announcement. Since asset prices should only react to the *unexpected* component of the announcement, the first step of the analysis requires to extract news from the aggregate announcement, that is to construct a *surprise* measure, while the second step commands to establish causality between this variable and the observe change in asset prices. These studies are subject to at least two problems. Firstly, many factors, some of them potentially even unobservable, may affect asset prices, so that the announcement effect, if any, can be attributed to an incorrect origin and thereby be overestimated or even be spurious. The second problem is *endogeneity*. A central bank announcement, for example, may be released due to a change in the economic outlook. In this case, it can be hard to disentangle whether movements in asset prices are due to the announcement or to the change in economic conditions.

The *event-study* methodology is the traditional approach adopted by financial studies in this area. Vector autoregression models (VARs) have been widely used to circumvent the issues above, especially in the monetary policy literature, but

this methodology has been subject to many critiques. A third popular strategy is to use survey data to measure expectations, but data availability is usually limited (in both time-series and cross-section), and survey-based measures are imperfect proxies of investors' beliefs. The availability of high-frequency data has helped overcome econometrics issues, as reviewed below.

Early contributions on the sensitivity of asset prices to economic news found weak support for this hypothesis in U.S. asset markets, except for monetary news. Schwert (1981) documented a negative, though small in magnitude, reaction of daily stock market returns to inflation (CPI) announcements. A broader set of studies found a significant negative response of stock prices to weekly money supply (M1) announcements made by the Federal Reserve (Lynge, 1981; Cornell, 1983; Pearce and Roley, 1983). Berkman (1978) was the first to document a similar pattern in bond markets. This evidence was subsequently corroborated in more comprehensive datasets of macroeconomic variables by Pearce and Roley (1985) and Hardouvelis (1987). The former found strong support for the reaction of stock prices to monetary policy news, weak support for reaction to inflation surprises and no support for real activity surprises. The latter confirmed that both interest rates and equity prices are systematically affected by monetary news, but found no comparable evidence for nonmonetary news.[10] Dwyer and Hafer (1989) and Hakkio and Pearce (1985) drew similar conclusions when investigating interest rate and foreign exchange markets, respectively, while Cook and Korn (1991) found a strong interest rate reaction to the employment report. Working in a VAR framework, Cutler et al. (1989) concluded that less than one third of the variance of equity returns is explained by macroeconomic news. Bomfim (2003) documented that U.S. interest rate announcements significantly affect stock market volatility with the caveaut that positive surprises tend to have a larger impact than negative surprises.

All these studies assume that the reaction of asset prices is the same over the business cycle. Later work showed that accounting for non-linearities alters this

[10]Pearce and Roley (1985) examine money supply, the Federal discount and surcharge rate, money supply, the consumer and the producer price indices, the unemployment rate, and the industrial production index. Hardouvelis (1987) extends this dataset to 15 macro-variables.

inference. Orphanides (1992) documented that the equity response to unemployment news depends on the average unemployment rate during the previous year. McQueen and Roley (1993) showed that equity prices strongly react to positive news about inflation, industrial production and the unemployment rate, but the response is negative in good business conditions, and positive in weak economic states.[11] More recently, Boyd, Hu and Jagannathan (2005) focused on unemployment news and found that, on average, announcements of rising unemployment have a positive impact on stock prices during economic expansions, but a negative impact during contractions.

Parallel work on the U.S. bond market at the daily frequency found that announcements of unanticipated unemployment revisions affect Treasury prices (Krueger and Fortson, 2003), and the unexpected component of major macro data releases significantly move long-term forward rates (Gürkaynak et al., 2005b).

A subsequent wave of studies in this area approached the same question using intra-day data. The availability of high-frequency data helped mitigate the identification issue by increasing the statistical power of the estimates.

A strand of work in this area focused on the effects of monetary policy and started measuring monetary policy shocks (i.e. *surprises*) directly from changes in the prices of federal funds futures around interest rate announcements, that is Federal Open Market Committee (FOMC) meetings. This technique was firstly proposed by Kuttner (2001), who documented the response of returns on bills and bonds to monetary policy announcements on days on which the FOMC changed its policy target. The analysis was subsequently extended to equity returns in Bernanke and Kuttner (2005) and to macroeconomic variables, such as output and prices, in Faust, Rogers, Swanson and Wright (2003) and Faust, Swanson and Wright (2004). Bernanke and Kuttner (2005) established a strong reaction of equity prices to surprise funds rate changes showing that a broad index of U.S. stocks experience an average daily increase of 1% in the aftermath of an unanticipated 25-basis-point cut in the target rate.

[11] McQueen and Roley (1993) rely on their own definition of business conditions. They mainly use the seasonally adjusted monthly industrial production index to define economic states.

Rigobon (2003) and Rigobon and Sack (2003, 2004) proposed an alternate identification scheme. They exploited heteroscedasticity in high frequency data, that is the increase in the variance of shocks around monetary policy announcements, to derive identifying restrictions on the effects of monetary policy. This methodology yields a larger negative impact on stock prices and a smaller positive impact on the yield curve than the so-called "event study" approach pointing to biases in the inference drawn when relying on that technique.

A third methodology, firstly proposed by Gürkaynak et al. (2005a) and Gürkaynak (2005), derives indirect measures of monetary policy shocks from financial market reactions. These papers rely on a principal components technique to identify two common factors that successfully describe asset prices dynamics around FOMC announcements. Most research in this area focuses on the United States, but recent contributions have turned the attention to non-U.S. central banks. For example, Brand and Turunen (2010) implement the methodology by Gürkaynak et al. (2005a) to the euro area yield curve and document a significant impact of announcements made by the European Central Bank (ECB) on medium and long-term interest rates.

As a fourth approach, Rosa (2008) estimates forecasting regressions to approximate the unexpected components of both FOMC and ECB announcements. He uses two predictors – the previous statement and the slope of the yield curve prior to the release of the statement –, and documents stronger market reactions to unexpected Federal Reserve announcements than to unexpected ECB announcements.

Finally, the recent adoption of unconventional monetary policies poses further challenges to the correct identification of monetary policy shocks. The approach proposed by Kuttner (2001), for example, relies on the assumption that futures prices reflect market expectations of future policy rates. This technique is uninformative in the absence of a clear measure of the central banks' policy stance, as in the current context of zero (or nearly zero) short-term interest rates. To circumvent this issue, Cahill, D'Amico, Li and Sears (2013) and Joyce, Lasaosa, Stevens and Tong (2011) estimate the surprise component of asset purchase announcements using survey expectations, while Rogers, Scotti and Wright (2014) define the monetary policy

surprise as the intraday change in government bond yields around the announcement. The latter measure captures spillover effects from a change in government bond yields (caused by monetary policy) onto asset prices under study, but cannot evaluate the direct impact of the monetary policy on government bond yields. Rogers, Scotti and Wright (2014) examine the effects of unconventional monetary policy implemented by the Federal Reserve, Bank of England, European Central Bank, and Bank of Japan on bond yields, stock prices and exchange rates and document a substantial pass-through with the caveaut that Japanese surprises exhibit the weakest effect.

Another set of contributions in this area documented that macroeconomic news explain a substantial fraction of price volatility immediately after the announcement. Ederington and Lee (1993) established this finding in foreign currency and U.S. treasury bond futures markets building on previous work by Harvey (1991) and Savanayana, Schneeweis and Yau (1992), respectively. Fleming and Remolona (1997) focused on the entire spectrum of maturities in the secondary market for U.S. Treasury securities, and Balduzzi, Elton and Green (2001) extended these studies in both cross-section and time-series validating these findings in a comprehensive set of macroeconomic variables over a longer sample period. Further contributions assessed the impact of macroeconomic and monetary announcement surprises on returns themselves rather than on return volatilities. Conditional mean jumps are documented in foreign exchange markets Andersen, Bollerslev, Diebold and Vega (2003), and also jointly in global stock, bond and foreign exchange markets (Andersen, Bollerslev, Diebold and Vega, 2007).

A growing body of studies has established the strong reaction of bond yields to U.S. scheduled macroeconomic announcements at the intra-day frequency (Fleming and Remolona, 1997, 1999, 2001; Green, 2004; Andersen, Bollerslev, Diebold and Vega, 2007; Pasquariello and Vega, 2007; Faust, Rogers, Wang and Wright, 2007; Beber and Brandt, 2008; Jiang, Lo and Verdelhan, 2011). Overall, these contributions uncover that bond movements tend to be most sensitive to employment news (i.e. releases of non-farm payrolls number in the U.S. Employment Report), and at intermediate maturity. Recent high-frequency work by Hördahl, Remolona

and Valente (2015) shows that not only bond yields, but also bond risk premia sharply respond to revisions to expectations about short-term interest rates. For a survey of the literature on announcement effects on foreign exchange volatility and foreign exchange returns, see Neely (2011) and Neely and Dey (2010), respectively.

A distinct strand of the literature has examined the effect of pre-scheduled announcements on expected returns. In the asset pricing terminology, these studies look at risk premia rather than factor betas (i.e. asset price sensitivity). Relative to the contributions above, the goal shifts from exploring the impact of monetary policy and/or macroeconomic announcements on asset prices to investigating whether investors demand a premium for bearing macroeconomic risk. Since the timing of pre-scheduled announcements is publicly available, if asset prices respond to news and investors care about those news, a premium should arise in equilibrium to compensate rational investors to hold securities on announcement days. Asset returns should therefore be predictably higher on announcement days, as rational investors anticipate the possibility of sharp asset price fluctuations around releases.

Early studies (Engle, Lilien and Robins, 1987; French, Schwert and Stambaugh, 1987) relied on time-varying conditional volatility models (i.e., ARCH-in-mean models) to estimate time-varying risk premia and found that expected returns are higher in periods of high ex-ante volatility. However, inference drawn from this approach is prone to model misspecification issues and thus not reliable (Pagan, 1991). Jones, Lamont and Lumsdaine (1998) were the first to propose pre-scheduled announcements as identification scheme in this context. The strategy simply prescribes to test whether expected returns are higher on announcement days relative to non-announcement days. This paper found that both average long-term Treasury bond returns and their volatilities are be higher on inflation and employment announcement days relative to all other days. Savor and Wilson (2013) adopted a similar methodology in equity markets and documented that in the 1958-2009 period the average excess daily return on a broad index of U.S. stocks (i.e., CRSP index) is 11.4 basis points on announcement days (FOMC, employment and inflation releases), but only 1.1 basis points on all other trading days.

These findings are related to a separate body of work investigating the earnings-announcement premium, that is the well documented phenomenon that stock returns are on average higher on days when earnings (Beaver, 1968; Chari et al., 1988; Ball and Kothari, 1991; Cohen et al., 2007; Frazzini and Lamont, 2006; Savor and Wilson, 2016) or dividends Kalay and Loewenstei (1985) are scheduled for release in the United States. These premia do not appear to be empirically explained by increases in systematic risk. On the contrary, Savor and Wilson (2014) document a robust risk-return trade-off in the United States on macroeconomic (employment and inflation) and interest rate announcement days. U.S. stock market, bond and currency returns are all found to positively load on the aggregate market return on those days, indicating that market beta is an important measure of systematic risk. Instead, the relationship between average returns and the market beta is negative on non-announcement days.

A growing body of asset pricing work has recently focused on scheduled monetary policy announcements in the United States, that is interest rate decisions announced after FOMC meetings. Lucca and Moench (2015) document that the high average U.S. equity returns earned on FOMC days are mainly collected ahead of FOMC announcements and show that major international equity indices exhibit similar patterns as well. In the 1994-2011 period, the S&P500 index increased on average 49 basis points in the 24 hours before FOMC announcements. This evidence is peculiar to U.S. monetary policy: other macroeconomic announcements do not generate similar patterns. This pervasive phenomenon – the *pre-FOMC announcement drift* – is interpreted as a puzzle, as the authors fail to rationalise it in a standard risk-based framework. On the contrary, Mueller, Tahbaz-Saleni and Vedolin (2015) establish both a pre- and a post- FOMC announcement premium in currency markets. They show that currency excess returns are especially high (i) for carry trade currencies, (ii) in periods when monetary policy uncertainty is high, and (iii) when the Federal Reserve implements a policy of monetary easing. This phenomenon is interpreted as compensation for monetary policy uncertainty. Finally, Cieslak, Morse and Vissing-Jorgensen (2015) document that the FOMC premium does not simply

pertain FOMC announcement days, but it is a broader phenomenon. Their analysis reveals that since 1994 the U.S. equity premium is earned entirely in even weeks (i.e., 0, 2,4, and 6) in FOMC cycle time, that is the window between a FOMC meeting and the following one. This phenomenon pertains other asset classes as well: it is substantial in broad indices of international stocks (stronger for emerging than developed markets), weak for U.S. Treasury bonds, robust in currency carry trade portfolios of G10 economies, but absent in dollar portfolios. These patterns are interpreted as a risk premium for news coming from the Federal Reserve via a systematic informal communication mechanism.

More generally, a wide literature in Economics has investigated the role and importance of central banks and the optimal conduct of monetary policy. A growing body of work over the last two decades has focused in particular on central bank communication, due to its increasingly importance as a powerful tool to move asset prices and to anchor expectations. Blinder, Ehrmann, Fratzscher, De Haan and Jansen (2008) survey the contributions in this area discussing, among others, the analogies and differences in the communication policies implemented by the Federal Reserve System, the Bank of England, and the European Central Bank, and reviewing recent studies on the effects of central bank communication on the volatility of financial assets (Kohn and Sack, 2004; Connolly and Kohler, 2004; Reeves and Sawicki, 2007) and those attempting to quantify the direction and magnitude of these effects on asset prices (Jansen and De Haan, 2005; Ehrmann and Fratzscher, 2007; Gerlach, 2007; Rosa and Verga, 2007; Berge et al., 2006). In a related study, Woodford, Hole, Bean and Bernanke (2005) reviews arguments for and against central bank transparency and concludes that the predictability of monetary policies has been enhanced over the last two decades thanks to a worldwide tendency to greater transparency and increased communication.

# 3 Human Capital and Asset Prices

Human capital, broadly defined as the set of knowledge, skills, health, and values that contribute to making people productive (Becker, 1964), has been the object

of extensive strands of academic research in both Economics and Finance. This concept, pioneered by Gary Becker, Jacob Mincer, and Theodore Schultz back in the 1960s, has been examined by hundreds of researchers spanning the fields of financial economics, microeconomics, and macroeconomics, making hard to do justice to this broad literature in few pages. Without pretending to be exhaustive, the review below crosses the three fields, narrowing the coverage to the contributions and research areas that are mostly related to the third paper of this thesis (Chapter 5).

The importance of embedding human wealth, the present discounted value of future labour income, in asset pricing work dates back to Mayers (1972) and Roll (1977). In theory, both tradable (e.g., stocks and bonds) and nontradable (e.g., human capital) assets should enter the optimal allocation problem of rational agents, who determine optimal savings and consumption choices by maximising the expected utility of their lifetime consumption. In this view, the traditional Capital Asset Pricing Model (CAPM) of Sharpe (1964) and Lintner (1965) is a misspecified model, being the market portfolio simply a proxy for financial wealth. Mayers (1972), who was the first to point out the importance of the human wealth share in the aggregate wealth of advanced economies, derived an extended version of the CAPM allowing for human capital. Roll (1977) attributed the poor performance of the CAPM in explaining cross-sectional differences in asset returns to measurement errors in the market portfolio (i.e., *Roll's critique*), thus interpreting human wealth as a relevant, but omitted, variable. The non-negligible magnitude of this component was later confirmed by Heaton and Lucas (2000), who found that about 48% of household wealth is due to human capital, Jorgenson and Fraumeni (1989), who refined this estimate to be 90%, and Lustig, Van Nieuwerburgh and Verdelhan (2013), who used a no-arbitrage model to estimate the prices of aggregate risk from bond yields and stock returns and calculated a 92% share as well.

The first test of the Mayers (1972)'s model did not find empirical support for the inclusion of a labor factor into the CAPM (Fama and Schwert, 1977). Empirical asset pricing papers have subsequently proposed aggregate proxies of human capital

facing the challenge that the cash-flow component of any non-tradable asset is observable, but the discount rate component is not.

An early wave of studies overcame this issue by imposing ad-hoc restrictions on the conditional expected human return. The human discount rate was assumed to be constant (Shiller, 1995), equal to the conditional expected return on financial wealth (Campbell, 1996), constant alongside unforecastable labour income growth (Jagannathan and Wang, 1996)[12]. Augmented versions of the CAPM embedding any of these measures were shown to outperform the CAPM itself confirming the validity of the Roll (1977)'s critique. In this vein, Lettau and Ludvigson (2001a,b) used aggregate labour income to approximate the non-stationary component of human capital and uncovered that *cay*, the cointegrated vector of consumption, labour income, and asset holdings, has forecasting power for stock returns, and also improves the performance of conditional versions of both the CAPM and the consumption CAPM.

In the context of general equilibrium models, Santos and Veronesi (2006) rationalise the ability of the labour income to consumption ratio to predict stock returns. They also document its power in explaining risk premia in the cross-section. Most recently, Palacios (2014) incorporates investment in human capital into a stochastic model with Duffie-Epstein preferences, and derives a three-factor asset pricing model that depends on the market portfolio, the capital share and the human investment.

A recent wave of empirical work has adopted the VAR methodology to back out human discount rates from aggregate consumption data conditioning on a set of standard predictors for consumption growth and stock returns (Lustig and Van Nieuwerburgh, 2007; Bansal, Kiku, Shaliastovich and Yaron, 2014). This approach relies on the implicit assumption that the true information set is known to the econometrician, a condition that fails to hold in practice. Working in this

[12]The human capital proxy derived by Jagannathan and Wang (1996) is aggregate labour income growth. Recently, Eiling (2013) introduces human capital heterogeneity in the same framework by estimating a model that includes labour income growth rates from five broad industries (i.e., goods producing, manufacturing, distribution, services, and the government). These proxies are intended to span aggregate labour income.

context, Lustig and Van Nieuwerburgh (2007) reported two findings. Firstly, they showed that major human capital measures proposed in early studies (Shiller, 1995; Jagannathan and Wang, 1996; Campbell, 1996) are inconsistent with consumption data, suggesting they are misspecified proxies of human capital. Secondly, they found that returns on human and financial wealth must be negatively correlated to generate a sensible consumption path. This puzzling finding was subsequently resolved by Bansal et al. (2014), who uncovered a positive correlation between these two sources of wealth by introducing the assumption of conditional heteroskedasticity between the stochastic discount factor and the consumption asset return.[13] To incorporate this condition in the empirical analysis, they augmented the set of VAR predictors to include a measure of realised volatility, which is interpreted as a proxy for macroeconomic uncertainty.[14]

The findings of Bansal, Kiku, Shaliastovich and Yaron (2014) find also support in recent work in labour macroeconomics (Hall, 2014), stemming from very different premises. According to the Mortensen-Pissarides (1994) model, the leading theoretical framework for labour market dynamics, fluctuations in unemployment are driven by the searching cost incurred by firms when posting vacancies. When the incentive for job creation falls, firms destinate fewer resources into the hiring process, and unemployment rises. In the original paper, and several successor ones, the driving force of this mechanism is a drop in productivity, but this assumption is firmly rejected by the data. Shimer (2005) shows that any realistic calibration of the search and matching model generates tiny fluctuations relative to the observed unemployment dynamics (i.e., *Shimer puzzle*). Incorporating frictions like sticky prices (Wash, 2003), sticky wages (Gertler, Sala and Trigari, 2008; Gertler and Trigari, 2009) or unemployment insurance (Nakajima, 2012) mitigates the issue, but does not solve the puzzle. Hall (2014) points to the discount rate as the key driving

[13] In Lustig and Van Nieuwerburgh (2007) the stochastic discount factor and the consumption asset return are conditionally homoskedastic.

[14] In parallel work, Campbell, Giglio, Polk and Turley (2012) derive an intertemporal CAPM with stochastic volatility and test it using the VAR methodology. They show that volatility risk help explain the cross-section of value and growth stocks and also small and large stocks under a reasonable calibration of the parameter of risk aversion.

force behind unemployment fluctuations showing that unemployment is high when the discount rate of the aggregate labour market is high. He derives a proxy of the job value, the present discounted value of a future stream of benefits for a firm from a new hire, and documents that discount rates in the aggregate financial market and the aggregate labour market are positively correlated in the United States. Under the assumption that workers take hiring/firing decisions as given, this evidence can be interpreted as the mirror image of the findings in Bansal, Kiku, Shaliastovich and Yaron (2014) and unemployment be an omitted state variable driving their results. From a macroeconomic perspective, Hall (2014) concludes that financial frictions are likely a missing channel in the traditional search and matching framework.

Two broad strands of literature in economics have relied on micro-level Current Population Survey (CPS) data to examine relative differences in earnings (i.e. human capital cash flows) in the United States, conditioning on demographic characteristics (i.e., sex, race, age), the educational attainment (i.e., no high school, diploma, college, and post-college degree), and expanding this set to any other observable variable capturing heterogeneity at individual level (e.g., years of work experience, occupation). These studies have two distinct goals. Specifically, the first strand aim to document trends in the U.S. wage structure over time, while the second strand seeks to measure casual effects of education on earnings (i.e. *returns to schooling*).

In the former area, researchers have uncovered a sharp rise in income inequality since 1980 (see Piketty and Saez (2003), among others), but have not reached consensus yet on the forces driving this phenomenon and its persistence. A large set of works documents a pervasive widening in earnings during the 1980s (Bound and Johnson, 1992; Katz and Murphy, 1992; Murphy and Welch, 1992; Juhn, Murphy and Pierce, 1993), both within and across groups. This strand attributes the trend to shifts in the supply of and demand for skills commenced decades earlier, coupled with milder protection of low-and middle-wage earnings via changes in labour market institutions (i.e., labour unions, minimum wage). These factors are also interpreted as persistent drivers of the widening wage differentials over time. With specific respect to education, the pattern is a substantial increase in earnings

differentials, especially for college holders relative to less educated workers, in the 1980s and again in the 1990s. This traditional view is challenged by a more recent wave of studies, known as *the revisionists*. These works attribute the dynamics to a fall in the real minimum wage in the early 1980s – a major one-off episode – and see the subsequent trends as the result of mechanical changes in the labour force composition (Card and DiNardo, 2002). Autor, Katz and Kearney (2008) provide a detailed review of this debate alongside recent empirical evidence in favour of the traditional view (see also contributions therein).

In terms of methodology, researchers in this broad area aggregate wage data in two steps. Firstly, they average wages by bundling together individuals with similar characteristics. Next, they take fixed-weighted averages of these means to obtain more aggregated estimates.

In the second area, namely the literature on the returns to education, labour micro-economists have extensively attempted to quantify the profitability of investment in education at individual level – the private returns to schooling – measuring causal effects of education on earnings.[15] Researchers have adopted a plethora of econometric techniques over time, aiming to improve the precision of the estimates and address the identification issue. Card (2001) and, more recently, Heckman, Lochner and Todd (2006) review econometric issues in this field, while Harmon, Oosterbeek and Walker (2003) survey the compelling evidence outlining the arguments for the empirical developments in this area.

Regardless the methodology used, the labour literature defines the rate of return to education as a *wage premia*, that is the marginal difference in earnings yielded by two alternate educational investments (e.g., college degree vs. diploma).[16] This definition has foundation in the human capital theory of Becker (1964), who derived the optimal investment in education by maximising the expected present value of

[15] Macro-economic studies quantify the social returns to schooling, that is the impact on national economic growth, and investigate whether the effects are large enough to justify public support to education. See Sianesi and Van Reenen (2003) for a review of the contributions in this area.

[16] Educational investments were initially measured in number of years of schooling, but researchers have subsequently argued that credentials, that is the degree obtained, are more relevant than years of schooling. This is known as the *sheepskin* effect, as diploma used to be written on sheepskin. See Hungerford and Solon (1987) and references therein.

the lifetime stream of wages earned by a representative investor.[17] As emphasised by Becker himself later on (Becker, 1994), this measure has the drawback of not accounting for the riskiness of the investment in education, that is the variance in its rate of return. In other words, wage premia are not a risk-adjusted measure of human capital returns, as they simply amount to a comparison of average levels, disregarding the role of risk.

Although valid and relevant, Becker's point has remained ignored for another decade. Palacios-Huerta (2003) is the first attempt to compare the risk-return trade off per unit of risk that is intrinsic to different financial and educational investments. Using the proportional increase in earnings per year associated with the last unit of education as human capital return measure, this study uncovered substantial differences in marginal risk-adjusted rates of returns between financial and human capital assets. In particular, the risk-adjusted human return earned by a representative college holder is found to be superior to its financial counterpart, while the opposite holds for a representative high school holder.

Recently, Brown, Fang and Gomes (2012) derive a risk-adjusted measure of returns to education from a life-cycle model. In analogy with financial assets, this study defines the price of human capital as the present discounted value of a future stream of wages. Expected earnings are estimated using U.S. micro-level panel data (i.e., the Panel Study of Income Dynamics), while the human discount rate is derived from a structural life-cycle model of consumption and savings decisions in the tradition of Carroll (1997). Capturing risk along several dimensions (i.e. rich stochastic labor income process, unemployment probability, unemployment insurance, a progressive income tax system, and a progressive Social Security system dispensing benefits in retirement), the authors find substantial (risk-adjusted) educational gains for diploma holders as in the wage premia literature. However, they also uncover decreasing, rather than increasing, returns to education for college holders in the recent period (1991-2007). This result stems from a substantial increase in earnings risk that offsets the increase in expected human returns.

[17]Wage premia were firstly estimated via Mincer(1974)-Becker(1964) regressions, that is regressing wages on the level of education, controlling for demographics.

Overall, life-cycle models examine age-earning profiles conditioning on education (i.e., no high school, diploma, and college degree), and focus on the portfolio choices of individual households over their lifetime. A substantial fraction of risk these models capture – and risk-adjusted human measures derived in this context as well – is therefore idiosyncratic rather than systematic. As the authors point out, this implies that the discount rate component of aggregate measures of human wealth – those relevant in the asset pricing field – should be lower. Benzoni and Chyruk (2015) survey theoretical and empirical contributions on the value and risk of human capital in the life-cycle earnings literature.

In theoretical models, risky human capital affects investors' portfolio decisions, as agents hedge their labour income risk by investing in financial markets (e.g., Bodie, Merton and Samuelson (1992); Danthine and Donaldson (2002); Santos and Veronesi (2006); Parlour and Walden (2011)). However, limited stock market participation is a well established fact. Mankiw and Zeldes (1991) and Brav, Constantinides and Geczy (2002), for example, show that workers who are less wealthy, less educated and more reliant on wage income as major source of wealth choose to not participate. Two recent papers find evidence consistent with the labour income hedging hypothesis using a panel data set of Swedish households. Betermier, Jansson, Parlour and Walden (2012) find that households adjust their portfolio holdings when switching jobs, and Betermier, Calvet and Sodini (2016) find that investors with high human capital and high exposure to macroeconomic risk tilt their portfolios away from value stocks. Berk and Walden (2013) rationalise limited stock market participation as an equilibrium outcome by introducing the labour market in the neoclassical asset pricing model. In equilibrium, less flexible workers (agents who do not switch jobs) hedge their labour income risk in the labour market and optimally choose to not participate in the stock market. With respect to the optimal portfolio allocation, the life-cycle model studied by Brown, Fang and Gomes (2012), as the one in Cocco, Gomes and Maenhout (2005), predicts that all education groups invest a similar share of savings in the stock market.

As a final note on the role of unemployment as a determinant of human wealth, aggregate human capital measures account for this component in level, as aggregate labour income data include unemployment benefits. In life-cycle models, unemployment enters as a transitory perturbation of the labour income stochastic process of an individual over the lifetime. For instance, Carroll (1997) introduced *background risk* into the life-cycle framework, by allowing labour income to be zero with positive probability over the life-cycle.[18] Brown, Fang and Gomes (2012) explicitly work with unemployment probabilities, assuming that each year agents have a probability of suffering an unemployed spell.

The three empirical papers which follow contribute to this wide existing literature on international asset pricing, the relation between announcements and asset prices, and the price and value of human capital, respectively.

[18] *Background risk* is defined as the possibility of a disastrous labour income shock over the life-cycle.

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# 3 | The International CAPM Redux*

## 1 Introduction

Are international equity investors compensated for bearing exchange rate risk? This question is increasingly relevant in light of the rapid acceleration of international financial integration over the past three decades. Figure 1 shows that aggregate foreign equity holdings as a percentage of global gross domestic product have increased steadily from roughly 3% in the 1980s to approximately 30% in 2011, alongside a steady dismantling of de-jure cross-border capital flow restrictions. The magnitudes are large, highlighting the importance of this issue – in the United States, for example, foreign equity holdings are currently worth roughly US$ 6 trillion.

In theory, exchange rate risk should matter to the pricing of equities and other risky assets in a world of real rigidities and deviations from purchasing power parity. These deviations affect the consumption of international investors, who invest abroad, but consume at home. In most plausible theoretical international asset pricing models, this effect on consumption leads to equilibrium compensation for the risk of low returns on foreign investments once these returns are expressed in real domestic terms. Reasonable though this rationale is, empirical work in international

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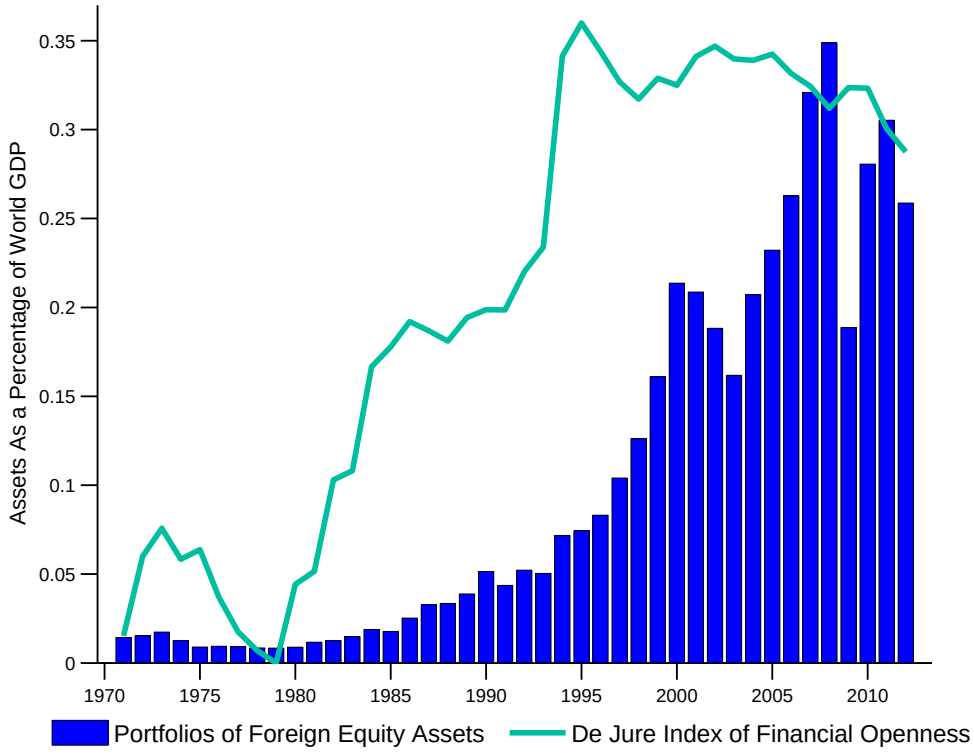
Figure 1
Portfolios of Foreign Equity Assets and Financial Openness

This figure shows the portfolio equity assets (bar plots) for 39 developed and emerging countries along with a de jure financial openness index. In each year, portfolio assets of all countries are summed up and divided by the corresponding world GDP. The financial openness index is based on the restrictions in capital flows listed by the IMF over time; a world index is obtained by GDP-weighting country indices. Estimates of foreign assets come from Lane and Milesi-Ferretti (2007). The financial openness index comes from Chinn and Ito (2008). The set of developed countries comprises Australia, Austria, Belgium, Canada, Hong Kong, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, India, Indonesia, Ireland, Italy, Japan, Kuwait, Malaysia, Mexico, Netherlands, New Zealand, Norway, Philippines, Poland, Portugal, Saudi Arabia, Singapore, South Africa, South Korea, Spain, Sweden, Switzerland, Taiwan, Thailand, Turkey, United Arab Emirates, United Kingdom, and United States. The data are annual and the sample period is 1971 to 2011.



asset pricing, with a few notable exceptions, has not been able to provide convincing evidence that currency risk is priced in international equity markets.

In this paper, we present a new three-factor model to capture the risks in international equity portfolios. These three factors are the return on a world market portfolio denominated in local currency terms, and two currency factors which effectively summarize variation in a broad cross-section of bilateral exchange rates, namely, the dollar factor and the carry factor.[19]

To test the model, we compile a comprehensive dataset of equity returns spanning value, size, and country aggregate portfolio returns from 25 developed markets and 21 emerging markets between February 1976 and April 2013. When estimated on these data, the model delivers an important role for currency risk in the cross-section of equity returns.

Importantly, in conditional asset pricing tests, our model outperforms the single-global factor World CAPM as well as the International CAPM estimated by Dumas and Solnik (1995), and delivers comparable performance to the global versions of the popular Fama and French (2012) factors. We show that our model is also relevant to those interested in delegated portfolios of international assets. It is often difficult to infer the amount of currency risk taken on by international investment managers, and our factors offer a simple metric of the extent of this issue, helping to explain a significant fraction of the variation in international mutual and hedge fund returns.

To better understand the dynamics captured by the empirical risk factors, we build a simple theoretical model in which exchange rates, currency risk factors, and equity market returns are all precisely defined. In leading international equity asset pricing models, exchange rate variation arises exogenously. In contrast, we assume that exchange rates are driven by the same shocks that drive the pricing kernels. To better explore the impacts of this important difference, we simplify all other aspects

[19]Verdelhan (2014) shows that a substantial fraction of the variation in bilateral exchange rates can be captured by these two factors. The dollar factor is the average excess return earned by an investor that borrows in the U.S. and invests in a broad portfolio of foreign currencies. The carry factor is the average excess return earned by an investor that goes short (long) in a portfolio of low (high) interest rate currencies.

of the model, assuming that financial markets are complete, and writing down the law of motion of the lognormal stochastic discount factors (SDFs) in all countries.

In our setup, country SDFs depend on country-specific shocks, as well as three global shocks, and each country's aggregate dividend growth rate also depends on the same set of shocks. To create a role for a pure equity risk factor (which affects all countries in the same way), we allow one of the global shocks to affect all SDFs similarly. Since changes in exchange rates are differences in log SDFs, this global shock does not show up in currency markets, resulting in a degree of segmentation between equity and currency markets. This simple innovation to the basic complete markets model, namely, the addition of the "equity" global shock, allows us to work in a very tractable framework in which (realized and expected) equity and currency returns can be written down in closed form.

The model reveals that the world aggregate equity return expressed in a common currency actually does contain all relevant information necessary for pricing international assets, meaning that there should be no role for bilateral exchange rates. The twist is that time-variation in the prices of global shocks – necessary to account for time-varying currency risk premia – confounds empirical estimation using a single-factor model, especially when the relevant state variables are unknown to the econometrician. In other words, a conditional version of the world CAPM holds in the model, but in practice the conditioning variables are unknown. We show that including the currency risk factors in our three-factor empirical model arises as a natural solution to the challenges of empirical estimation in this framework.

We use the model to shed light on our main empirical results, calibrating it to match a large set of equity and currency moments, and generating simulated data. We apply exactly the same estimation procedure as we do on real data to these simulated data, and replicate our empirical finding that the new "International CAPM Redux" outperforms both the World CAPM of Sharpe (1964) and Lintner (1965), and the "Classic" International CAPM of Adler and Dumas (1983). We note that time-variation in the prices of risk appears key to this result.

Finally, we use the model to revisit a longstanding and important issue in international finance, namely, optimal currency hedging in international portfolios.[20] The prevailing wisdom arises from the mean-variance efficient framework of Black (1989), in which the optimal amount of currency hedging depends on the mean and standard deviation of world market returns, as well as on exchange rate volatility. In other words, Black (1989) shows that the optimal amount of currency hedging is constant through time and across investor location. In contrast, in our model, in which exchange rates are driven by many of the same risks affecting equities, the optimal amount of investment in currency portfolios such as carry and dollar is time-varying and investor-location-specific, and depends on the prices of risk of both country-specific and global shocks.

Our paper relates to two very large strands of literature on international equity markets and on currency risk. A short paragraph in the introduction of this paper would not do justice to this literature. Instead, we propose a four-page review of the most relevant work. In the interest of space, the material is placed at the start of the Appendix.

The paper is organized as follows. Section 2 shows the empirical specifications of the various international asset pricing models that we test. Section 3 describes the data on which we conduct these asset pricing tests, and Section 4 discusses the results. Section 5 presents our simple theoretical model which endogenizes exchange rates in complete markets and gives rise to the empirical specification of the International CAPM Redux. Section 6 shows how we can use the model to shed light on optimal hedging in a world with endogenous exchange rates, and Section 5 concludes. All robustness checks and additional results that are mentioned but not reported in the paper are provided in the Online Appendix.

[20]This continues to be a hot topic for both institutional and retail investors. For example, four Exchange-Traded Funds (ETFs) that seek to completely hedge the currency risk component of international equity returns are among the ten fastest growing ETFs over the last six months, collecting more than 30 billion dollars over this short period.

# 2 Specifying and Testing International Asset Pricing Models

We compare the performance of our three-factor model against a range of alternative asset pricing models: the World CAPM, the International CAPM, and a global version of the Fama-French-Carhart model. In this section, we briefly present the empirical specifications of these leading international asset pricing models and contrast them with our three-factor specification.

## 2.1 The World CAPM and the International CAPM

The World CAPM is a simple extension of the CAPM to global markets. The additional assumption necessary in the global context is that PPP always holds, meaning that currency risk is rendered irrelevant. In the World CAPM, global market risk is the single source of systematic risk driving asset prices, and international investors should only earn a premium for exposure to this source of risk. Empirically, the measure of global market risk has generally been the excess return on the world equity market portfolio, $WMKT$, denominated in a common currency, generally U.S. dollars. Writing $r^{\$}_{i,t+1} - r_{f,t}$ for the excess return of asset $i$ at time $t+1$ expressed in dollars, the empirical specification of this model is:

$$r^{\$}_{i,t+1} - r_{f,t} = \alpha^{i}_{WCAPM} + \beta^{i}_{WMKT} WMKT_{t+1} + \epsilon_{t+1}. \tag{2}$$

In the international CAPM of Adler and Dumas (1983), PPP does not hold instantaneously, and exchange rates are an additional source of exogenous risk. This model is operationalized in the empirical work of Dumas and Solnik (1995). With the same notation as before, and $r^{GBP}_{t+1}$, $r^{JPY}_{t+1}$, $r^{DEM}_{t+1}$ representing excess currency returns denominated in Pound Sterling, Japanese Yen, and Euro/Deutsche Mark, respectively, the (unconditional) model specification is:

$$\begin{aligned} r^{\$}_{i,t+1} - r_{f,t} &= \alpha^{i}_{ICAPM} + \beta^{i}_{WMKT} WMKT_{t+1} \\ &+ \beta^{i}_{GBP} r^{GBP}_{t+1} + \beta^{i}_{JPY} r^{JPY}_{t+1} + \beta^{i}_{DEM} r^{DEM}_{t+1} + \epsilon_{t+1}. \end{aligned} \tag{3}$$

## 2.2 Fama-French Global Factor Model

Since their discovery, the Fama and French (1993) three factors and the Carhart (1997) four factors, although not based on a particular theoretical model, have become standard benchmarks in empirical asset pricing. These models ignore exchange rate risk, and offer an explanation for patterns in international average stock returns based on loadings on size, value, and momentum premia:

$$\begin{aligned} r^{\$}_{i,t+1} - r_{f,t} &= \alpha^i_{FF} + \beta^i_{WMKT} WMKT_{t+1} \\ &+ \beta^i_{SMB} SMB_{t+1} + \beta^i_{HML} HML_{t+1} + \beta^i_{WML} WML_{t+1} + \epsilon_{t+1} \end{aligned} \quad (4)$$

where $WMKT$ is defined as before, $SMB$ is small minus big, capturing the size premium, $HML$ is high minus low (book-to-market), capturing the value premium, and $WML$ is (short-term) winners minus losers, capturing the effect of momentum.

## 2.3 The International CAPM Redux

We present now intuitively our three-factor specification. A proper derivation is presented in Section 5.

The International CAPM read literally recommends the use of all bilateral exchange rates as additional risk factors, which is somewhat cumbersome empirically. However, we know from recent research in currency markets that a large set of bilateral exchange rates can be summarized using carry and dollar factors, which are well able to capture systematic variation in bilateral exchange rates. In addition to these currency factors, one may need additional factors to summarize equity risk unrelated to currency risk.

This heuristic description is simply here to introduce our empirical model, and we spend a great deal of time rationalizing the model, following the discussion of our empirical results. For now, we describe our international CAPM Redux model simply as:

$$\begin{aligned} r^{\$}_{i,t+1} - r_{f,t} &= \alpha^i_{CAPMredux} + \beta^i_{LWMKT} LWMKT_{t+1} \\ &+ \beta^i_{Dollar} Dollar_{t+1} + \beta^i_{Carry} Carry_{t+1} + \epsilon_{t+1}, \end{aligned} \quad (5)$$

where $LWMKT_{t+1}$ denotes the excess return on the world market portfolio denominated in local currencies, and the construction of $Carry_{t+1}$ and $Dollar_{t+1}$ is described below.

## 2.4 Time-varying Quantities and Risk Prices

While we write all these models in their unconditional form, we estimate all of them conditionally, using rolling windows, to account for the possibility that betas and market prices of risk vary over time. Time-variation in the models' parameters is not a luxury, but a key feature of any international asset pricing exercise. To see this point clearly, let us assume that financial markets are complete.

When the law of one price holds on financial markets and investors can form portfolios freely, there exists a SDF $M_{t+1}$ that prices any return $R^i_{t+1}$ such that $E_t\left(M_{t+1}R^i_{t+1}\right) = 1$.[21] The same condition holds for the risk-free rate $R_f$. Assuming that the returns and SDF are lognormal, the Euler equation implies:

$$E_t\left(r^i_{t+1} - r_{f,t} + \frac{1}{2}var_t(r^i_{t+1})\right) = \underbrace{-\frac{cov_t(m_{t+1}, r^i_{t+1})}{var_t(m_{t+1})}}_{\beta^i_t}\underbrace{var_t(m_{t+1})}_{\Lambda_t} \tag{6}$$

where lower letters denote logs. Expected excess returns are the product of the quantity of risk, $\beta^i_t$, which is asset-specific, and the market price of risk, $\Lambda_t$.

It is well-known since Bekaert (1996) and Bansal (1997) that in a lognormal model in complete markets, the log currency risk premium equals half the difference between the conditional volatilities of the log domestic and foreign SDFs. Since currency risk premia are time-varying (as shown by the large literature on uncovered interest rate parity and the forward premium puzzle), log SDFs must be heteroskedastic.[22] That is, empirical estimation must (at least) account for time-varying market prices of risk ($\Lambda_t$).

With this set of empirical specifications in hand, we turn now to the data.

[21] The law of one price on financial markets implies that for any payoffs $X$ and $Y$, their prices satisfy $P(aX + bY) = aP(X) + bP(Y)$ for any real numbers $a$ and $b$.

[22] When SDFs and returns are not lognormal, a similar result implies that the higher moments of the SDF must be time-varying. Lustig and Verdelhan (2015) extend the Bekaert (1996) and Bansal (1997) result to incomplete markets: in the case of lognormal shocks, the expected currency risk premium depends additionally on the conditional volatility of the incomplete market wedge.

# 3 Data

This section describes our test assets and risk factors.

## 3.1 International Equity Returns

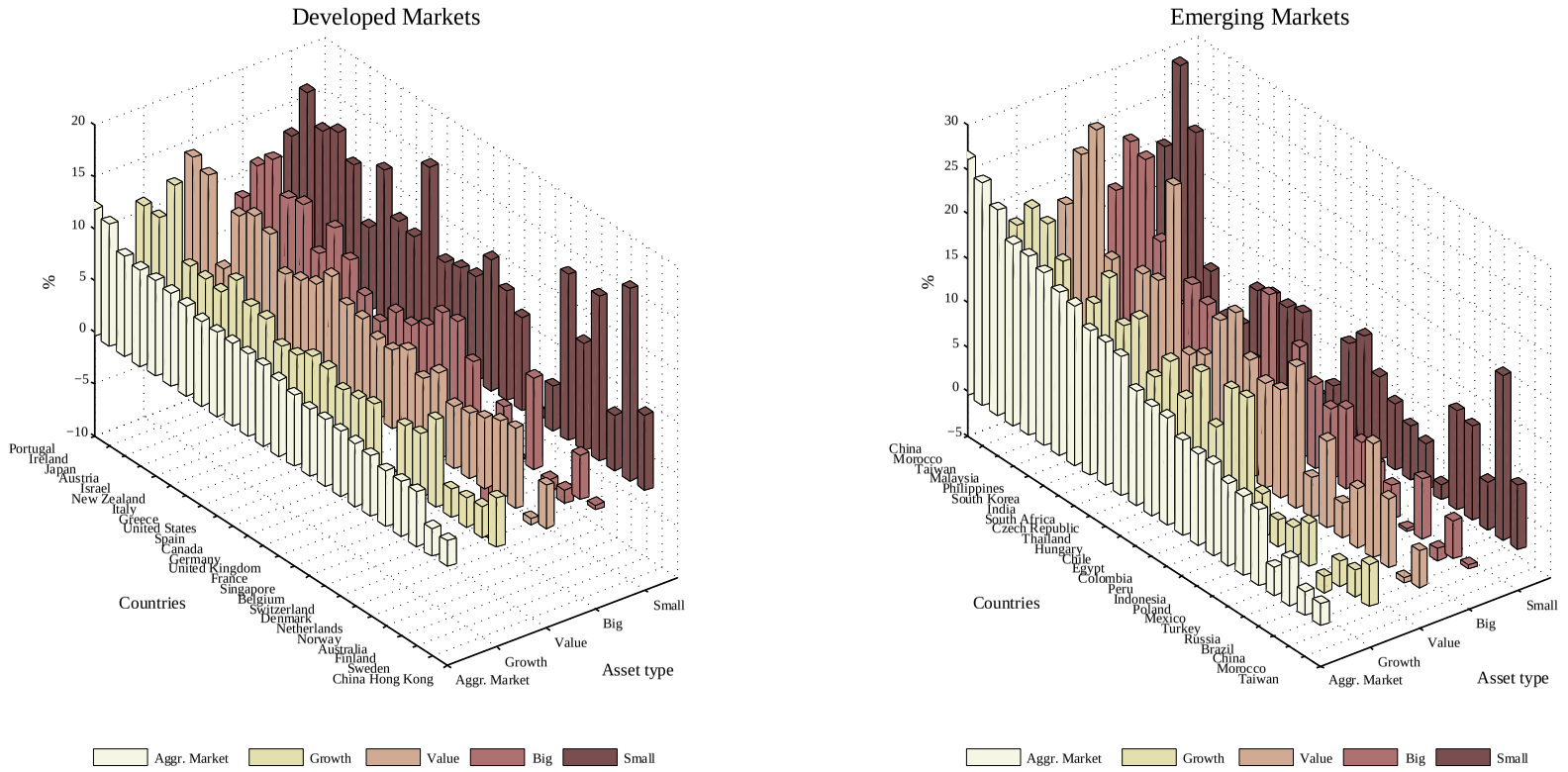
Our equity return data span 46 countries, comprising 225 different indices, over the period from January 1976 to April 2013. The coverage of countries follows the constituents of the 2011 Morgan Stanley Capital International (MSCI) Global Investable Market Indices. Following MSCI's approach, countries are classified into two categories, namely, 25 developed markets and 21 emerging markets. For each of these countries, Datastream reports daily total return series denominated in U.S. dollars for five different MSCI indices, namely, (i) the aggregate market, (ii) an index of growth stocks, (iii) an index of value stocks, (iv) an index of large market-capitalization stocks, and (v), an index of small market-capitalization stocks. Monthly returns are obtained from end-month to end-month. The risk-free rate is the U.S. 30-day Treasury bill rate, obtained from Kenneth French's website. Countries and asset types enter the equity data set at different points in time, depending on data availability. There are 54 test assets at the beginning of the sample in 1976, covering 18 developed markets and three asset types, namely, aggregate market, value, and growth. The size of the cross-section progressively increases from 1986 onwards.

The MSCI portfolios offer a challenging cross-section of returns to explain. Figure 2 provides a pictorial description of this fact – in the figure, for each asset type, countries are sorted according to their average market excess returns. The figure shows that these test assets exhibit large cross-sectional variation in average aggregate equity excess returns, across both developed and emerging countries, and across the different types of indices.

Our two global equity factors, $WMKT_{t+1}$ in U.S. dollars and $LWMKT_{t+1}$ in local currencies are monthly series, and we construct them using the daily MSCI World Index series denominated either in U.S. dollars or in local currrencies, obtained from Datastream. Monthly returns are computed from end-month to

Figure 2
Annualized Average Excess Returns by Country Group and Asset Type

This figure shows annualized average excess returns (in percentage) for Developed Markets (left plot) and Emerging Markets (right plot). Monthly returns are computed on five types of MSCI country equity indices (aggregate market, growth stock, value stock, big cap and small cap indices) and reported in excess of the U.S. one month Treasury bill rate. In both graphs, sorting is based on aggregate market excess returns: countries are ranked from the highest to the lowest value of those returns. Average excess returns are computed over different samples due to data availability. For data coverage refer to Table A.1. Daily indices are from Datastream; monthly returns are computed from end-of-month series; annualized returns are obtained multiplying monthly sample average excess returns by 12. The sample is February 1976 to April 2013.



end-month; excess returns simply subtract off the U.S. 30-day Treasury bill rate. The size, value, and momentum international Fama and French (2012) factors and the U.S. size, value and momentum Fama and French factors are obtained from Kenneth French's website. All these series are denominated in U.S. dollars. International series are available from July 1990, except for the momentum series that start in November 1990.

Our use of a world equity index built without any exchange rate data is to avoid any misattribution of currency risk to equities and vice versa. This "synthetic" local factor may be raise concerns about real-world implementability of our model. To address such concerns, we show that this factor can easily be replicated using existing mutual fund returns. The Appendix presents details about the construction of a factor mimicking portfolio (FMP) for $LWMKT$, and shows that this FMP is highly correlated with $LWMKT$, and delivers virtually identical returns. The Appendix also shows that the pricing errors from estimating the model using the FMP rather than $LWMKT$ are virtually identical to the ones we obtain in our benchmark results.

We turn next to describing our exchange rate data.

## 3.2 Exchange Rates, Carry, and Dollar

We obtain daily spot and one-month forward exchange rate series (midpoint quotes) quoted in British pounds for the same set of countries as above by merging data from Datastream, Reuters, Barclays, and additional sources. The Online Appendix describes these series in detail.

Assuming that the covered interest parity condition (CIP) holds, the difference between the (log) forward and spot exchange rates (i.e., the forward discount) is equal to the interest rate differential (in log-form) between the foreign and domestic nominal one-month risk-free rates. Countries enter the currency data set at different points in time according to the availability of their forward rate series. There are 15 currencies at the beginning of the sample in 1976 and 28 at

the end. The maximum monthly coverage is 34, as the Euro replaces national Euro area currencies from January 1999 onwards.

Following Lustig and Verdelhan (2005, 2007), at each time $t$, we create six currency portfolios by sorting all available currencies in our data set by their forward discounts. These portfolios are rebalanced at the end of each month. The excess returns on the carry factor in each month $t$, denoted $Carry_t$, are constructed as the difference between the returns on the top portfolio minus the return on the bottom portfolio constructed in this fashion. The average excess return earned by a U.S. investor on the carry trade strategy is 7.65%.

To construct the dollar excess return, we assume that in each period an investor borrows in the U.S. and invests in all other currencies in our data set. The excess returns on this strategy in each month are denoted by $Dollar_t$. The correlation between the dollar factor and the carry factor has historically been low – in our data set, it is 0.18 over the full sample period, but rises to 0.40 in the post-1990 period, driven primarily by the incidence of currency and financial crises in this latter period.

We also expand our set of test assets beyond equity markets by adding two sets of six currency portfolios, either sorting countries by their short-term interest rates (i.e., Carry portfolios) or by their dollar betas (i.e., Dollar portfolios). The construction of these portfolios is described in greater detail in Verdelhan (2015).

## 3.3 Mutual Fund and Hedge Fund Returns

In our empirical work, we also use our model to evaluate the exposure of international mutual funds and hedge funds to currency risk.

Monthly mutual fund data are from CRSP. The sample includes all funds classified as “Foreign Equity Funds" according to the CRSP fund style code. The sample period is 11/1990–4/2013, that is the horizon over which the Fama-French global factors are constructed. This choice ensures a large cross-section of mutual fund returns. At each time $t$ we compute returns only if CRSP provides the corresponding total net asset value (NAV) under management. There are 74 funds

with 5 years of past returns as of October 1995, but up to 1148 funds by the end of the sample, which collectively manage roughly US$ 800 billion.

Monthly hedge fund data are from the updated version of the consolidated hedge fund database built by Ramadorai (2013) and Patton, Ramadorai and Streatfield (2013).[23] The sample period for these data is 1/1994–4/2013. From this universe, we select funds with self-reported strategies falling into "Macro" or "Emerging" categories, and we do not select any funds-of-funds. The database includes fund returns net of management and incentive fees, and fund assets under management (AUM). All series are denominated in U.S. dollars. The number of hedge funds varies over time – with 85 funds at the beginning of the sample and 362 towards the end of the sample, which collectively manage roughly US$ 149 billion.[24]

We now turn to the results of our asset pricing tests.

# 4 Time-series and Cross-sectional Asset Pricing Tests

## 4.1 Time-series Tests

As described above, there are numerous reasons to expect that capturing the role of currency risk will require allowing for time-variation in the prices and quantities of risk. This importance of using a conditional model is also consistent with the findings of Dumas and Solnik (1995) when testing international asset pricing models.

Harvey (1991) and Dumas and Solnik (1995) capture time-variation in factor risk premia by conditioning on a set of instruments. As Cochrane (2001) notes, since the econometrician does not know the true set of state variables, conditional asset pricing tests are joint tests of the set of variables employed as conditioning information and whether the asset pricing model minimizes pricing errors.

[23]Ramadorai (2013) and Patton, Ramadorai and Streatfield (2013) consolidate data from the TASS, HFR, CISDM, Morningstar, and BarclayHedge databases.

[24]For mutual fund series, when NAV are annual or quarterly, we linearly interpolate monthly values. We adopt the same procedure for single missing observations, which we interpolate using the two adjacent values. Likewise, for hedge fund series, single missing observations of the AUM are linearly interpolated. We never interpolate returns.

Our principal approach in our asset pricing tests, therefore, is to estimate time-variation in factor loadings using simple rolling window regressions in the spirit of Lewellen and Nagel (2006). Following their implementation, we use 60-month rolling windows for our regressions. This choice means that the maximum number of rolling regressions we run for a single country is 388, and the minimum is 161 – this variation is a result of country-specific data availability. In the Online Appendix, we verify that our results are robust to the use of other window sizes (namely, 36- and 72-month windows).[25]

Figure 3 shows that there is indeed substantial time-variation in the factor betas estimated using our model across rolling windows. For each country in the data set and each risk factor in the model, the figures report the average rolling factor loading (the central dot in each figure), as well as the range between the minimum and the maximum estimated rolling factor loadings (the two ends of the line in each figure).[26] Across countries, especially for the currency factors, there are significant differences both in the magnitude of the risk exposures, and the degree to which they vary over time. While these features are evident in both sets of markets, they are more pronounced in emerging markets.

With few exceptions (primarily in developed countries), estimated carry loadings switch sign over time, and dollar betas are more volatile than carry betas. Extreme examples of time variation include the dollar loadings of Indonesia, and the carry loadings of Turkey. The Netherlands and the United States show the most stable exposures to currency risk. The heterogeneity across countries is so pronounced that it is difficult to identify common patterns. For example, while Japan and Switzerland are typical carry-trade funding countries, the carry factor loadings of their equity returns often move in opposite directions.

[25]Results from the unconditional estimations are omitted to save space, but intercepts $\hat{\alpha}$, adjusted R-squared statistics and beta estimates are available upon request.

[26]Country-by-country time-series of the betas, along with Newey and West (1987) standard error bands computed with the optimal number of lags according to Andrews (1991), are available upon request. They confirm that time-variation in these risk-exposures is not driven by outliers, and is usually statistically significant.

Figure 3
Time-Varying Factor Betas

This figure shows, for each country in the dataset, the interval between the minimum and the maximum value (black arrow) of a time-series of 60-month rolling factor betas along with their average values (white dot). Monthly equity excess returns are regressed on a constant, the global equity factor (LWMKT) and the dollar and carry factor over 60-month rolling windows. Test assets are excess returns on MSCI aggregate market indices denominated in U.S. dollars. The sample period is February 1976 to April 2013.

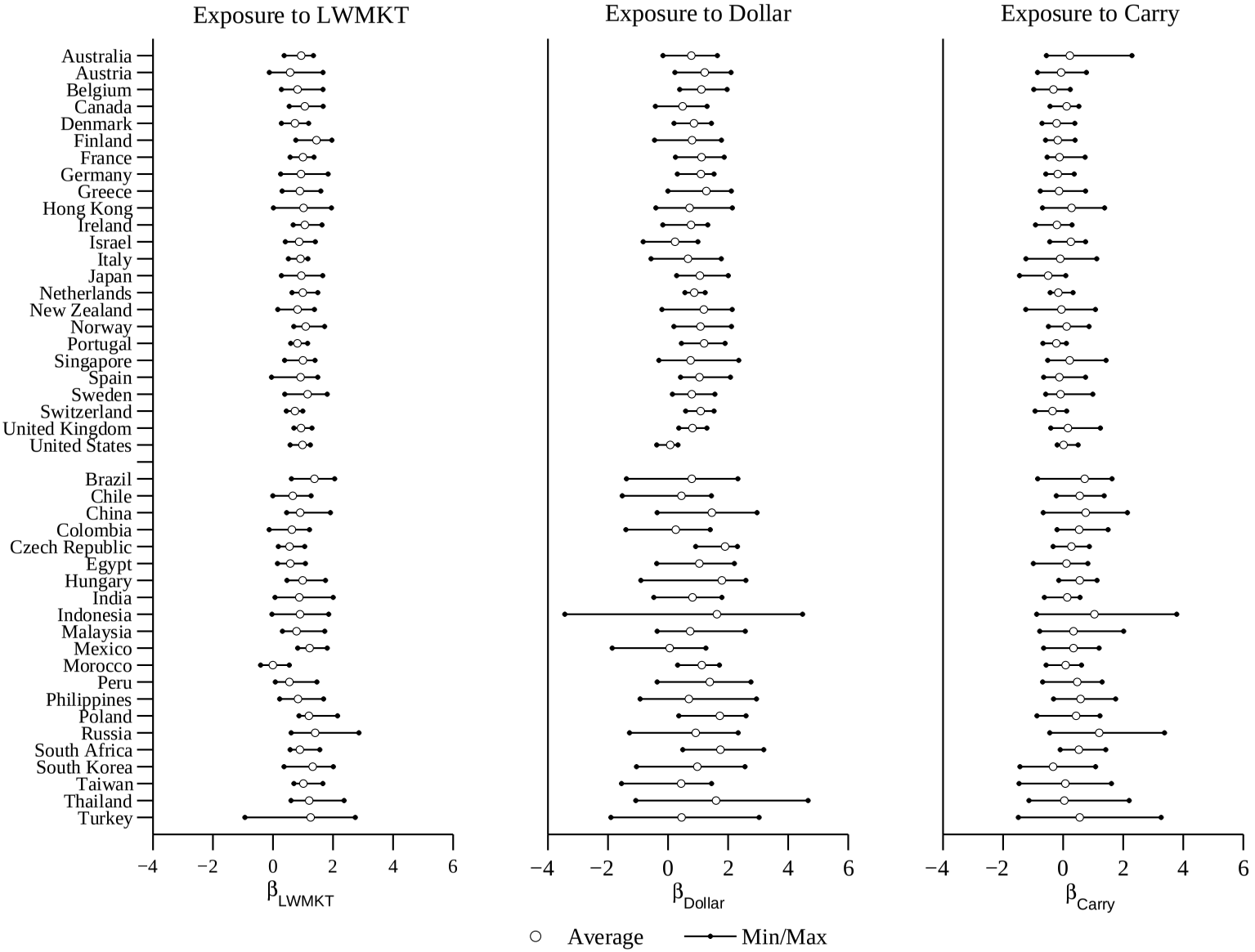


Figure 4
CAPM Redux: Significant Time-Varying Exposure to Global Factors

This figure reports the time-series of the share of test assets with significant exposure to the global equity and currency factors ($\beta_t^{**}$). Monthly equity excess returns are regressed on a constant and the global equity, dollar and carry factor over 60-month rolling windows. The size of the cross-section varies across time according to data coverage (# Test Assets). Test assets are equity excess returns on MSCI aggregate market, value, growth, small cap and big cap indices for developed (left graphs) and emerging (right graphs) markets. Standard errors are Newey-West. Statistical significance is tested at the 5% level. The sample period is February 1976 to April 2013.

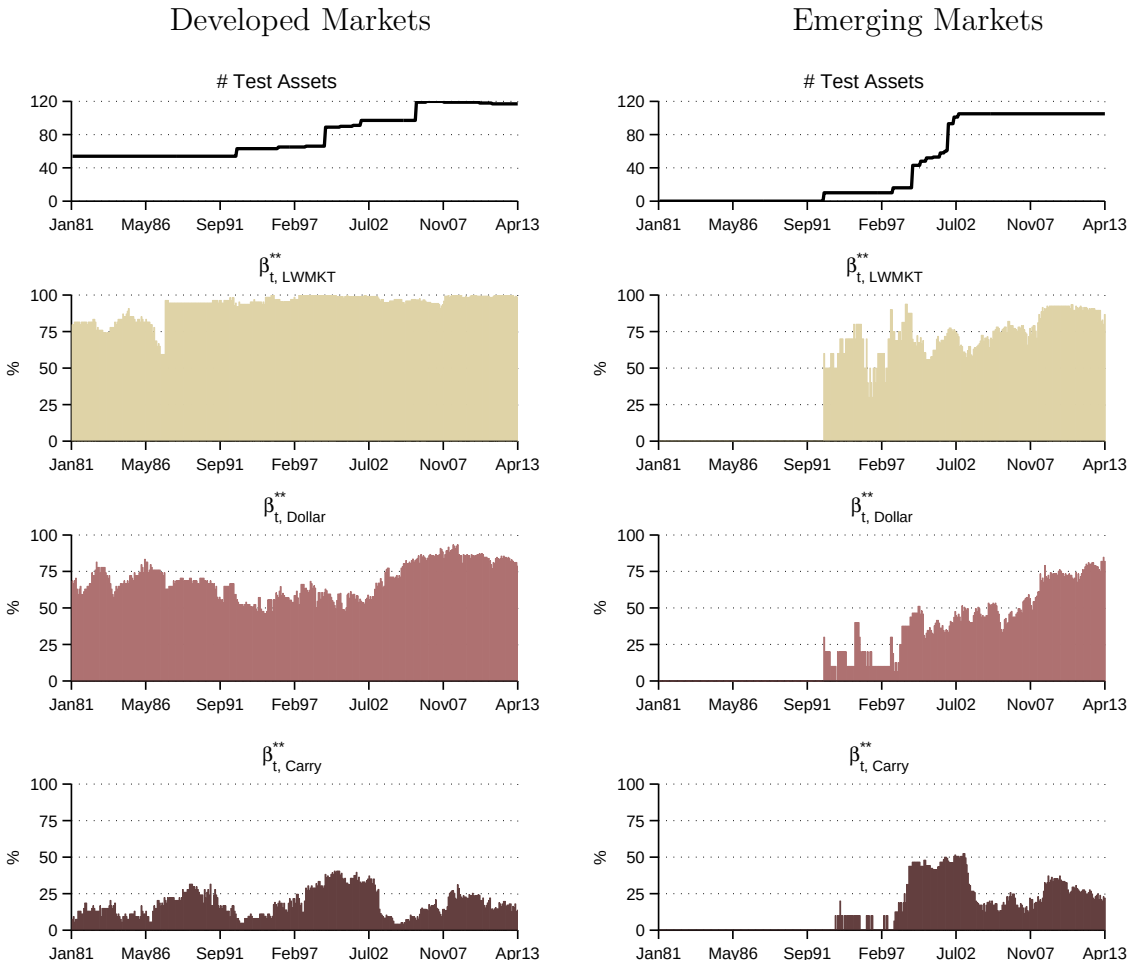



Figure 4 presents a high-level overview of the time-series of the estimated coefficients from our international CAPM Redux model. The figure shows results from 60-month rolling-window regressions of test asset returns on the factors in our new model. Developed markets are on the left- and emerging markets on the right-hand side of the figure.

The top panel in both columns shows the number of test assets in our asset pricing tests. Until late 1991, the set of test assets (including aggregate country indexes and value and size sorted equity portfolios) is restricted to those from the developed countries. Emerging countries begin entering the data in 1991, but really only constitute a significant fraction of the data from the late 1990s onwards. The set of assets continues expanding well into the 2000s, posing some challenges for time-series tests, since we suffer from large $N$, small $T$ problems in our unbalanced panel. We explain how we conduct GRS tests in this setting below.

The panels below show the percentage of betas from the model that are statistically significant in each rolling window. Despite the overlap between these windows and the mechanical persistence of these estimates over shorter periods, it is clear that there is considerable time-variation in the statistical significance of these beta estimates. The equity factor, $LWMKT$ is virtually always significant for the developed markets, and by the end of the sample period, for the emerging markets as well. The Dollar factor also shows considerable statistical significance for the developed markets, with over 50% of the set of test assets having statistically significant loadings on this factor, and has increasing significance as an explanatory variable for emerging market assets, reaching statistical significance in 75% of test asset regressions by the end of the sample. Finally, the Carry factor shows substantial time variation in its statistical significance, peaking after crises, with statistical significance seen for between 10% and 50% of test assets depending on the rolling window. The fact that these factors are statistically significant is important, especially in light of recent literature in asset pricing which casts doubt on second-stage results from standard two-pass cross-sectional asset pricing tests when first-stage betas are statistically insignificant (see, for example, Bryzgalova, 2015).

Table 1 shows the first comparison between our model and the competitor international asset pricing models, using the usual "GRS" $F$-test of Gibbons, Ross, and Shanken (1989) on each model to test whether all test-asset intercepts are jointly zero. The table compares the models by reporting the share of rolling windows in which the GRS test rejects the null hypothesis that rolling alphas implied by a

Table 1
Rolling GRS Tests

This table compares the World CAPM, the International CAPM, the Fama-French four-factor model (4FF) and the CAPM Redux by reporting the share of rolling windows in which the GRS test rejects the null hypothesis that rolling alphas implied by a given asset pricing model are jointly zero at the 5% level. Intercepts are estimated by regressing country-i equity excess returns on a constant and the appropriate set of factors over rolling-windows of 60 months. Under the assumption of normality, the GRS test statistic is F-distributed and defined as $\frac{T-N-K}{N}\left(1+E_T(f)'\hat{\Omega}^{-1}E_T(f)\right)^{-1}\hat{\alpha}'\hat{\Sigma}^{-1}\hat{\alpha}\sim F_{(N,T-N-K)}$, where T, N and K denote the sample size, the number of test assets, and the number of factors, respectively, $\hat{\Omega}$ is the variance-covariance matrix of the factors $f$ and $\hat{\Sigma}$ is the variance-covariance matrix of the estimated residuals. Five sets of test assets are considered (i.e., excess returns on MSCI aggregate market/value/growth indices and small/large capitalization indices). Developed and emerging markets are treated separately in Panel I and II and jointly in Panel III. Fama-French factors are obtained by combining U.S. factors with their global counterparts. The sample period is February 1976 to April 2013.

| Model | Test assets | | | | |
|---|---|---|---|---|---|
| | Aggr. Market | Value | Growth | Small | Big |
| | I: Developed Markets | | | | |
| World CAPM | 28.68 | 3.62 | 1.29 | 4.91 | 1.81 |
| Int. CAPM | 29.72 | 4.91 | 0.26 | 4.91 | 1.03 |
| 4FF | 6.20 | 2.07 | 4.65 | 0.00 | 4.13 |
| CAPM Redux | 8.01 | 2.58 | 1.81 | 2.84 | 2.33 |
| | II: Emerging Markets | | | | |
| World CAPM | 10.59 | 3.88 | 5.17 | 3.36 | 3.10 |
| Int. CAPM | 10.34 | 0.78 | 2.84 | 5.43 | 3.10 |
| 4FF | 3.88 | 2.58 | 0.00 | 2.07 | 3.10 |
| CAPM Redux | 6.72 | 0.00 | 0.26 | 1.03 | 5.68 |
| | III: All Markets | | | | |
| World CAPM | 11.89 | 1.03 | 2.07 | 6.46 | 3.62 |
| Int. CAPM | 17.31 | 2.84 | 1.81 | 6.72 | 2.33 |
| 4FF | 2.07 | 2.07 | 2.84 | 3.10 | 4.39 |
| CAPM Redux | 5.43 | 0.00 | 2.58 | 5.17 | 1.29 |

given asset pricing model are jointly zero, at the 5% level. To combat the "large $N$, small $T$" issue, we report this fraction separately for each category of test assets, i.e., the country indexes, as well as all value, growth, small, and big stock portfolios across developed and emerging markets, as well as for all markets together.

The table shows that our model exhibits far lower fractions of rejections of the GRS null than the World CAPM and the International CAPM, especially when applied to the country indexes. Interestingly, this outperformance of our model in the time series domain is particularly pronounced for the developed markets

rather than for the emerging markets. The Fama-French-Carhart four factor model has performance comparable to that of our model across both emerging and developed markets, but for the set of Value test assets, our model generally beats even this model, despite the fact that the Fama-French model was first constructed to explain the value and size premiums.

Using the time-series estimates of factor betas, we turn to the factor risk premiums in the cross-section of test asset returns.

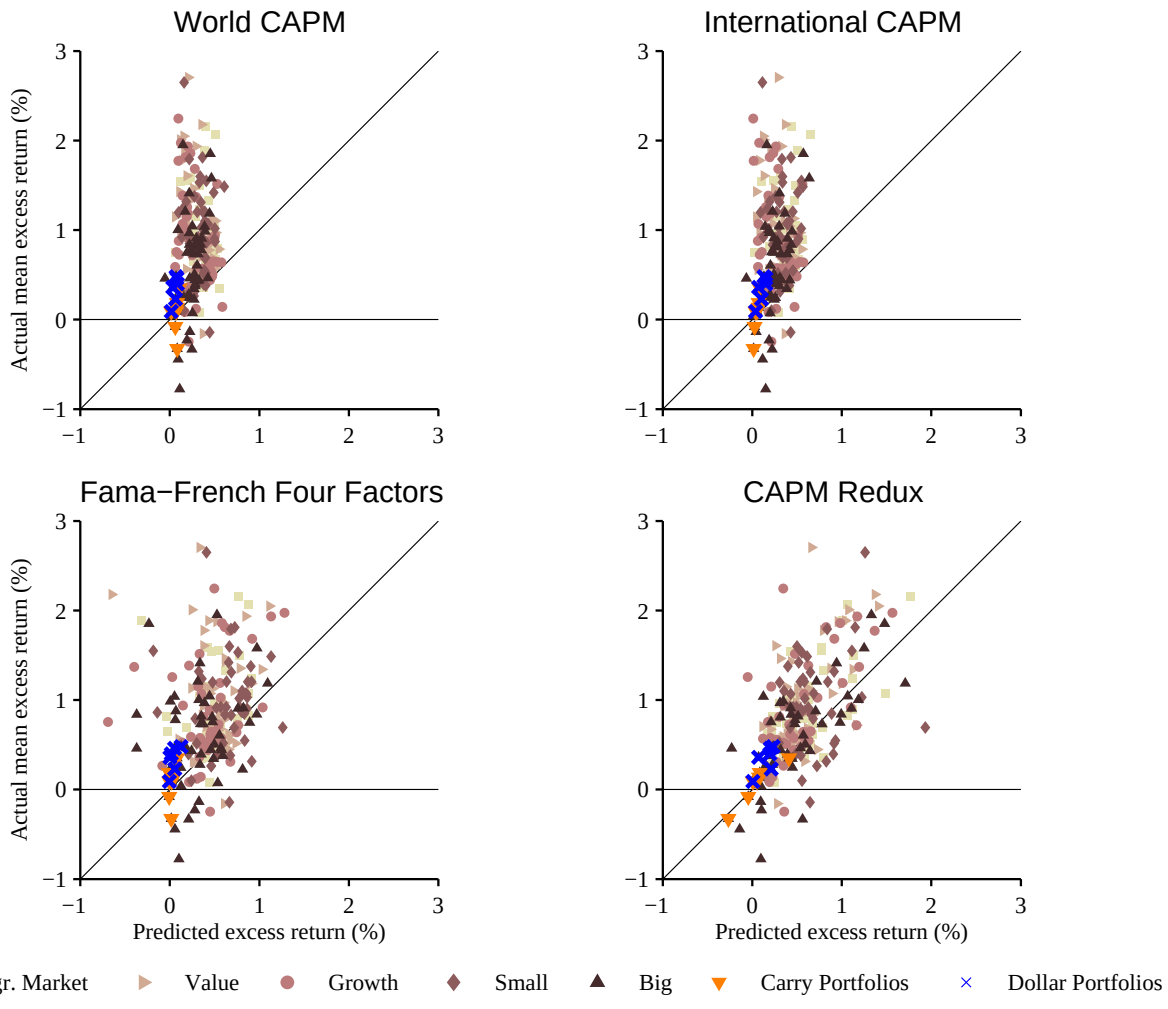
## 4.2 Cross-sectional Tests

As a preliminary exercise, Figure 5 provides a pictorial representation of the relative performance of the models in the cross-section. The vertical axis in these plots is common to all models, and reports the realized average excess returns (in percentage) of our entire cross-section of test assets over the full sample period. The horizontal axis reports the average excess returns of the same test assets as predicted by each model, and varies across the four models we inspect. For the purposes of these plots, we compute predicted returns by using the factor loadings for each asset estimated using the 60-month rolling windows described above. We then multiply these estimated factor loadings by factor means computed over the same time window, and average these conditional predictions across all periods. As usual, better performing models will generate points which lie close to the $45^{\circ}$line, and deviations from this line indicate pricing errors.

The figure shows that the World CAPM and the International CAPM underestimate realized average excess returns: the large cross-sectional variation observed in the data is not matched because there is little cross-sectional variation in the predictions of these models, leading to a more vertical line. From the figure it is apparent that this poor performance is not driven by a particular type of equity asset; the differences between predictions and realizations are similar for the various different types of assets. The Fama-French-Carhart four-factor model does do substantially better than the more theoretically grounded models, but there is still significant deviation from the $45^{\circ}$line.

Figure 5
Realized versus Predicted Average Excess Returns

This figure plots realized average excess returns against those predicted by the World CAPM, the International CAPM, the Fama-French four-factor model and the CAPM Redux. All models are estimated conditionally. For each country the predicted excess returns are computed as follows: a) Time-varying factor betas are estimated using 60-month rolling windows; b) At each time t estimated conditional betas are multiplied by the corresponding factor means computed over the same time-window (from time *t-59* to time *t*); c) Predicted values are averaged. For each country the realized average excess returns are computed as follows: a) In each 60-month rolling window actual excess returns are averaged and b) 60-month rolling return means are averaged. Test assets are equity excess returns for all asset types and all countries, six currency portfolios sorted on forward discounts (Carry portfolios) and six currency portfolios sorted on dollar exposure (Dollar portfolios). Average returns are in percentage. The straight line is the 45-degree line through the origin. Fama -French factors are U.S. (global) factors prior to (from) November 1990. The sample is February 1981 to April 2013.



Our model significantly improves the visual relationship between predicted and realized average returns. These improvements come from two sources. First, we explicitly model the currency component embedded in foreign equity market returns. Second, we reduce the noise in measured currency risks by relying on two global currency risk factors rather than a selected few currency excess returns.

We confirm this result in a number of ways in the Online Appendix. First, we find that the cross-sectional average $\bar{R}^2$ from these time-series regressions is generally higher for our model than for the competition. At each date in the sample, the global equity, dollar, and carry factors explain a larger share of the time-series variation in international equity returns than the world CAPM. We also find that our model outperforms the International CAPM as we move from the distant past towards the recent past. This latter finding suggests that the increasing integration of global markets might account for the increasing explanatory power of global currency factors.

Next, we relax the no-arbitrage condition that pins down the market prices of risk and estimate them using the cross-section of excess returns. We conduct our cross-sectional tests using the standard two-pass cross-sectional approach of Fama and MacBeth (1973, henceforth FMB). For each of the models in our comparisons, we run FMB tests in three different ways. In the first variant of these FMB tests (which we denote as FMB$^1$), we estimate the factor loadings using unconditional time-series regressions over the full sample. We then compute market prices of risk ($\lambda$) via a cross-sectional regression of average excess returns of the test assets on these unconditional factor loadings. In the second variant (which we denote FMB$^2$), first-stage betas continue to be obtained using unconditional time-series regressions over the full sample as in FMB$^1$. However, in the second stage, we run $T$ cross-sectional regressions, one for each time period, of country excess returns on these estimated factor loadings. The average market prices of risk are then computed as simple averages of the slope coefficients obtained from these $T$ cross-sectional regressions. In the third variant (denoted FMB$^{TV}$), we obtain time-varying factor loadings using our rolling regressions over 60-month windows.

In each period $t+1$, we estimate market prices of risk $\lambda_{t+1}$ using cross-sectional regressions of test asset returns on these time-varying factor loadings estimated using windows ending in period $t$. Average market prices of risk are once again simple averages of $\lambda_{t+1}$ over all periods $T$. In all three variants of these tests, we omit constants in the second stage of the FMB procedure. To ensure that the three tests are comparable, the FMB$^1$ and FMB$^2$ tests are carried out on the second-stage estimation sample of the FMB$^{TV}$ procedure.

Table 2 reports the estimates of the average market prices of risk along with Shanken (1992)-corrected standard errors (in parentheses) from these tests across models which are in blocks of rows. The columns identify the set of test assets on which we run these tests. The first set of test assets includes aggregate market excess returns, and value and size-sorted portfolios for the developed markets in the sample (120 assets), as well as 12 currency portfolios. The second set of test assets expands the equity cross-section to include assets from emerging markets, leading to a total of 225 equity assets plus 12 currency portfolios.

The table shows that across all model specifications, cross-sections of test assets, and testing approaches, the world equity factor is priced, whether it is measured in local currency or U.S. dollar terms. The statistical significance of this result holds at the five percent level or better. This finding supports the early evidence in the international finance literature that international investors are compensated for taking on risk that is correlated with returns on the world equity market portfolio.

In contrast, there is little evidence to support the pricing of currency risk in models other than our own. Looking across the three currencies included in the International CAPM, only the British pound appears to carry a significant currency premium in our sample. Moreover, this result holds only when time-variation is taken into account (FMB$^{TV}$). The importance of using a conditional model is consistent with the findings of Dumas and Solnik (1995), however even allowing for time-variation in factor loadings and risk premia, the German mark/Euro and Japanese Yen are not priced over the sample period for the wider cross-section of assets. This result is potentially attributable to the longer sample period, the larger

Table 2

Asset Pricing Fama-MacBeth Tests

This table reports results from the Fama-MacBeth asset pricing procedure for the World CAPM, the International CAPM, the Fama-French four-factor model (4FF, U.S. and global factors are combined) and the CAPM Redux. In FMB$^1$, the average market prices of risks ($\lambda$) are obtained via a cross-sectional regression of average excess returns on the (unconditional) first-step betas. In FMB$^2$, the $\lambda$s correspond to the average across T cross-sectional regressions of excess returns on the same unconditional betas. In FMB$^{TV}$, in the first step of the procedure time-varying (TV) betas are estimated over 60-month rolling windows ending at time $t$. In the second-step, the market prices of risk are estimated at each time t+1 via a cross-sectional regression of country excess returns on the first-step conditional betas. The $\lambda$s are obtained as average of these second-stage estimates. In all cases, the second stage of the procedure does not include a constant and factors are added to the set of test assets. *MAPE* (*RMSE*) denotes the Mean Absolute Pricing Error (Root of Mean Square pricing Errors). Shanken (1992)-corrected standard errors are in brackets. *, **, *** denote statistical significance of the average prices of risk at the 10%, 5% and 1% level, respectively. Test assets are country equity excess returns for developed markets (DM, left panel) or all markets (right panel), six currency portfolios sorted on forward discounts (Carry portfolios) and six currency portfolios sorted on dollar exposure (Dollar portfolios). The sample period is February 1981 to April 2013.

| Model | | DM Equity + FX Portfolios | | | All Equity + FX Portfolios | | | Factor |
|---|---|---|---|---|---|---|---|---|
| | | FMB$^1$ | FMB$^2$ | FMB$^{TV}$ | FMB$^1$ | FMB$^2$ | FMB$^{TV}$ | Means |
| | $\lambda_{WMKT}$ | 0.623** | 0.679*** | 0.683*** | 0.735*** | 0.803*** | 0.811*** | 0.518 |
| World | | (0.258) | (0.262) | (0.263) | (0.266) | (0.269) | (0.261) | (0.252) |
| CAPM | *MAPE* | 0.251 | 0.260 | 0.308 | 0.320 | 0.328 | 0.441 | |
| | *RMSE* | 0.333 | 0.352 | 0.408 | 0.425 | 0.438 | 0.603 | |
| | $\lambda_{WMKT}$ | 0.661** | 0.698*** | 0.597** | 0.755*** | 0.825*** | 0.736*** | 0.518 |
| | | (0.257) | (0.261) | (0.254) | (0.269) | (0.276) | (0.253) | (0.252) |
| | $\lambda_{FX}^{GBP}$ | 0.203 | 0.188 | 0.387** | 0.309 | 0.12 | 0.468*** | 0.190 |
| | | (0.198) | (0.215) | (0.166) | (0.214) | (0.255) | (0.174) | (0.160) |
| Int. | $\lambda_{FX}^{JPY}$ | 0.074 | 0.032 | −0.043 | −0.143 | −0.123 | 0.095 | -0.012 |
| CAPM | | (0.246) | (0.281) | (0.210) | (0.274) | (0.348) | (0.205) | (0.183) |
| | $\lambda_{FX}^{EUR}$ | 0.018 | 0.039 | 0.108 | −0.105 | −0.160 | 0.146 | 0.118 |
| | | (0.219) | (0.250) | (0.188) | (0.221) | (0.254) | (0.177) | (0.177) |
| | *MAPE* | 0.239 | 0.238 | 0.286 | 0.303 | 0.305 | 0.393 | |
| | *RMSE* | 0.326 | 0.331 | 0.410 | 0.402 | 0.406 | 0.550 | |
| | $\lambda_{WMKT}$ | 0.620** | 0.682** | 0.709** | 0.699*** | 0.771*** | 0.780*** | 0.459 |
| | | (0.264) | (0.270) | (0.288) | (0.267) | (0.271) | (0.283) | (0.257) |
| | $\lambda_{SMB}$ | 0.158 | 0.013 | 0.317** | 0.189 | 0.261 | 0.321** | -0.036 |
| 4FM | | (0.154) | (0.222) | (0.146) | (0.173) | (0.249) | (0.141) | (0.121) |
| (Combined) | $\lambda_{HML}$ | −0.147 | −0.118 | −0.164 | −0.221 | −0.244 | −0.148 | 0.324 |
| | | (0.178) | (0.198) | (0.152) | (0.187) | (0.196) | (0.142) | (0.130) |
| | $\lambda_{MOM}$ | 0.197 | 0.239 | 0.399 | −0.049 | 0.016 | 0.204 | 0.672 |
| | | (0.259) | (0.260) | (0.245) | (0.332) | (0.375) | (0.238) | (0.216) |
| | *MAPE* | 0.251 | 0.234 | 0.284 | 0.315 | 0.330 | 0.357 | |
| | *RMSE* | 0.331 | 0.315 | 0.377 | 0.414 | 0.434 | 0.496 | |
| | $\lambda_{LWMKT}$ | 0.542** | 0.588** | 0.526** | 0.680*** | 0.720*** | 0.572** | 0.418 |
| | | (0.248) | (0.256) | (0.242) | (0.253) | (0.264) | (0.241) | (0.235) |
| CAPM | $\lambda_{Dollar}$ | 0.187 | 0.189 | 0.241* | 0.101 | 0.132 | 0.293** | 0.234 |
| Redux | | (0.143) | (0.157) | (0.130) | (0.147) | (0.159) | (0.126) | (0.119) |
| | $\lambda_{Carry}$ | 0.436** | 0.493* | 0.450*** | 0.428** | 0.699** | 0.418*** | 0.615 |
| | | (0.207) | (0.255) | (0.169) | (0.185) | (0.282) | (0.149) | (0.130) |
| | *MAPE* | 0.237 | 0.235 | 0.288 | 0.302 | 0.289 | 0.362 | |
| | *RMSE* | 0.328 | 0.333 | 0.412 | 0.403 | 0.394 | 0.500 | |

cross-section of test assets, and our use of a methodology based on rolling windows instead of instrumental variables to model time-variation.

The table also shows that there is no statistical evidence for a value or momentum premium in the cross-section of test assets, either conditionally or unconditionally. However there is some evidence to support the existence of a size premium, especially when we evaluate the Fama-French model allowing for time-variation in factor loadings.

In contrast with these results, we find evidence to support the pricing of currency risk when we measure this risk using the dollar and carry factors. This is even true to some extent unconditionally, in the sense that the carry factor is priced using $\text{FMB}^1$ across developed and developed plus emerging cross-sections, and using $\text{FMB}^2$ in the broader developed plus emerging cross-section.

The results supporting our model become substantially stronger when we account for time-variation in factor loadings as well as in risk premia, using $\text{FMB}^{TV}$. The prices of dollar and carry risk are all significantly different from zero at the 5% level or better, in both cross-sections.

Additional evidence on the models is provided when we inspect the prices of risk of the equity, dollar, and carry factors and compare them with the average excess returns of those factors which are provided in the final column of Table 2. The no-arbitrage condition implies, since the beta of each factor on itself is obviously one, that the market price of risk of each factor should be equal to the average of the factor. While the sample is short, leading to difficulties in estimating this relationship precisely, we do see that these factor means are relatively close to the estimated factor risk premia. For the other models, the price of equity risk is much higher, further removed from its sample mean. As noted, however, the sample is indeed short, leading to substantial imprecision, and susceptibility to Daniel and Titman's (1997) critique. We cannot of course rule out a characteristics-based behavioral explanation of the cross-sectional variation in test asset returns.

For each model, Table 2 also reports the cross-sectional Mean Absolute Pricing Error (MAPE) and the cross-sectional Root Mean Square Error (RMSE). Consistent

with the evidence discussed above, our factor model delivers relatively smaller RMSEs than the competition, although these differences are not substantial.

The Online Appendix reports a number of additional robustness checks including re-estimating the model on samples of expanding sizes, either moving forwards in time, beginning in 2/1976 or moving backwards in time, beginning in 12/2013. The dollar and carry factors appear priced even in samples that exclude the recent financial crisis (i.e., before 2007). Small perturbations in the estimation sample do not imply abrupt changes in the estimates. The magnitude of these currency premia varies smoothly over time suggesting that the price of risk (not merely risk exposure) is time-varying. We also find that the significance of the pricing results for the model increases over time. This pattern is certainly related to the increase in power arising from a broader cross-section of asset returns available to test the model, as well as the increasing global capital market integration over time.

One obvious question is whether our strong results are simply a consequence of the currency return component embedded in dollar-denominated test asset returns, and the pricing of these currency components by the carry and dollar factors. Figure 6 shows that a simple story of this nature would be insufficient to explain the somewhat involved dynamics of equity and currency risk.

The figure reports the cross-sectional average correlation coefficient between country aggregate market equity returns expressed in local currency and the dollar factor (left panel) and the carry factor (right panel). Clearly, equity and currency risk are not orthogonal to one another, and demonstrate significant time-variation in their relationship. Particularly over the second half of the sample period that we consider, these correlations are significant and increasing, sometimes non-linearly. These correlations reach 60% (50%) at the end of our sample period for the dollar (carry) factor. We discuss this issue in more detail when we present the model in the next section.

Next, we show that a large proportion of mutual and hedge funds investing internationally are also exposed to currency risk, which we are well able to detect using our model.

## 4.3 Global Risk in Mutual and Hedge Fund Returns

Mutual fund managers can choose whether or not to hedge the currency exposure in their foreign equity investments.[27] While many mutual fund brochures remain vague about currency risk, we find that their realized returns reveal a clear and economically significant exposure to exchange rate fluctuations.

As an initial exercise, we simply document the explanatory power of our model for these returns over the entire sample period. A clear pattern emerges in the time-series, as shown in the top panel of Figure 7: over time, despite the number of international funds rising throughout the sample period, there is a steady increase in the percentage of funds exposed to the global equity, dollar, and carry factors in the data. The percentage of mutual funds significantly exposed to dollar or carry risk increases from roughly 60% of funds levels in the mid 1990s to close to 100% of funds by the end of the sample. The average cross-sectional $R^2$ at each date increases from around 60% in the early 1990s to more than 90% at the end of our sample. Clearly, international mutual fund returns are exposed to currency risk, and as a consequence, their investors bear significant currency risk as well.[28]

The bottom panel of the same figure assesses the growth in international hedge funds' exposures to the factors in our model. While exposure to the global equity factor appears stable throughout the period, the combined exposure to the dollar and carry factors has increased notably: very few hedge funds load significantly on the currency factors in the 1990s but close to half the funds do so by the end of the sample.

[27] According to the Wall Street Journal (August 4, 2013), for example, Fidelity Overseas manages $2.3 billion without hedging its foreign currency risk, while Oakmark International, which has $19.5 billion in assets, "hedges up to 80% of its exposure to a given currency when the currency's exchange rate against the dollar is more than 20% above what the management team considers fair value." It is not clear whether this fair value is known to outside investors or not.

[28] We check the robustness of our findings on longer rolling windows of 120 months, instead of 60 months. The number of funds with available data decreases. There are in this case only 568 funds at the end of the sample, but those funds as a group still manage more than $620 billions. The results are broadly similar as above. On average over the whole period, 80% of the mutual funds are significantly exposed to currency risk, representing 91% of the total funds under management in our dataset.

Overall, while hedge fund returns appear to be less exposed to the factors in our model than mutual funds, it is nonetheless true that a large number of hedge fund returns load significantly on our risk factors. This means that their exposure to currency risk can potentially be reproduced at a much lower cost than by incurring the substantial fees involved in investing in a hedge fund. On average across the sample, one can reject the null joint hypothesis that their returns do not load on the equity and currency factors for close to 60% of international hedge funds, representing a similar share of the money invested in the "Macro" and "Emerging" hedge fund sectors.

Figure 8 presents a simple comparison across the models that we consider, regressing monthly mutual fund and hedge fund returns on the competing models, and plotting CDFs of the $t$-statistics of (positive) alphas (using bootstrap standard errors) obtained from these different models. For mutual funds, in the left panel, the figure reveals that the performance of our model is better than that of the World CAPM and the International CAPM, but not as good as that of the Carhart four-factor model.[29].

The right-hand panel of the figure shows the same CDF of $t$-statistics of positive alphas for hedge funds. The figure reveals that the performance of our model is better even than that of the Carhart four-factor model, delivering a slightly lower number of funds with statistically significant alphas.

Having described our empirical results, we now turn to describing a simple theoretical model that helps explain our results.

[29]As we document in the Online Appendix, the MSCI World indices include only developed countries. We can account for all emerging markets in our dataset by constructing our own market index using lagged market capitalization as weighting scheme. Using this extended global equity factor we have verified that our model and the Carhart four-factor model share similar performances.

Figure 6

Cross-Sectional Average Time-Varying Correlation Between Equity Returns in Local Currency and Each Currency Factor

This figure shows, for each 60-month rolling window, the average cross-sectional correlation between country aggregate market returns expressed in local currency and each currency factor (dollar/carry factor in the left/right graph, respectively). Specifically, at each time t we compute pairwise correlations between the equity return of each country-$i$ and a given currency factor and report the average correlation across countries. In each rolling window, a 95% confidence interval (dash-dotted line) is obtained via pairwise bootstrapping (1000 bootstrap samples). The lower/upper bound corresponds to the $2.5^{th}$ and $97.5^{th}$ percentile of the bootstrap distribution of the cross-sectional average correlation in that time-window. The sample is February 1976 to April 2013.

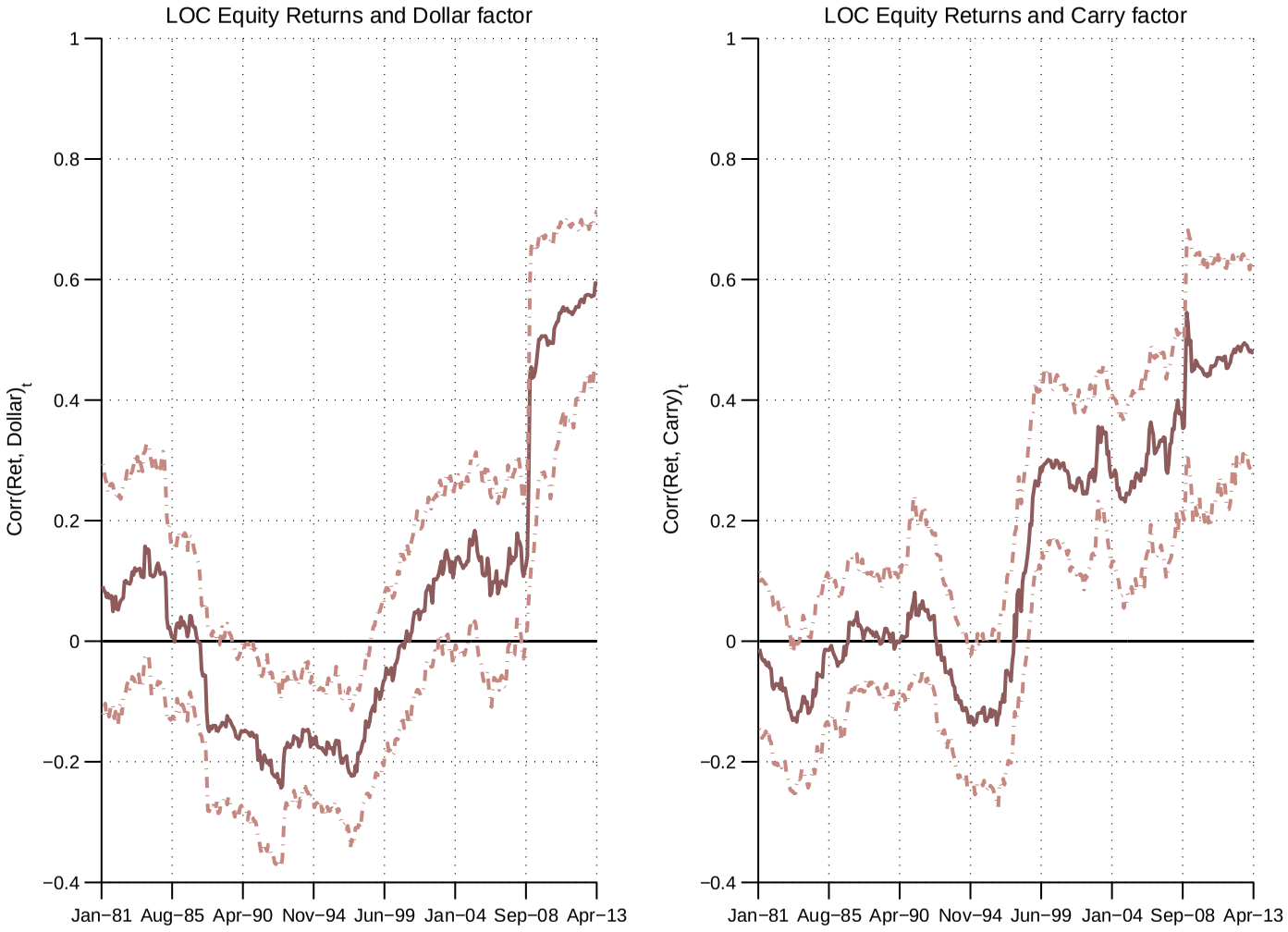


Figure 7
Mutual Funds' and Hedge Funds' Significant Exposure to Global Factors

This figure plots the time-series of the share of hedge funds (top three graphs) and mutual funds (bottom three graphs) with significant exposure to the CAPM Redux factors (LWMKT, Dollar and Carry). At each point in time monthly hedge fund/mutual fund returns are regressed on a constant and the three global factors over rolling-windows of 60 months. The first and fourth panels report the number of funds with available data. The second and fifth panels report the percentage of funds with significance exposure to LWMKT at the 5% significance level (i.e., $t$-statistic on the equity factor above 1.96 in absolute value). Standard errors are obtained via bootstrapping (1000 bootstrap samples). The third and last panel report the percentage of funds with significance exposure to both the dollar and the carry factor at the 5% significance level (i.e. p-value of the $F$-test below 5% under the null hypothesis that both FX loadings are zero). Funds are equally-weighted. Hedge fund data are from the updated version of the consolidated hedge fund database by Ramadorai (2013) and Patton, Ramadorai and Streatfield (2013). The sample includes all funds classified as "Macro" or "Emerging" according to the strategy code. The sample period is January 1994 to April 2013. The mutual fund sample includes all funds classified as "Foreign Equity Funds" according to the CRSP fund style code. The sample period is November 1990 to April 2013. All data are monthly.

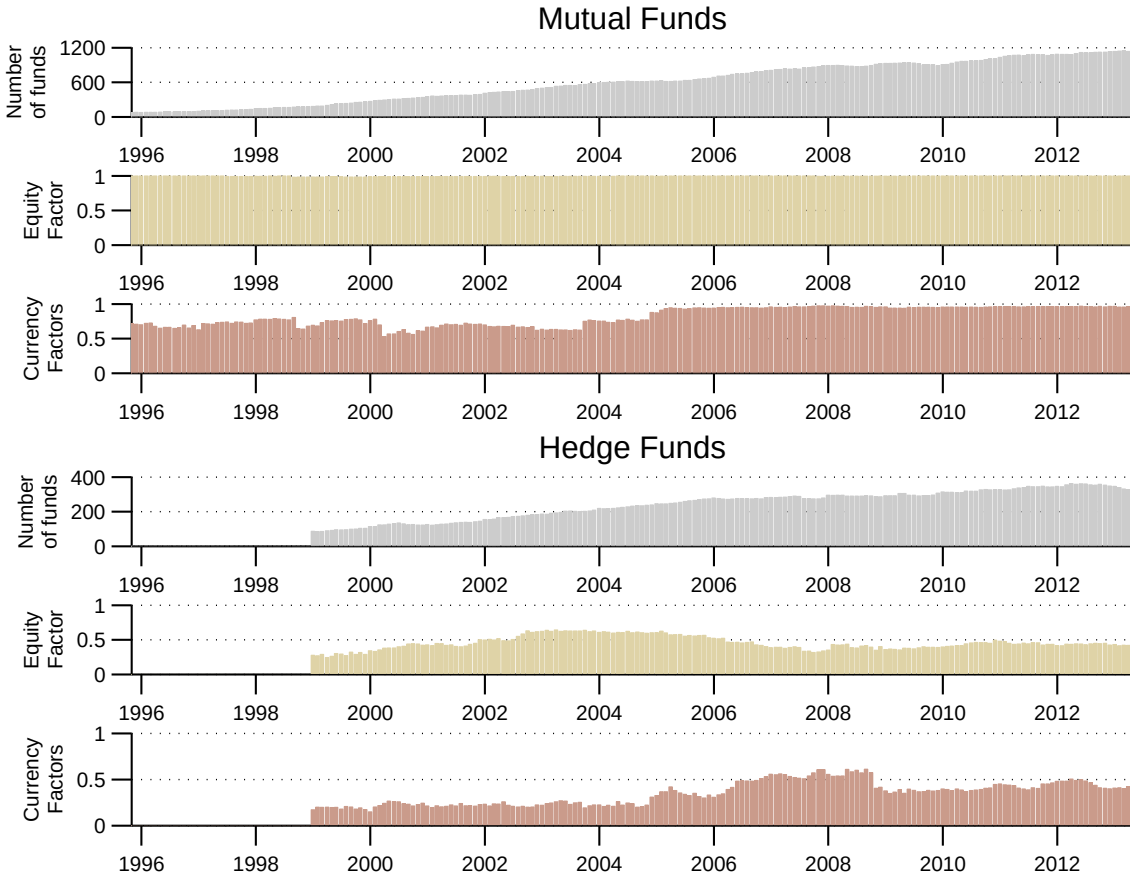
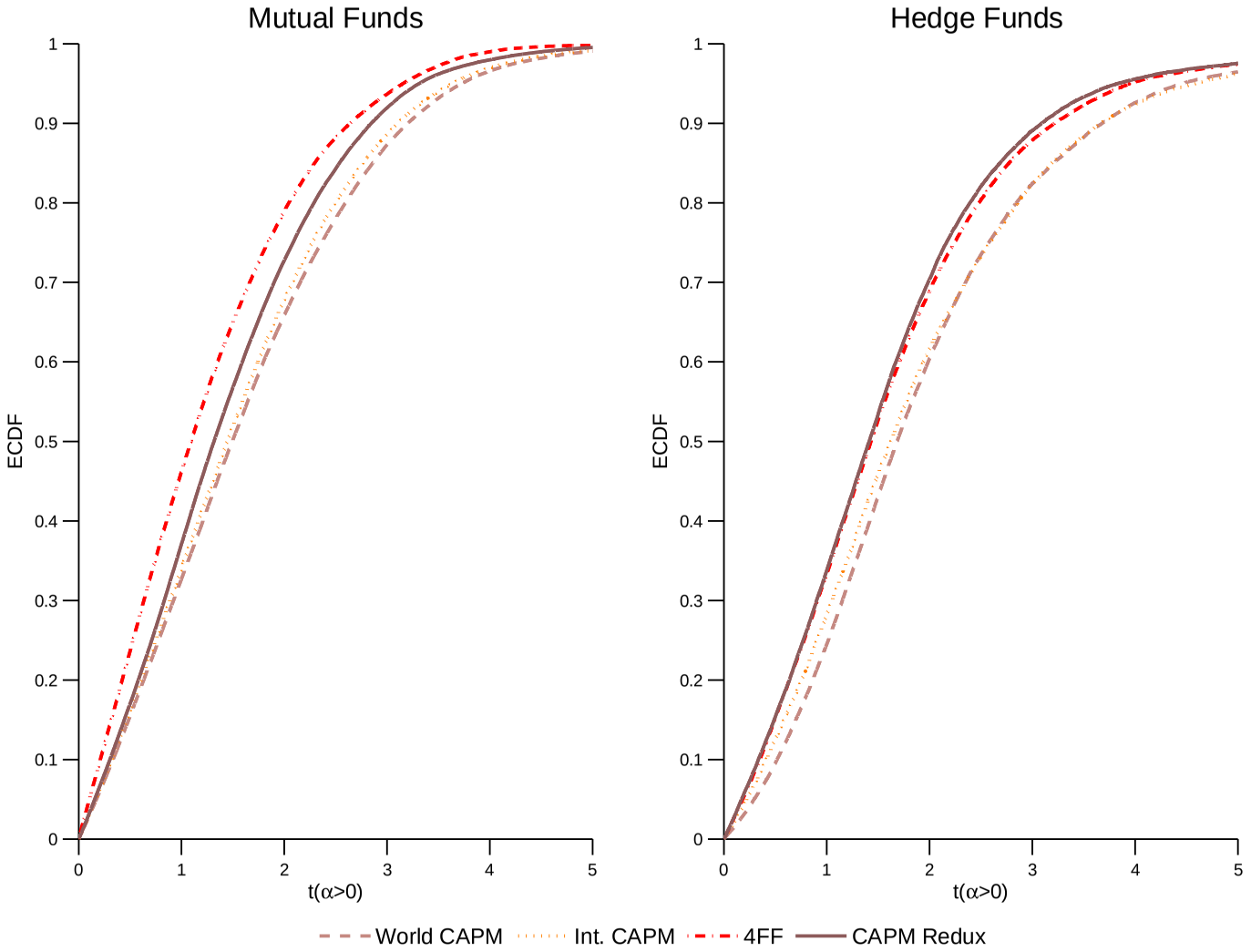


Figure 8

Statistical Significance of Mutual Funds' and Hedge Funds' Alphas

This figure shows the empirical cumulative distribution function (ECDF) of t-statistics of positive hedge funds' (left graph) and mutual funds' (right graph) rolling alphas. Alphas are obtained by estimating the World CAPM, the International CAPM, the Fama-French four-global factor model (4FF) and the CAPM Redux over rolling-windows of 60 months. Standard errors are obtained via bootstrapping (1000 bootstrap samples). Hedge fund data are from the updated version of the consolidated hedge fund database by Ramadorai (2013) and Patton, Ramadorai and Streatfield (2013). The sample includes all funds classified as "Macro" or "Emerging" according to the strategy code. The sample period is January 1994 to April 2013. The mutual fund sample includes all funds classified as "Foreign Equity Funds" according to the CRSP fund style code. The sample period is November 1990 to April 2013. All data are monthly.



# 5 The International CAPM Redux Model

We begin by discussing the theoretical literature on international asset pricing, to provide context about how our simple model fits in to the broader asset pricing literature. We then move to a formal presentation of the model.

The International CAPM is very general, but with this strength comes some costs. The most important one is that exchange rate shocks, while priced, are exogenous to the model, despite potentially being linked to world market returns. This can be seen clearly in Adler and Dumas (1983), which forms the basis of the empirical work of Dumas and Solnik (1995).

Our simple International CAPM Redux model is a modest attempt to plug this gap in the literature. In our model, exchange rates, currency risk factors, and equity market returns all depend on the same shocks. Our focus is to explore what this change buys us in a relatively stripped-down setting. In our model, financial markets are complete, and we specify the law of motion of the lognormal stochastic discount factors (SDFs) in all countries. The SDFs are posited to depend on country-specific shocks, as well as three global shocks. We assume that each country's aggregate dividend growth rate depends on the same shocks as the country-specific SDF. In order to specify a role for a pure equity risk factor affecting all countries in the same way, we posit that one global shock affects all SDFs in the same way.

To preview the main insight obtained from the model, we find that the world stock market return in U.S. dollars depends on all global shocks, meaning that there should be no role for bilateral exchange rates, i.e., the World CAPM should work. Yet the model shows that the world equity return expressed in U.S. dollars bundles state variables with different time-dynamics together. This makes it difficult to use this single factor to uncover risk exposures in international asset pricing since the state variables are unknown to the econometrician. We find in this setting that our three factor empirical model is the best way to capture the various sources of global risk posited in the theory. We now turn to describing our model more formally.

## 5.1 SDFs

In the tradition of Backus, Foresi, and Telmer (2001), we assume that pricing kernels $m^i_{t+1}$ are exponentially affine:

$$\begin{aligned} -m^i_{t+1} &= \alpha + \chi z^i_t + \sqrt{\gamma z^i_t} u^i_{t+1} + \tau z^w_t + \sqrt{\delta^i z^w_t} u^w_{t+1} + \sqrt{\kappa z^i_t} u^g_{t+1} + \sqrt{\omega z^w_t} u^c_{t+1}, \\ z^i_{t+1} &= (1-\phi)\theta + \phi z^i_t - \sigma \sqrt{z^i_t} u^i_{t+1}, \\ z^w_{t+1} &= (1-\phi^w)\theta^w + \phi^w z^w_t - \sigma^w \sqrt{z^w_t} u^w_{t+1} \end{aligned}$$

where $u^i_{t+1}, u^w_{t+1}, u^g_{t+1}, u^c_{t+1}$ are i.i.d, mean-zero, variance-one Gaussian shocks, $m^i_{t+1}$ is the log SDF of country $i$, and $z^i_t$ and $z^w_t$ are the state variables that govern the conditional volatility of the SDF. Each SDF is heteroskedastic because currency risk premia are driven by the conditional variances of the SDFs. Inflation is not a priced risk in this model – the SDFs can be interpreted as nominal SDFs. We use the U.S. dollar as the base currency and drop the superscript $i$ to describe any U.S. variable.

A similar model is studied in Lustig, Roussanov, and Verdelhan (2011, 2014) and in Verdelhan (2014).[30] The key distinguishing feature of our model is the introduction of equity-specific shocks $u^c_{t+1}$. As we shall see, these shocks affect both dividends and SDFs, but not exchange rates, and drive the world equity return factor.

The SDFs depend on country-specific shocks, $u^i_{t+1}$, and three global shocks, $u^w_{t+1}$, $u^g_{t+1}$, and $u^c_{t+1}$. We refer to the volatilities of the SDF related to these three shocks (namely, $\sqrt{\delta^i z^w_t}$, $\sqrt{\kappa z^i_t}$, and $\sqrt{\omega z^w_t}$) as the market prices of risk of these shocks.

The first shock $u^w_{t+1}$ is priced similarly in each country up to a scaling factor, denoted $\delta^i$. Examples of such a shock might be a global financial crisis which affects prices in all countries in a perfectly correlated fashion, but with differential intensity. To be parsimonious, we model differences in exposure $\delta^i$ as the only source of heterogeneity in countries' SDFs, and fix all the other parameters of the SDFs to be the same across countries. Countries also differ in their aggregate dividend growth rates.

[30] In those papers, the prices of risk (i.e., the square roots in the law of motion of the SDFs), depend on both the country-specific and global state variables in order to differentiate between unconditional and conditional currency risk premia. For the sake of clarity, we leave this difference aside here, but the model can be easily extended in this direction. Earlier examples of affine models in international finance include Frachot (1996) and Brennan and Xia (2006).

The model thus features two sources of heterogeneity: one in the SDF and one in the dividend growth rates. The first source of heterogeneity is necessary to account for the cross-section of interest rates and currency excess returns. As Lustig, Roussanov, and Verdelhan (2011) show, high interest rate countries must be characterized by low exposure ($\delta^i$) to world shocks for the high interest rate currency to depreciate in bad times — the key mechanism of any risk-based explanation of carry trade profits. This first source of heterogeneity entails differences in global equity betas, but if it were not for the differences in dividend growth rates, the equity betas would line up with the carry betas. In the data, they do not, consistent with our modeling choice of two sources of cross-country differences.

The second shock $u^g_{t+1}$ is priced differently across countries, even if countries share the same exposure $\kappa$. An example of this might be a productivity shock that affects some economies more than others. Finally, as described earlier, the third shock $u^c_{t+1}$ is priced in exactly the same way in all countries.

## 5.2 Exchange Rates

When markets are complete, log changes in exchange rates correspond to the differences between domestic and foreign log pricing kernels (Bekaert, 1996, Bansal, 1997):[31]

$$\begin{aligned}\Delta s^i_{t+1} &= m_{t+1} - m^i_{t+1},\\ &= \chi(z^i_t - z_t) + \sqrt{\gamma z^i_t}u^i_{t+1} - \sqrt{\gamma z_t}u_{t+1}\\ &+ (\sqrt{\delta^i} - \sqrt{\delta})\sqrt{z^w_t}u^w_{t+1} + \sqrt{\kappa}(\sqrt{z^i_t} - \sqrt{z_t})u^g_{t+1},\end{aligned}$$

where the exchange rate is defined in foreign currency per U.S. dollar. Therefore, an increase in the exchange rate corresponds to an appreciation of the U.S. dollar. Although financial markets are complete, real exchange rates are not necessarily constant as soon as some frictions exist in the goods markets (e.g., non-traded goods, or trading costs). The exchange rate between country $i$ and the domestic

[31] This result derives from the Euler equations of the domestic and foreign investors buying any asset $R^i$ that pays off in foreign currency: $E_t[M_{t+1}R^i S^i_t/S^i_{t+1}] = 1$ and $E_t[M^i_{t+1}R^i] = 1$. When markets are complete, the pricing kernel is unique and thus exchange rates are defined as $S^i_{t+1}/S^i_t = M_{t+1}/M^i_{t+1}$, or in logs $\Delta s^i_{t+1} = m_{t+1} - m^i_{t+1}$.

economy depends on the country-specific shocks $u^i$ and $u$, as well as on the global shocks $u^w$ and $u^g$, but not on the global shocks $u^c$ since their prices of risk are the same across countries.

## 5.3 Equity Returns

In order to define equity returns, the model posits a dividend growth process in each country $i$:

$$\Delta d^i_{t+1} = \mu_D + \psi z^i_t + \psi_w z^w_t + \sigma_D \sqrt{z^i_t} u^i_{t+1} + \sigma^w_D \sqrt{z^w_t} u^w_{t+1} + \sigma^g_D \sqrt{z^i_t} u^g_{t+1} + \sigma^{c,i}_D \sqrt{z^w_t} u^c_{t+1},$$

where the innovations are the same as those described above. Dividend growth rates respond to both country-specific and global shocks. The only systematic difference across countries comes from the impact of global shocks $u^c_{t+1}$ on dividend growth, governed by the parameters $\sigma^{c,i}_D$. This source of heterogeneity drives the differences in global equity market betas.

Innovations to the log gross equity excess return are then:[32]

$$\begin{aligned} r^{e,i}_{t+1} - E_t(r^{e,i}_{t+1}) &= (\sigma_D - k_1 B^i_{pd}\sigma)\sqrt{z^i_t} u^i_{t+1} \\ &+ (\sigma^w_D - k_1 C^i_{pd}\sigma^w)\sqrt{z^w_t} u^w_{t+1} + \sigma^g_D \sqrt{z^i_t} u^g_{t+1} + \sigma^{c,i}_D \sqrt{z^w_t} u^c_{t+1}, \end{aligned}$$

where $r^{e,i}_{t+1}$ is the logarithmic gross rate of return on each country's stock market index denominated in that country's currency.

[32] In this model, the log price-dividend ratio is affine in the state variables, $z^i_t$ and $z^w_t$:

$$pd^i_t = A^i_{pd} + B^i_{pd} z^i_t + C^i_{pd} z^w_t,$$

where the constants $A^i_{pd}$, $B^i_{pd}$, and $C^i_{pd}$ are defined as function of the SDF and dividend growth parameters. The model is solved in closed form using the standard log-linear approximation for the log gross return on the aggregate dividend claim:

$$r^{e,i}_{t+1} \approx k_0 + k_1 pd^i_{t+1} - pd^i_t + \Delta d^i_{t+1},$$

where $k_0$ and $k_1$ are defined by the Taylor approximation of the log price-dividend ratio $pd^i_t$ around its mean. More precisely, the Euler equation applied to the stock market return implies that the coefficients $A^i_{pd}$, $B^i_{pd}$ and $C^i_{pd}$ are defined by:

$$\begin{aligned} A^i_{pd} &= -\alpha + k_0 + k_1 A^i_{pd} + k_1 B^i_{pd}(1-\phi)\theta + k_1 C^i_{pd}(1-\phi^w)\theta^w + \mu_D, \\ B^i_{pd} &= k_1 B^i_{pd}\phi + \psi - \chi + \frac{1}{2}(\sqrt{\gamma} + k_1 B^i_{pd}\sigma - \sigma_D)^2 + \frac{1}{2}(\sigma^g_D - \sqrt{\kappa})^2, \\ C^i_{pd} &= k_1 C^i_{pd}\phi^w + \psi_w - \tau + \frac{1}{2}(\sqrt{\delta^i} + k_1 C^i_{pd}\sigma^w - \sigma^w_D)^2 + \frac{1}{2}(\sqrt{\omega} - \sigma^{c,i}_D)^2. \end{aligned}$$

## 5.4 International CAPM Redux

The equity return of country $i$ expressed in U.S. dollars, denoted $r_{t+1}^{e,i,\$}$, is simply derived from the equity return of country $i$ in local currency and the change in the exchange rate:

$$\begin{aligned} r_{t+1}^{e,i,\$} - E_t\left(r_{t+1}^{e,i,\$}\right) &= \sqrt{\gamma z_t}u_{t+1} + \left(\sigma_D - k_1 B_{pd}^i \sigma - \sqrt{\gamma}\right)\sqrt{z_t^i}u_{t+1}^i \\ &+ \left(\sigma_D^w - k_1 C_{pd}^i \sigma^w - \sqrt{\delta^i} + \sqrt{\delta}\right)\sqrt{z_t^w}u_{t+1}^w + \sigma_D^{c,i}\sqrt{z_t^w}u_{t+1}^c \\ &+ \left((\sigma_D^g - \sqrt{\kappa})\sqrt{z_t^i} + \sqrt{\kappa z_t}\right)u_{t+1}^g. \end{aligned}$$

The foreign equity return expressed in U.S. dollars therefore depends on the foreign and U.S. specific shocks ($u^i$ and $u$), as well as the global shock captured by the carry and dollar factors ($u^w$) and ($u^g$) and the world equity shock ($u^c$).

The expected equity excess return from the perspective of a U.S. investor is:

$$\begin{aligned} E_t r_{t+1}^{e,i,\$} - r_{f,t} + \frac{1}{2}Var_t\left[r_{t+1}^{e,i,\$}\right] &= -Cov_t\left[m_{t+1}, r_{t+1}^{e,i,\$}\right] \\ &= \gamma z_t + \left(\sigma_D^w - k_1 C_{pd}^i \sigma^w - \sqrt{\delta^i} + \sqrt{\delta}\right)\sqrt{\delta} z_t^w \\ &+ \sigma_D^{c,i}\sqrt{\omega} z_t^w - \left((\sigma_D^g - \sqrt{\kappa})\sqrt{z_t^i} + \sqrt{\kappa z_t}\right)\sqrt{\kappa z_t} \end{aligned} \tag{7}$$

Note that our assumption of complete markets implies that we only need to verify the Euler condition for one country's investor. As soon as the Euler equation is satisfied for the U.S. investor, for example, it implies that the Euler condition for any foreign investor is also satisfied. To see this point, recall that the Euler condition for the U.S. investor is $E_t\left[M_{t+1}R_{t+1}^i S_t^i/S_{t+1}^i\right] = 1$, which implies $E_t\left[M_{t+1}^i R_{t+1}^i\right] = 1$ — the Euler condition of the representative investor in country $i$ — as well as $E_t\left[M_{t+1}^j R_{t+1}^i (S_t^i/S_{t+1}^i)(S_{t+1}^j/S_t^j)\right] = 1$, the Euler condition of the representative investor in country $j$.

In order to better understand our empirical approach, next, we express various factors in the language of the model, namely, the world equity return in U.S. dollars $WMKT_{t+1}$, the world equity return in local currencies $LWMKT_{t+1}$, and the carry and dollar factors.

## 5.5 Equity Factors

The innovations to the average world equity market return in local currency terms, which we define for ease of exposition as the simple average of local equity returns, are:

$$LWMKT_{t+1} = \overline{r^{e,i}_{t+1}} - E_t(\overline{r^{e,i}_{t+1}}) = (\sigma^w_D - k_1 \overline{C^i_{pd}} \sigma^w)\sqrt{z^w_t} u^w_{t+1} + \sigma^g_D \overline{\sqrt{z^i_t}} u^g_{t+1} + \overline{\sigma^{c,i}_D} \sqrt{z^w_t} u^c_{t+1},$$

where $\overline{x}$ denotes the cross-country average of a variable $x$.

Country-specific shocks average out, and the world equity market return is only driven by the global shocks $u^w_{t+1}$, $u^g_{t+1}$, and $u^c_{t+1}$. If the law of large numbers applies, the cross-sectional mean of $z^i_t$ is constant (and equal to $\theta$). The world equity return in U.S. dollars is:

$$\begin{aligned} WMKT_{t+1} = \overline{r^{e,i,\$}_{t+1}} - E_t\left(\overline{r^{e,i,\$}_{t+1}}\right) &= \sqrt{\gamma z_t} u_{t+1} + \left((\sigma^g_D - \sqrt{\kappa})\overline{\sqrt{z^i_t}} + \sqrt{\kappa z_t}\right) u^g_{t+1} \\ &+ \left(\sigma^w_D - k_1 \overline{C^i_{pd}} \sigma^w - \overline{\sqrt{\delta^i}} + \sqrt{\delta}\right) \sqrt{z^w_t} u^w_{t+1} \\ &+ \overline{\sigma^{c,i}_D} \sqrt{z^w_t} u^c_{t+1}. \end{aligned}$$

## 5.6 Currency Factors

We first express currency excess returns, which are returns on the following strategy: the investor borrows at the domestic risk-free rate, converts the amount into foreign currency and lends at the foreign risk-free rate, converting back the proceeds at the end of the investment period and paying back the debt. The log currency

excess return is thus:[33]

$$rx_{t+1}^i = r_{f,t}^i - r_{f,t} - \Delta s_{t+1}^i.$$

The systematic components of currency excess returns are driven by at least two risk factors (Lustig, Roussanov, and Verdelhan, 2011, 2014, and Verdelhan, 2014), i.e., carry and dollar factors.

The carry factor is the excess returns of a strategy that invests in high- and borrows in low-interest rate currencies:

$$Carry_{t+1} = \frac{1}{N_H}\sum_{i\in H} rx_{t+1}^i - \frac{1}{N_L}\sum_{i\in L} rx_{t+1}^i,$$

where $N_H$ ($N_L$) denotes the number of high (low) interest rate currencies in the sample.

For the sake of exposition, assume that most of the cross-country difference in interest rates is due to their exposure ($\delta^i$) to the world state variable. In this case, baskets of high and low interest rate currencies will exhibit the same level of country-specific volatility, assuming that the law of large numbers holds.[34] Under this assumption, innovations to the carry factor only depend on shocks $u^w$, not on shocks $u^g$.

$$Carry_{t+1} - E_t\left(Carry_{t+1}\right) = \left(\overline{\sqrt{\delta^i}}^L - \overline{\sqrt{\delta^i}}^H\right)\sqrt{z_t^w}u_{t+1}^w,$$

[33]The risk-free rate in country $i$, denoted $r_{f,t}^i$, is given by:

$$r_{f,t}^i = -E_t\left[m_{t+1}^i\right] - \frac{1}{2}Var_t[m_{t+1}^i] = \alpha + \left(\chi - \frac{1}{2}(\gamma+\kappa)\right)z_t^i + \left(\tau - \frac{1}{2}(\delta^i+\omega)\right)z_t^w.$$

The interest rate difference, or forward discount, between country $i$ and the U.S is therefore equal to:

$$r_{f,t}^i - r_{f,t} = \left(\chi - \frac{1}{2}(\gamma+\kappa)\right)(z_t^i - z_t) - \frac{1}{2}\left(\delta^i - \delta\right)z_t^w.$$

The expected currency excess return is therefore:

$$\begin{aligned} E_t\left(rx_{t+1}^i\right) &= r_{f,t}^i - r_{f,t} - E_t\left(\Delta s_{t+1}^i\right) = \frac{1}{2}Var_t(m_{t+1}) - \frac{1}{2}Var_t(m_{t+1}^i) \\ &= \frac{1}{2}(\gamma+\kappa)(z_t - z_t^i) + \frac{1}{2}(\delta - \delta^i)z_t^w. \end{aligned}$$

If the SDFs were not hetereoscedastic, the expected currency excess returns would be constant and the uncovered interest rate parity, which is strongly rejected in the data, would be satisfied in the model.

[34]In practice of course, the number of currencies is small (the dataset used in this paper contains at most 39 currencies). As a result, the law of large numbers is only an approximation used here to provide intuition.

where $\overline{x^i}^H$ and $\overline{x^i}^L$ denote the cross-sectional average of the variable $x$ across countries in the high- and low-interest rate portfolios: $\overline{x^i}^H = \frac{1}{N_H}\sum_{i\in H} x^i$, $\overline{x^i}^L = \frac{1}{N_L}\sum_{i\in L} x^i$.

The dollar risk factor is the average of all currency excess returns defined in U.S. dollars:

$$Dollar_{t+1} = \frac{1}{N}\sum_i rx^i_{t+1},$$

where $N$ denotes the number of currencies in the sample. In large baskets of currencies, foreign country-specific shocks average out (again assuming that there are enough currencies in the baskets for the law of large numbers to apply). As a result, innovations to the dollar risk factor depend on both U.S.-specific and world shocks, but not on country-specific shocks:

$$\begin{aligned} Dollar_{t+1} - E_t\left(Dollar_{t+1}\right) &= \sqrt{\gamma z_t}u_{t+1} \\ &+ \left(\sqrt{\delta} - \overline{\sqrt{\delta^i}}\right)\sqrt{z^w_t}u^w_{t+1} + \sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z^i_t}}\right)u^g_{t+1}. \end{aligned}$$

If the U.S. SDF exhibits the average exposure to shocks $u^w$ (i.e., $\overline{\delta^i} = \delta$), the dollar factor is orthogonal to the carry factor. Under that assumption, the dollar factor captures shocks $u^g$, while the carry factor captures shocks $u^w$.

## 5.7 Connecting the Model to the Empirics

The world equity return, $LWMKT_{t+1}$, built from returns in local currencies, does not span all systematic shocks: it depends on shocks $u^w$, $u^g$, and $u^c$, but not on the U.S. shocks $u$. These shocks $u$ are systematic from the perspective of the U.S. investor, although not from the perspective of other investors. They appear in the cross-section of equity excess returns because the returns need to be expressed in one common currency (e.g., the U.S. dollar).

In contrast, the world equity return in U.S. dollars $WMKT_{t+1}$ contains all the systematic shocks that drive foreign equity returns. At first sight, it appears as a sufficient tool to measure aggregate risk, without the need to add any bilateral exchange rates. This result contrasts with the findings of Adler and Dumas (1983)

because the exchange rate is endogenous in our setup. In practice, however, even in this simple model, the world equity return in U.S. dollars is an imperfect measure of risk in the (obviously general) case when the econometrician does not know the country-specific and global state variables.

To see this point, consider the time-variation in the quantity and market price of aggregate risk, starting with the betas. The betas on the global shocks $u^w$ and $u^c$ are constant, while the betas on the global shocks $u^g$ are not. The total beta on the world equity return in U.S. dollars $WMKT_{t+1}$ is time-varying, following the dynamics of the country-specific state variables $z$ and $z^i$:

$$\beta^i_{WMKT,t} = 1 + \frac{\sigma^w_D - k_1 C^i_{pd}\sigma^w - \sqrt{\delta^i} + \sqrt{\delta}}{\sigma^w_D - k_1 \overline{C^i_{pd}}\sigma^w - \overline{\sqrt{\delta^i}} + \sqrt{\delta}} + \frac{\sigma^{c,i}_D}{\overline{\sigma^{c,i}_D}} + \frac{(\sigma^g_D - \sqrt{\kappa})\sqrt{z^i_t} + \sqrt{\kappa z_t}}{(\sigma^g_D - \sqrt{\kappa})\overline{\sqrt{z^i_t}} + \sqrt{\kappa z_t}} \quad (8)$$

In our simple setup, all country-specific state variables $z^i$ are characterized by the same persistence and volatility. However, in actual data, country-specific state variables seem very likely to evolve at different frequencies, confounding estimation.

Market prices of risk are also time-varying in our simple model. Expected excess returns are the product of betas and market prices of risk, i.e., the market price of world equity risk is the ratio of the expected excess return, defined in Equation (7), to the beta in Equation (8). The price of risk thus also varies with the country-specific and global state variables ($z$, $z^i$, and $z^w$). If those state variables were known to the econometrician, a conditional asset pricing experiment could recover the time-variation in the quantities and prices of risk. In practice however, the state variables are not known, forcing reliance on rolling windows or other such empirical approaches.

The issue becomes even more acute if the heteroskedasticity of the $u^w$ or $u^c$ shocks were to depend on both global and country-specific state variables.[35] In that

[35]Lustig, Roussanov, and Verdelhan (2011, 2014) and Verdelhan (2014) consider this variation in order to reproduce some features of the currency markets. In the context of our model, the law of motion of the log SDF would be:

$$-m^i_{t+1} = \alpha + \chi z^i_t + \sqrt{\gamma z^i_t}u^i_{t+1} + \tau z^w_t + \sqrt{\delta^i z^w_t + \lambda z^i_t}u^w_{t+1} + \sqrt{\kappa z^i_t}u^g_{t+1} + \sqrt{\omega z^w_t}u^c_{t+1}$$

We do not consider that variation here as it does not admit a closed-form solution for equity returns and betas.

case, the aggregate market beta $\beta^i_{WMKT,t}$ also depends on the global state variable $z^w$. The critical issue here is that if the local and global state variables evolve at different frequencies, it becomes impossible to use a single beta to perfectly summarize the time-variation in two state variables.

Taken together, the world equity return expressed in U.S. dollars bundles variables with different time-dynamics together. This makes it difficult to use this single factor to uncover risk exposures in international asset pricing. While our simple model with endogenous exchange rates suggests that we should go back to the World CAPM, the heteroskedasticity of the SDF implies that this single-factor model would struggle to accurately capture the risk-return tradeoff in practice.

This motivates our choice of a multiple-factor model. With $LWMKT$, $Carry$, and $Dollar$, we can summarize all the shocks in the system, and we can allow all quantities and prices of risk to have their own time-dynamics. When we empirically implement the model in the data, the only difference lies in the approach to aggregation, i.e., we weight countries and currencies by stock market capitalizations, instead of the equal weights that we employ in the model for ease of exposition.

The model has three additional implications that we check in the data. First, the carry and dollar factors should matter even for equity excess returns expressed in local currency. Likewise, the equity factor $LWMKT_{t+1}$, built without introducing exchange rates, should be correlated with the carry and dollar factors. Second, despite exchange rates being endogenous, exchange rate shocks in this model do not span equity returns. Shocks $u^c$ affect equity returns but do not affect exchange rates, because their impact is exactly the same across countries. Equity and currency markets therefore appear segmented to an extent. Third, while using bilateral exchange rates as risk factors directly is consistent with the model, they are also driven by country-specific shocks that are irrelevant to asset pricing and thus weaken the identification of aggregate risk.

## 5.8 Shortcomings of the Model

The model is parsimonious and can be solved in closed form. But its simplicity entails some shortcomings when compared to the data. The main weakness of the model lies in its implied betas.

In the data, global equity and currency betas are clearly time-varying. Moreover, the global equity, carry, and dollar betas are not highly correlated. The time-series correlations (obtained country-by-country) range from $-0.7$ to 0.7 for the $LWMKT-Dollar$, $LWMKT-Carry$, and $Dollar-Carry$ pairs. Across countries, these correlations are close to zero. Our asset pricing estimates allow for these factor-specific dynamics in the quantities of risk.

But in the model, betas are either constant or too smooth. A univariate regression of equity excess returns in U.S. dollars on the carry factor, for example, delivers a constant beta; the data suggest otherwise. Time-variation in all betas can be obtained in the model by assuming that the market prices of risk (i.e., the square roots in the SDFs) depend not only on one but on two state variables. In this case, however, the model does not admit a closed-form. For the sake of clarity, we choose to focus on a model that delivers all objects of interest in closed-form.

## 5.9 Simulating the Model

We simulate the model in order to better understand the drivers of our empirical results. We begin by calibrating key parameters of the model, and present these calibration parameters in Table 3.

All parameters governing the state variable dynamics and the parameters $\tau$, $\chi$ and $\gamma$ governing the stochastic discount factor are from Lustig, Roussanov, and Verdelhan (2014). The boundaries of the parameter $\delta$ (i.e., $\delta_l$ and $\delta_h$) are determined to match the mean interest-rate differential of the Japanese yen and the Australian dollar against the U.S. dollar over the period January 1976 to April 2013 (-3.18 % and 3.27 %, respectively). The parameter $\alpha$ is calibrated to match an average annualized nominal risk-free rate of 4.5% given the parameter values for $\chi$, $\gamma$, $\kappa$, $\tau$, $\omega$ and the average $\delta$. The remaining parameters are set to match the empirical

Table 3
Parameter Values

This table reports the parameter values used to simulate the model. The boundaries of the parameter $\delta$ (i.e., $\delta_l$ and $\delta_h$) are determined to match the mean interest-rate differential of the Japanese yen and the Australian dollar against the U.S. dollar over the period January 1976 to April 2013 (-3.18 % and 3.27 %, respectively). All parameters governing the state variable dynamics and the parameters $\tau$, $\chi$ and $\gamma$ governing the stochastic discount factor are from Lustig, Roussanov and Verdelhan (2014). The parameter $\alpha$ is calibrated to match an average annualized nominal risk-free rate of 4.5% given the parameter values for $\chi$, $\gamma$, $\kappa$, $\tau$, $\omega$ and the average $\delta$. The remaining parameters are set to match the empirical pairwise correlations between the equity and currency factors. The law of motion of the stochastic discount factor (Panel A), the state variable dynamics (Panel B) and the dividend growth process (Panel C) are reported at the top of each panel.

Panel A: Stochastic discount factor

$$-m^i_{t+1} = \alpha + \chi z^i_t + \sqrt{\gamma z^i_t} u^i_{t+1} + \tau z^w_t + \sqrt{\delta^i z^w_t} u^w_{t+1} + \sqrt{\kappa z^i_t} u^g_{t+1} + \sqrt{\omega z^w_t} u^c_{t+1}$$

| SDF | | | | | | Heterogeneity | | |
|---|---|---|---|---|---|---|---|---|
| $\alpha$ (%) | $\chi$ | $\gamma$ | $\kappa$ | $\tau$ | $\omega$ | $\delta_{HC}$ | $\delta_L$ | $\delta_H$ |
| 4.5 | 0.89 | 0.04 | 2.2 | 0.06 | 2.2 | 0.36 | 0.10 | 0.61 |

Panel B: State variable dynamics

$$z^i_{t+1} = (1-\phi)\theta + \phi z^i_t - \sigma\sqrt{z^i_t} u^i_{t+1} \qquad z^w_{t+1} = (1-\phi^w)\theta^w + \phi^w z^w_t - \sigma^w\sqrt{z^w_t} u^w_{t+1}$$

| $\phi$ | $\theta$ (%) | $\sigma$ (%) | $\phi^w$ | $\theta^w$ (%) | $\sigma^w$ (%) |
|---|---|---|---|---|---|
| 0.91 | 0.77 | 0.68 | 0.99 | 2.09 | 0.28 |

Panel C: Dividend growth rate

$$\Delta d^i_{t+1} = \mu_D + \psi z^i_t + \psi_w z^w_t + \sigma_D\sqrt{z^i_t} u^i_{t+1} + \sigma^g_D\sqrt{z^i_t} u^g_{t+1} + \sigma^w_D\sqrt{z^w_t} u^w_{t+1} + \sigma^{c,i}_D\sqrt{z^w_t} u^c_{t+1}$$

| Dividends | | | | | | Heterogeneity | | |
|---|---|---|---|---|---|---|---|---|
| $\mu_D$ (%) | $\psi$ | $\psi_w$ | $\sigma_D$ | $\sigma^g_D$ | $\sigma^w_D$ | $\bar{\sigma^c_D}$ | $\sigma^{c,L}_D$ | $\sigma^{c,H}_D$ |
| 2.71 | 0 | -1.1 | 0.12 | 0.05 | 0.01 | 0.19 | 0.16 | 0.23 |

Table 4

Descriptive Statistics of Simulated Equity and Foreign Exchange Series

This table compares cross-sectional moments of simulated equity and foreign exchange series (Panel A) with their empirical counterparts (Panel B). Moments are defined as the cross-sectional mean and standard deviation of the time-series mean, standard deviation, and first-order autocorrelation (AC1) of the following variables: country equity returns denominated in either local currency ($r^{ei}$) or U.S. Dollars ($r^{e\$}$), the price-dividend ratio in log-form ($p-d$), the dividend growth rate ($\Delta d$), the country risk-free rates ($r_f^i$), the log changes in bilateral exchange rates ($\Delta s$) and the interest rate differences between a foreign and the domestic currency ($i^*-i$). The last column reports time-series moments for the United States. The model is calibrated using parameter values reported in Table 3. It is simulated for a population of 45 countries and one million periods. Empirical moments refer to developed countries. Except for the price-dividend ratio, means and standard deviations are annualized and reported in percentage terms.

| Moments | A: Simulation | | B: Data | | |
|---|---|---|---|---|---|
| | Mean | Std | Mean | Std | US |
| $\mathrm{E}(r^{ei})$ | 10.39 | 1.83 | 10.05 | 2.43 | 10.29 |
| $\sigma(r^{ei})$ | 9.93 | 0.84 | 20.38 | 3.73 | 15.23 |
| $\mathrm{AC1}(r^{ei})$ | 0.02 | 0.00 | 0.11 | 0.06 | 0.06 |
| $\mathrm{E}(r^{e\$})$ | 10.36 | 1.83 | 10.34 | 1.75 | 10.29 |
| $\sigma(r^{e\$})$ | 14.88 | 2.37 | 22.84 | 3.83 | 15.23 |
| $\mathrm{AC1}(r^{e\$})$ | 0.00 | 0.00 | 0.07 | 0.05 | 0.06 |
| $\mathrm{E}(p-d)$ | 5.46 | 0.37 | 3.53 | 0.36 | 3.60 |
| $\sigma(p-d)$ | 0.03 | 0.01 | 0.40 | 0.11 | 0.47 |
| $\mathrm{AC1}(p-d)$ | 0.99 | 0.01 | 0.98 | 0.01 | 0.99 |
| $\mathrm{E}(\Delta d)$ | 4.95 | 0.01 | 7.65 | 2.60 | 6.02 |
| $\sigma(\Delta d)$ | 9.96 | 0.81 | 14.79 | 5.18 | 7.18 |
| $\mathrm{AC1}(\Delta d)$ | 0.02 | 0.00 | 0.20 | 0.17 | 0.46 |
| $\mathrm{E}(r_f^i)$ | 4.50 | 1.93 | 5.10 | 2.17 | 5.03 |
| $\sigma(r_f^i)$ | 1.64 | 0.10 | 1.33 | 0.50 | 0.97 |
| $\mathrm{AC1}(r_f^i)$ | 0.99 | 0.00 | 0.84 | 0.15 | 0.97 |
| $\sigma(\Delta s)$ | 12.48 | 1.82 | 10.15 | 2.91 | - |
| $\mathrm{AC1}(\Delta s)$ | -0.01 | 0.00 | 0.00 | 0.09 | - |
| $\mathrm{E}(i^*-i)$ | 0.00 | 1.93 | 0.93 | 2.82 | - |
| $\sigma(i^*-i)$ | 0.19 | 0.04 | 1.03 | 0.37 | - |
| $\mathrm{AC1}(i^*-i)$ | 0.93 | 0.01 | 0.73 | 0.24 | - |

pairwise correlations between the equity and currency factors (reported in the Online Appendix). The model delivers reasonable interest rates, exchange rates, equity, and currency excess returns as can be seen in Table 4 below.

Figure 9 shows what happens when we estimate our empirical specifications

on these simulated data from the model – it is the counterpart to Figure 5, but estimated using simulated data from the model. The vertical axis corresponds to average excess returns from simulated data. The horizontal axis corresponds to predicted excess returns from the World CAPM, the International CAPM, and our three-factor model all estimated using the simulated data, using rolling-windows in order to capture time-variation in quantities and prices of risk. The figure presents a striking demonstration of the difference: the World CAPM and the International CAPM both fit the simulated data rather poorly, while our model accurately describes the cross-section of returns.

Our next step is to attempt to use our model framework to explore the portfolio choice and hedging strategy of global investors.

# 6 Optimal Currency Hedging

Black (1989) derives a striking result: the optimal amount of currency hedging is the same for every investor in the world. It depends on only three variables: the average world market portfolio expected excess return ($\mu_m$), the average world market portfolio excess return variance ($\sigma_m^2$), and the average exchange rate variance ($\sigma_e^2$).

Black (1989) shows that the fraction of exchange risk hedged should be equal to:

$$\frac{\mu_m - \sigma_m^2}{\mu_m - \frac{1}{2}\sigma_e^2}.$$

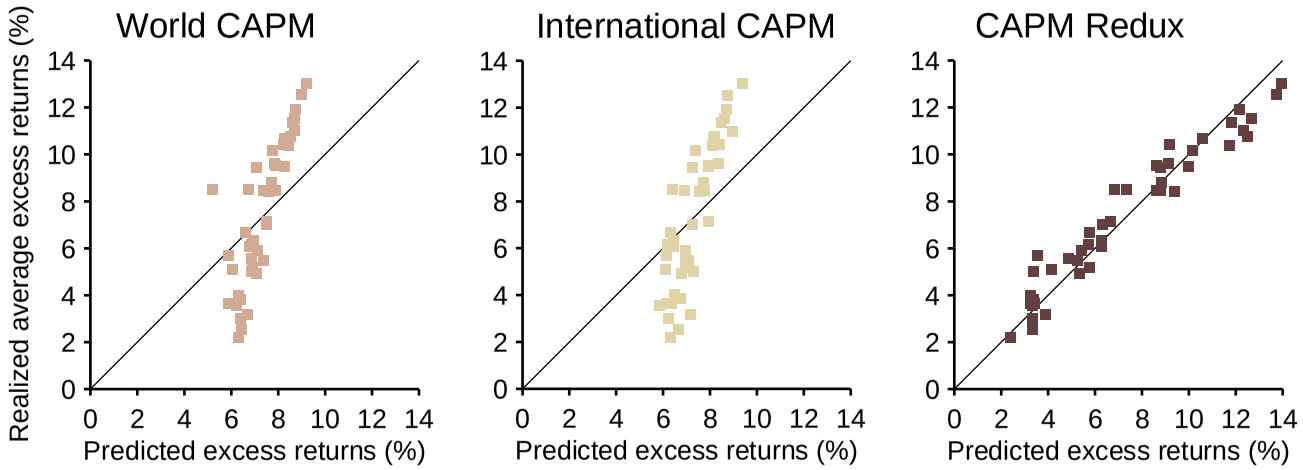
This striking result continues to serve as something of a benchmark to this day. But it is derived in a mean-variance efficient world that does not encompass our model. We thus revisit the issue in our framework, by solving for the optimal portfolio and then inspecting the optimal currency portfolio investment strategy of an atomistic international investor.

## 6.1 Optimal Portfolio Problem

In order to define an optimal portfolio, we naturally need to start with an objective. We interpret each stochastic discount factor of country $i$ as describing the intertemporal marginal rate of substitution of a representative investor in country $i$ characterized by constant relative risk-aversion $\Gamma$.

Figure 9
Realized vs Predicted Average Excess Returns: Simulated Data

This figure plots annualized realized average excess returns against those predicted by the World CAPM, the International CAPM and the CAPM Redux from simulated series. For each country the predicted excess returns are computed as follows: a) Time-varying factor betas are estimated using 60-month rolling windows; b) At each time t estimated conditional betas are multiplied by the corresponding factor means computed over the same time-window (from time *t-59* to time *t*); c) Predicted values are averaged. For each country the realized average excess returns are computed as follows: a) In each 60-month rolling window actual excess returns are averaged and b) 60-month rolling return means are averaged. Average returns are in percentage. The straight line is the 45-degree line through the origin. The sample has length T=447 and a cross-section of 45 countries



Consider the optimal one-period portfolio problem of each investor with initial wealth $W^i$, in the absence of labor income:

$$max_{X^i}\ E[u(X^i)] \text{ subject to } E(M^i X^i) = W^i,$$

where $X^i$ denotes the payoffs next period. The optimal portfolio is therefore (Cochrane, 2005):

$$\widehat{X^i} = W^i \frac{M^{-\frac{1}{\Gamma}}}{E(M^{1-\frac{1}{\Gamma}})}.$$

Solving for the optimal return $\widehat{R^i} = \frac{\widehat{X^i}}{W^i}$ and expressing it in log terms leads to:

$$\widehat{r^i_{t+1}} = -\frac{1}{\Gamma} m^i_{t+1} - \left(1 - \frac{1}{\Gamma}\right) E_t(m^i_{t+1}) - \frac{1}{2}\left(1 - \frac{1}{\Gamma}\right)^2 Var_t(m^i_{t+1}). \tag{9}$$

Substituting in the expressions for our pricing kernels $m$, the optimal return is:

$$\begin{aligned} \widehat{r^i_{t+1}} &= \frac{1}{\Gamma}\left(\sqrt{\gamma z^i_t}u^i_{t+1} + \sqrt{\delta^i z^w_t}u^w_{t+1} + \sqrt{\kappa z^i_t}u^g_{t+1} + \sqrt{\omega z^w_t}u^c_{t+1}\right) \\ &+ \left(\alpha + \chi z^i_t + \tau z^w_t\right) - \frac{1}{2}\left(1 - \frac{1}{\Gamma}\right)^2 \left(\gamma z^i_t + \delta^i z^w_t + \kappa z^i_t + \omega z^w_t\right). \end{aligned} \tag{10}$$

The optimal return depends on the three global shocks in the model as well as the country-specific shock $u^i$.

## 6.2 Optimal Portfolio Solution and Hedging

Let us assume that each investor has access to a risk-free asset and four ETFs which replicate risky returns. In the case of the U.S. investor, the four ETFs are $LWMKT$, $Dollar$, $Carry$, and $MKT_{US}$ (the U.S. stock market return in U.S. Dollars). A U.K. investor would similarly have access to $LWMKT$, $Pound$, $Carry$, and $MKT_{UK}$. We find the optimal wealth allocation ($\omega^{LWMKT}$, $\omega^{Dollar}$, $\omega^{Carry}$, $\omega^{MKT_i}$, $\omega^{rf_i}$) by assuming that each investor wants to replicate the unexpected component of the optimal return by investing in all these five assets. The optimal wealth allocation therefore satisfies:

$$\begin{aligned} \widehat{r^i_{t+1}} - E_t\left(\widehat{r^i_{t+1}}\right) &= \omega^{LWMKT} LWMKT_{t+1} + \omega^{Dollar} Dollar_{t+1} + \omega^{Carry} Carry_{t+1} \\ &+ \omega^{MKT_i} r^{MKT,i}_{t+1} + \omega_{rf_i} r^i_{f,t} \\ \text{where } \omega^{LWMKT} &+ \omega^{Dollar} + \omega^{Carry} + \omega^{MKT_i} + \omega^{rf_i} = 1. \end{aligned}$$

We ignore the small expected component of the optimal return, but it could be replicated by adding for example a bond to the portfolio.[36] Since all returns are known in closed-form, we find that the optimal wealth allocation for the U.S. investor at time $t$ is:

$$\omega_t^{LWMKT} = \frac{\frac{\sqrt{\gamma}\sigma_D^c}{\Gamma}\left(\sqrt{\kappa z_t} - \left(1 - \frac{\sqrt{\omega}}{\sqrt{\gamma}\sigma_D^c}(\sigma_D - k_1 B_{pd}\sigma)\right)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) - \frac{\sqrt{\omega}}{\sigma_D^c}\sigma_D^g\sqrt{z_t}\right)}{\overline{\sigma_D^{c,i}}(\sigma_D - k_1 B_{pd}\sigma)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) - \sqrt{\gamma}\sigma_D^g\left(\overline{\sigma_D^{c,i}}\sqrt{z_t} - \sigma_D^c\overline{\sqrt{z_t^i}}\right)},$$

$$\omega_t^{MKT} = \frac{\frac{1}{\Gamma}\sqrt{\omega} - \omega_t^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c},$$

$$\omega_t^{Dollar} = \frac{1}{\Gamma} - \frac{1}{\sqrt{\gamma}}\left(\frac{\frac{1}{\Gamma}\sqrt{\omega} - \omega_t^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c}\right)(\sigma_D - k_1 B_{pd}\sigma),$$

$$\omega_t^{Carry} = \frac{\frac{1}{\Gamma}\sqrt{\delta} - \omega_t^{LWMKT}(\sigma_D^w - k_1\overline{C_{pd}^i}\sigma^w) - \omega_t^{MKT}(\sigma_D^w - k_1 C_{pd}\sigma^w)}{\left(\overline{\sqrt{\delta^i}}^L - \overline{\sqrt{\delta^i}}^H\right)},$$

$$\omega_t^{rf} = 1 - \omega_t^{LWMKT} - \omega_t^{Dollar} - \omega_t^{Carry} - \omega_t^{MKT}.$$

Note that these shares of wealth in the different assets are time-varying. Date by date, the investor can replicate the unexpected component of the optimal portfolio. Doing so, however, requires the estimation of the model and a measure of the country-specific market price of risk (the variable $z$ in the example of the U.S. investor).

In this framework, optimal portfolio returns are clearly defined, but in the absence of additional frictions, optimal currency hedging is not. The same global shocks affect both equity and currency returns. Investing without any exchange rate risk, for example in a world index of local-currency equity returns, would still expose the investor to many of the same shocks that affect exchange rates. It is thus not clear why a given shock should be hedged when it affects exchange rates but not when it affects equity returns.

One can, however, distinguish between the ETFs that are explicitly investing in currency portfolios versus those that are investing in global or local markets

[36] Our basis of assets is incomplete as it does not replicate the expected component of the optimal return. Moreover, we do not impose the market clearing conditions (equity should be in strictly positive supply and currency and bonds in net zero supply). The basis of assets is here purely illustrative.

without any exchange rate risk. We define the following hedge ratio to summarize this allocation:

$$\text{Hedge ratio}_t = \frac{\omega_t^{LWMKT} + \omega_t^{rf} + \omega_t^{MKT}}{\omega_t^{LWMKT} + \omega_t^{Dollar} + \omega_t^{Carry} + \omega_t^{rf} + \omega_t^{MKT}} = \omega_t^{LWMKT} + \omega_t^{rf} + \omega_t^{MKT}.$$

It is the ratio of the optimal exchange rate insensitive component of the portfolio to the total size of the portfolio. In our setup, this fraction is investor- and time-specific, which contrasts with the universal hedging result of Black (1989). While the optimal portfolio return does not depend on the set of traded assets but only on the risk-aversion and SDF of the representative investor in each country, the hedge ratio depends on the set of traded assets used to implement the optimal portfolio.

Our framework abstracts from real-word frictions that could be added at the cost of loosing the tractability of closed-form results. Through simulations, the model could be extended notably to add short-sale and base-currency constraints.

# 7 Conclusion

In this paper, we show that currency risk is priced in international equity portfolios. We do so using a new set of three factors, namely, a global equity factor denominated in local currencies, and two currency factors, dollar and carry, which are constructed from the space of global currency excess returns. Our results on the pricing of currency risk using this model are obtained from a comprehensive set of equities from 46 developed and emerging countries spanning value, growth, and country index returns from 1976 to the present. We also find that our model outperforms the World and International CAPM, as well as the Fama-French three and four factor models in our sample, and find evidence of substantial exposure to currency risk in the universe of global mutual funds and hedge funds, suggesting that whether exposure to international markets is direct or indirect, currency risk is a significant factor affecting the returns of international investors.

We set up a simple complete-markets model of international equities with endogenous exchange rates, and show that this model is well able to replicate our key empirical findings. The model also enables us to derive the optimal asset

portfolio weights for an atomistic international investor. In our framework, the optimal investment in currency assets for such an investor varies over time and across investor locations, contrary to the famous result of Black (1989) who operates in a simpler mean-variance preference setting.

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# Appendix

## A A Brief Review of the Literature on International Asset Pricing

Early theoretical work in international asset pricing generally assumed that consumption and investment opportunity sets do not differ across countries. This restrictive assumption was relaxed in later work, which eliminated the assumption of investor indifference between domestic and foreign markets.[37] With country-specific investor heterogeneity in preferences, or homogenous preferences but differences in relative prices faced by investors in different countries, currency risk matters for asset pricing. Such relative price differences naturally arise in a world in which there are deviations from purchasing power parity (PPP); moreover in such a world, investing in foreign currencies is risky, since adverse exchange rate movements imply low income from foreign investment expressed in domestic currency terms.[38] In this context, Solnik (1974), Sercu (1980), Stulz (1981), and Adler and Dumas (1983) incorporated currency risk into theoretical asset pricing models, allowing for differences across countries in consumption opportunity sets.

Specifically, Solnik (1974) modeled exchange rates as cross-country relative prices of consumption baskets, and Sercu (1980) generalized the approach, allowing stock returns expressed in local currencies to be correlated with exchange rates. Stulz (1981) and Adler and Dumas (1983) assumed stochastic country-specific inflation in addition to deviations from PPP.[39]

[37] See Glen and Jorion (1993) and Stulz (1995) for a comprehensive survey of the theoretical and empirical literature on international portfolio choice and asset pricing. More recently, Campbell, De Medeiros and Viceira (2010) consider optimal currency hedging in international equity investment, while Kroencke, Schindler and Schrimpf (2013) study the optimal portfolio allocation across equity and exchange rate investment styles.

[38] Empirical work suggests that PPP holds only in the long run. See Froot and Rogoff (1995) for a comprehensive survey of the early literature on PPP.

[39] In the model of Adler and Dumas (1983), there are $L+1$ countries, and a set of $m = n+L+1$ assets – other than the base-currency deposit – comprised of $n$ portfolios of equities, $L$ foreign currency deposits, and the world market portfolio. In that model, the expected return on asset $j$ is:

$$E_t\left(r_{j,t+1} - r_{f,t}\right) = \sum_{i=1}^{L} \lambda_{i,t} cov_t(r_{j,t+1}, r_{n+i,t+1}) + \lambda_{m,t} cov_t(r_{j,t+1}, r_{m,t+1}) \qquad \text{(C.0)}$$

where $r_{j,t+1} - r_{f,t}$ is the nominal return on the equity portfolio $j$ in excess of the risk-free rate

Early empirical work in international asset pricing generally extended the standard Capital Asset Pricing Model (CAPM) (Sharpe (1964) and Lintner (1965)) to global markets, or simply augmented the domestic CAPM with the addition of a few international factors. In this early literature, unconditional tests of these models generally proved inconclusive (see, for example, Stehle Richard (1977), Solnik (1974) and Korajczyk and Viallet (1989).

Later conditional studies yielded more promising results: for example, Bekaert and Harvey (1995) provided evidence that countries' capital markets become increasingly globally integrated over time, Chan, Karolyi and Stulz (1992) identified a time-varying global market premium in U.S. equity markets, and Karolyi and Stulz (1996) studied co-movement in Japanese and U.S. stock markets.[40] Ferson and Harvey (1993) studied the predictability of foreign equity returns and showed that most of it is related to time-variation in global risk premia, and Bekaert and Hodrick (1992), and Bekaert (1995, 1996) studied the predictability of equity and currency returns.

Harvey (1991) introduced a novel approach to the literature, modelling time-variation in both the exposure and the price of risk by conditioning on common and country-specific set of instruments, and finding that time-variation reveals differences across countries, but fails to fully predict conditional expected returns. Implementing a variation on this methodology, and incorporating currency risk explicitly, Dumas and Solnik (1995) found that an exchange rate risk model outperforms a simple "World CAPM" with global market equity returns but no explicit role for currency risk. In their empirical work, Dumas and Solnik (1995) assumed that currency risk for a U.S. investor is well captured by three major currencies: the Japanese Yen, the British Pound, and the German Mark. Providing evidence that currency

(denoted $r_{f,t}$) of the currency in which returns are measured, and $r_{m,t+1}$ is the excess return on the world market portfolio. The covariance terms $cov_t(r_{j,t+1}, r_{n+i,t+1})$ measure the quantity of exchange rate risk. The time-varying coefficients $\lambda_{i,t}$ are the world prices of exchange rate risk. The time-varying coefficient $\lambda_{m,t}$ is the world price of market risk. The unconditional version of this model is presented later in this section.

[40] A large related literature studies the integration of emerging and developed equity markets [see, e.g., Bekaert and Harvey (2000), Bekaert, Harvey and Lundblad (2003), Bekaert, Harvey, Lundblad and Siegel (2007b), Bekaert, Hodrick and Zhang (2009), Bekaert, Harvey, Lundblad and Siegel (2011), and Bekaert and Harvey (2014)].

risk is priced in global capital markets, they attributed the widespread failure of the World CAPM to a misspecification problem. These findings were subsequently corroborated by De Santis and Gerard (1998), who extended the analysis to account for volatility dynamics. Harvey, Solnik and Zhou (2002) also provided support for currency risk in international equity returns. Using latent factors, they find that their first factor premium resembles the expected return on the world market portfolio, while their second factor premium is related to foreign exchange risk. These studies provided strong empirical support for the predictions of international asset pricing theory, but were generally implemented on a small sample of assets from developed countries, over relatively short sample periods.[41]

More recently, a variety of alternate explanations have been proposed for differences in international average equity returns (see Lewis (2011) for the most complete and recent survey of international asset pricing). These explanations include global economic risks Ferson and Harvey (1994); inflation risk Chaieb and Errunza (2007); liquidity risk (Karolyi, Lee and van Dijk, 2012; Bekaert, Harvey and Lundblad, 2007a; Malkhozov, Mueller, Vedolin and Venter, 2014) factors including momentum and a global cash-flow-to-price factor (Hou, Karolyi and Kho, 2011); and investability restrictions (Karolyi and Wu, 2014). These specifications do not account for currency risk; nevertheless they significantly outperform the World CAPM. This is also a feature of the work of Fama and French (2012), who present international versions of both the Fama and French (1993) three-factor model, and the Carhart (1997) four-factor model. Fama and French (2012) find that their global models perform reasonably well in unconditional time-series tests on global size and book-to-market sorted portfolios as well as size and momentum sorted portfolios, but perform worse when the same portfolios are constructed at a regional level (i.e., North America, Japan, Asia Pacific, and Europe).

It is worth mentioning here that our work has close links with recent advances in research on currencies. It is well-known (see Meese and Rogoff (1983) that

[41] Dumas and Solnik (1995) considered a set of four countries (United Kingdom, Germany, Japan and the United States) from March 1970 to December 1991. De Santis and Gerard (1998) take into account the same countries from June 1973 to December 1994.

economic models of exchange rate determination generally lack empirical support in the short-run, with few identifiable links between nominal exchange rates and economic fundamentals.[42] However, a related literature has attempted to explain returns on portfolios of currencies as compensation for risk. Lustig, Roussanov and Verdelhan (2011) show that the cross-section of currency portfolios sorted by interest rates can be well-explained by a "slope" factor – which corresponds to the "carry" factor that we employ in this paper. Lustig, Roussanov and Verdelhan (2011), Menkhoff, Sarno, Schmeling and Schrimpf (2012), Maggiori (2012), and Lettau, Maggiori and Weber (2014) link the currency factors to measures of volatility and downside risk in equity and currency markets. Verdelhan (2015) shows that the carry factor and the "dollar" factor together account for a large share of the variation in bilateral exchange rates, and provides evidence that dollar risk is priced in the cross-section of currencies. These two currency risk factors are nearly orthogonal to one other, and capture two distinct sources of currency risk relevant for explaining variation in currency returns. Our use of carry and dollar as risk-factors in our international asset pricing model is a product of these applications of risk-based arbitrage pricing Ross (1976) to currency markets.

# B Factor Mimicking Portfolio

Figure B.1 reports the characteristics of the a factor mimicking portfolio for the global equity factor denominated in local currencies.

Table B.1 compares the pricing errors obtained with the LWMKT factor or its factor mimicking portfolio.

[42] At high frequencies, a number of papers, including Evans and Lyons (2002), and Froot and Ramadorai (2005) show that order flow in exchange rate markets is helpful at predicting and explaining exchange rate movements, but there is considerable debate about the source of this explanatory power, with both rational and behavioral explanations rationalizing these findings.

Figure B.1
Factor Mimicking Portfolio For LWMKT (Equity Factor)

This figure presents details about a factor mimicking portfolio (FMP) for the global equity factor denominated in local currencies (LWMKT). The FMP is built over 60-month non-overlapping windows in two steps. In the first step LWMKT is regressed on a constant, the excess returns on the eight extreme Fama-French U.S. portfolios and the excess returns on a set of maximum 30 mutual funds with no exposure to the currency factors and high statistical exposure to LWMKT over the same time window. The coefficients are estimated under two constraints: they sum up to one and are all no greater than one in absolute value. In the second step, the FMP is obtained as a weighted average of the excess returns on the same set of traded assets, where weights are given by the first-step coefficients. Panel A shows the number of mutual funds embedded in the FMP in each non-overlapping window. Panel B shows the correlation between the FMP series and LWMKT series over the same horizon. Panel C plots the difference in excess returns between FMP and LWMKT at each point in time. The correlation between these two series is reported at the bottom of the graph. The sample period is April 1978 to April 2013.

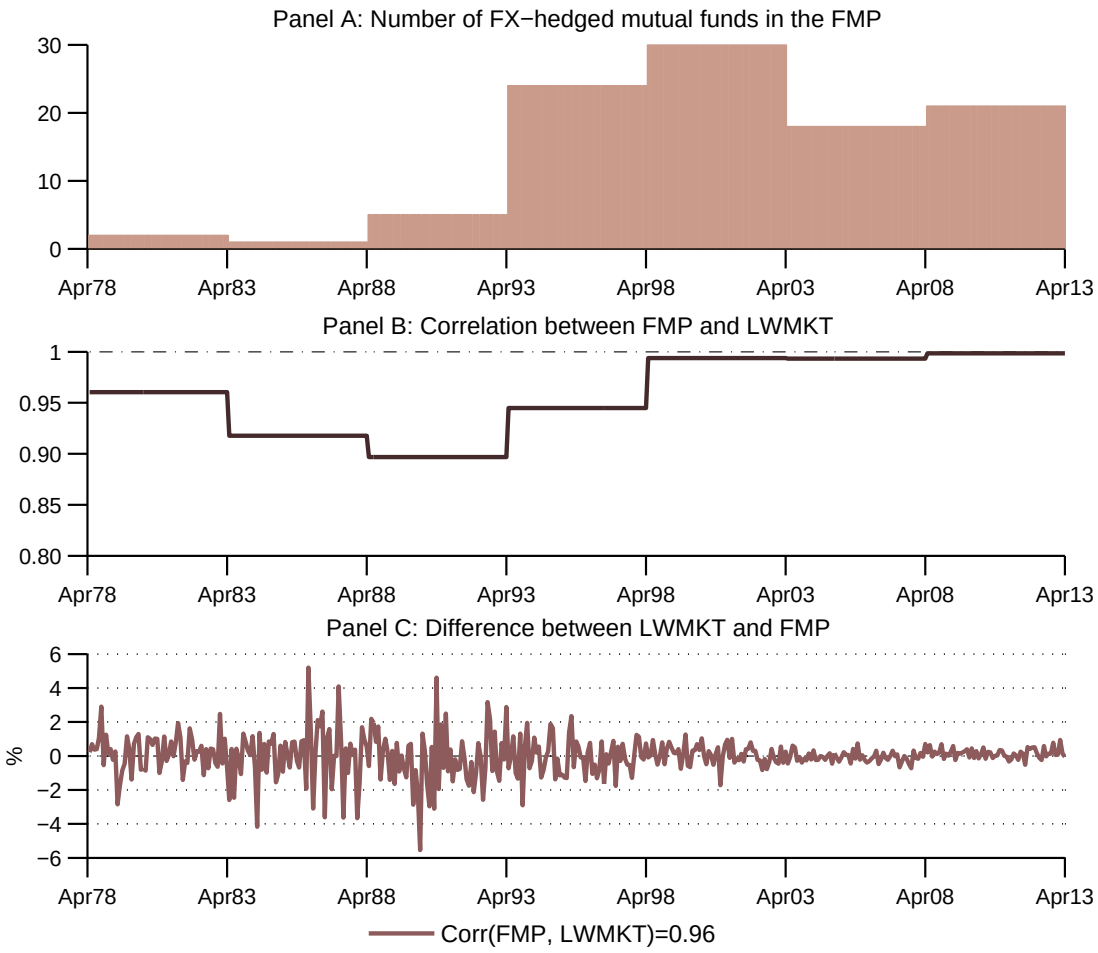


Table B.1
FMP vs. LWMKT: Pricing Error Tests

This table evaluates the CAPM Redux pricing errors when the global equity factor is defined as either the LWMKT factor or its factor mimicking portfolio. Pricing errors are compared along four statistical metrics: the Mean Pricing Error (MPE), the Mean Absolute Pricing Error (MAPE), the Time-Series Mean Absolute Pricing Error (TS-MAPE) and the Root Mean Squared Error (RMSE). For TS-MAPE: At each point in time, we compute the cross-sectional average absolute pricing error and report the time-series average of this statistic. For all other metrics: we compute the metrics at country level and report the cross-sectional average of each statistic. Details about the construction of the FMP are reported in Table B.1. The sample period is April 1978 to April 2013.

| | FMP | LWMKT |
|---|---|---|
| MPE | 0.271 | 0.280 |
| MAPE | 0.782 | 0.777 |
| TS MAPE | 0.538 | 0.540 |
| RMSE | 0.845 | 0.846 |

## C Solving The Model

Under the standard assumption that asset returns and the SDF are jointly log-normal, the Euler equation implies that the expected equity return of any country i, $r^{ei}_{t+1}$, satisfies:

$$0 = \ln E_t \ \exp\{m^i_{t+1} + r^{ei}_{t+1}\} = E_t[m^i_{t+1} + r^{ei}_{t+1}] + \frac{1}{2} Var_t[m^i_{t+1} + r^{ei}_{t+1}]. \quad \text{(C.1)}$$

To solve for the price-dividend ratio of the aggregate dividend claim, we proceed in three steps.

First, we rely on the standard log-linear approximation for the log gross return on that claim:

$$r^{ei}_{t+1} \approx k_0 + k_1 pd^i_{t+1} - pd^i_t + \Delta d^i_{t+1}, \quad \text{(C.2)}$$

where $r^{ei}_{t+1}$ denotes the logarithmic gross real rate of return on each country's stock market index denominated in that country's currency and $k_0$ and $k_1$ are the coefficients for the expansion of the log price-dividend ratio $pd^i_t$ around its mean, $\bar{pd}$[43].

Second, we conjecture that the log price-dividend ratio is affine in the country-specific and the global state variables, $z^i_t$ and $z^w_t$:

$$pd^i_t = A^i_{pd} + B^i_{pd} z^i_t + C^i_{pd} z^w_t. \quad \text{(C.3)}$$

Third, we posit that the dividend growth process of country $i$ follows:

$$\Delta d^i_{t+1} = \mu_D + \psi z^i_t + \psi_w z^w_t + \sigma_D \sqrt{z^i_t} u^i_{t+1} + \sigma^g_D \sqrt{z^i_t} u^g_{t+1} + \sigma^w_D \sqrt{z^w_t} u^w_{t+1} + \sigma^{c,i}_D \sqrt{z^w_t} u^c_{t+1}, \quad \text{(C.4)}$$

Substituting Equation (C.3) and Equation (C.4) into Equation (C.2) yields the following expression for the country-i return in local currency:

$$\begin{aligned} r^{ei}_{t+1} \ \approx \ & k_0 + k_1(A^i_{pd} + B^i_{pd} z^i_{t+1} + C^i_{pd} z^w_{t+1}) - pd^i_t + \mu_D + \psi z^i_t + \psi_w z^w_t + \sigma_D \sqrt{z^i_t} u^i_{t+1} \\ & + \sigma^g_D \sqrt{z^i_t} u^g_{t+1} + \sigma^{c,i}_D \sqrt{z^w_t} u^c_{t+1} + \sigma^w_D \sqrt{z^w_t} u^w_{t+1}. \end{aligned} \quad \text{(C.5)}$$

[43]The coefficients $k_0$ and $k_1$ are given by $k_0 = \log(1 + exp(\bar{pd})) - \frac{\bar{pd}\ exp(\bar{pd})}{1+exp(\bar{pd})}$ and $k_1 = \frac{exp(\bar{pd})}{1+exp(\bar{pd})}$, where $\bar{pd}$ denotes the mean of the log price-dividend ratio.

Combining the law of motion of each state variable with this expression (Equation C.5), substituting back into Equation (C.1) and solving for $pd_t^i$ yield:

$$\begin{aligned} pd_t^i &= -\alpha + k_0 + k_1 A_{pd}^i + k_1 B_{pd}^i(1-\phi)\theta + k_1 C_{pd}^i(1-\phi^w)\theta^w + \mu_D + [k_1 B_{pd}^i \phi + \psi \\ &\quad -\chi + \frac{1}{2}(\sqrt{\gamma} + k_1 B_{pd}^i \sigma - \sigma_D)^2 + \frac{1}{2}(\sigma_D^g - \sqrt{\kappa})^2] z_t^i + \\ &\quad [k_1 C_{pd}^i \phi^w + \psi_w - \tau + \frac{1}{2}(\sqrt{\delta^i} + k_1 C_{pd}^i \sigma^w - \sigma_D^w)^2 + \frac{1}{2}(\sqrt{\omega} - \sigma_D^{c,i})^2] z_t^w. \end{aligned}$$

The linear coefficients $A_{pd}^i$, $B_{pd}^i$ and $C_{pd}^i$ can be obtained by solving 3 equations in 3 unknowns:

$$\begin{aligned} A_{pd}^i &= -\alpha + k_0 + k_1 A_{pd}^i + k_1 B_{pd}^i(1-\phi)\theta + k_1 C_{pd}^i(1-\phi^w)\theta^w + \mu_D, \\ B_{pd}^i &= k_1 B_{pd}^i \phi + \psi - \chi + \frac{1}{2}(\sqrt{\gamma} + k_1 B_{pd}^i \sigma - \sigma_D)^2 + \frac{1}{2}(\sigma_D^g - \sqrt{\kappa})^2, \\ C_{pd}^i &= k_1 C_{pd}^i \phi^w + \psi_w - \tau + \frac{1}{2}(\sqrt{\delta^i} + k_1 C_{pd}^i \sigma^w - \sigma_D^w)^2 + \frac{1}{2}(\sqrt{\omega} - \sigma_D^{c,i})^2. \end{aligned} \quad \text{(C.6)}$$

Solving yields:

$$A_{pd}^i = \frac{-\alpha + k_0 + \mu_D + k_1[B_{pd}^i(1-\phi)\theta + C_{pd}^i(1-\phi^w)\theta^w]}{(1-k_1)}, \quad \text{(C.7)}$$

and two solutions each for $B_{pd}^i$ and $C_{pd}^i$:

$$\begin{aligned} B_{pd}^i &= \frac{1 - k_1(\phi + \sqrt{\gamma}\sigma - \sigma\sigma_D) \pm \sqrt{[1 - k_1(\phi + \sqrt{\gamma}\sigma - \sigma\sigma_D)]^2 + 2k_1^2\sigma^2(\chi ...}}{\kappa_1^2\sigma^2} \\ &\quad \frac{\sqrt{... - \psi - \frac{1}{2}(\gamma + \kappa) - \frac{1}{2}\sigma_D^2 + \sqrt{\gamma}\sigma_D - \frac{1}{2}\sigma_D^{g2} + \sigma_D^g\sqrt{\kappa})}}{\kappa_1^2\sigma^2}, \\ C_{pd}^i &= \frac{1 - k_1(\phi^w + \sqrt{\delta^i}\sigma^w - \sigma^w\sigma_D^w) - \sqrt{[1 - k_1(\phi^w + \sqrt{\delta^i}\sigma^w - \sigma^w\sigma_D^w)]^2 + ...}}{k_1^2\sigma_w^2} \\ &\quad \frac{\sqrt{... + 2k_1^2\sigma_w^2(\tau - \psi_w - \frac{1}{2}\delta^i - \frac{1}{2}\sigma_D^{2,w} + \sqrt{\delta^i\sigma_D^w} - \frac{1}{2}(\sqrt{\omega} - \sigma_D^{c,i})^2)}}{k_1^2\sigma_w^2} \end{aligned} \quad \text{(C.8)}$$

In both cases, we consider only the smallest root, as the other one is not consistent with the assumption that $E[pd^i] = \overline{pd^i}$ .

**Optimal Hedging: Full Solution for the U.S. investor** The set of optimal conditions for the U.S. (domestic) investor is:

$$[1]\quad \omega^{LWMKT}(\sigma_D^w - k_1\overline{C_{pd}^i}\sigma^w) + \omega^{Carry}\left(\overline{\sqrt{\delta^i}}^L - \overline{\sqrt{\delta^i}}^H\right)$$
$$+\omega^{MKT}(\sigma_D^w - k_1 C_{pd}\sigma^w) = \frac{1}{\Gamma}\sqrt{\delta} \qquad [u_{t+1}^w]$$

$$[2]\quad \omega^{LWMKT}\overline{\sigma_D^{c,i}} + \omega^{MKT}\sigma_D^c = \frac{1}{\Gamma}\sqrt{\omega} \qquad [u_{t+1}^c]$$

$$[3]\quad \omega^{LWMKT}\sigma_D^g\overline{\sqrt{z_t^i}} + \omega^{Dollar}\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) + \omega^{MKT}\sigma_D^g\sqrt{z_t} = \frac{1}{\Gamma}\sqrt{\kappa z_t} \qquad [u_{t+1}^g]$$

$$[4]\quad \omega^{MKT}(\sigma_D - k_1 B_{pd}\sigma) + \omega^{Dollar}\sqrt{\gamma} = \frac{1}{\Gamma}\sqrt{\gamma} \qquad [u_{t+1}]$$

$$[5]\quad \omega^{LWMKT} + \omega^{Dollar} + \omega^{Carry} + \omega^{MKT} + \omega^{rf} = 1.$$

That is:

$$[2]\quad \omega^{MKT} = \frac{\frac{1}{\Gamma}\sqrt{\omega} - \omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c},$$

$$[4]\quad \omega^{Dollar} = \frac{1}{\Gamma} - \frac{1}{\sqrt{\gamma}}\left(\frac{\frac{1}{\Gamma}\sqrt{\omega} - \omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c}\right)(\sigma_D - k_1 B_{pd}\sigma)$$

$$[3]\quad \omega^{LWMKT}\sigma_D^g\overline{\sqrt{z_t^i}} + \left(\frac{1}{\Gamma} - \frac{1}{\sqrt{\gamma}}\left(\frac{\frac{1}{\Gamma}\sqrt{\omega} - \omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c}\right)(\sigma_D - k_1 B_{pd}\sigma)\right)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right)$$
$$+\frac{\frac{1}{\Gamma}\sqrt{\omega} - \omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c}\sigma_D^g\sqrt{z_t} = \frac{1}{\Gamma}\sqrt{\kappa z_t},$$

$$[1]\quad \omega^{Carry} = \frac{\frac{1}{\Gamma}\sqrt{\delta} - \omega^{LWMKT}(\sigma_D^w - k_1\overline{C_{pd}^i}\sigma^w) - \omega^{MKT}(\sigma_D^w - k_1 C_{pd}\sigma^w)}{\left(\overline{\sqrt{\delta^i}}^L - \overline{\sqrt{\delta^i}}^H\right)},$$

$$[5]\quad \omega^{LWMKT} + \omega^{Dollar} + \omega^{Carry} + \omega^{MKT} + \omega^{rf} = 1.$$

Solving equation [3] for $\omega^{LWMKT}$ yields:

$$\omega^{LWMKT}\sigma_D^g\overline{\sqrt{z_t^i}} + \left(\frac{1}{\Gamma} - \frac{\sqrt{\omega}}{\Gamma\sqrt{\gamma}\sigma_D^c}(\sigma_D - k_1 B_{pd}\sigma)\right)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) +$$
$$+\frac{\omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sqrt{\gamma}\sigma_D^c}(\sigma_D - k_1 B_{pd}\sigma)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) + \frac{\sqrt{\omega}}{\Gamma\sigma_D^c}\sigma_D^g\sqrt{z_t} -$$
$$\frac{\omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c}\sigma_D^g\sqrt{z_t} = \frac{1}{\Gamma}\sqrt{\kappa z_t}.$$

That is,

$$\omega^{LWMKT}\sigma_D^g\overline{\sqrt{z_t^i}} + \frac{\omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sqrt{\gamma}\sigma_D^c}(\sigma_D - k_1 B_{pd}\sigma)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) - \frac{\omega^{LWMKT}\overline{\sigma_D^{c,i}}}{\sigma_D^c}\sigma_D^g\sqrt{z_t}$$
$$= \frac{1}{\Gamma}\left(\sqrt{\kappa z_t} - \left(1 - \frac{\sqrt{\omega}}{\sqrt{\gamma}\sigma_D^c}(\sigma_D - k_1 B_{pd}\sigma)\right)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) - \frac{\sqrt{\omega}}{\sigma_D^c}\sigma_D^g\sqrt{z_t}\right).$$

That is,

$$\omega^{LWMKT}\left(\sqrt{\gamma}\sigma_D^g\left(\sigma_D^c\overline{\sqrt{z_t^i}} - \overline{\sigma_D^{c,i}}\sqrt{z_t}\right) + \overline{\sigma_D^{c,i}}(\sigma_D - k_1 B_{pd}\sigma)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right)\right)$$
$$= \frac{\sqrt{\gamma}\sigma_D^c}{\Gamma}\left(\sqrt{\kappa z_t} - \left(1 - \frac{\sqrt{\omega}}{\sqrt{\gamma}\sigma_D^c}(\sigma_D - k_1 B_{pd}\sigma)\right)\sqrt{\kappa}\left(\sqrt{z_t} - \overline{\sqrt{z_t^i}}\right) - \frac{\sqrt{\omega}}{\sigma_D^c}\sigma_D^g\sqrt{z_t}\right).$$

This system of linear equations leads to the optimal wealth allocation for the U.S. investor described in the main text.

# 4 | One Central Bank To Rule Them All*

## 1 Introduction

Central bank policies and decisions have a large impact on security markets. Equity prices respond strongly to Federal Open Market Committee (FOMC) interest rate decisions (Bernanke and Kuttner (2005)). Investors also demand a very high risk premium to bear risks associated with FOMC scheduled announcements. Average stock market returns and Sharpe ratios are 20-40 times higher on days when the FOMC is making its scheduled announcements relative to non-announcement days (Savor and Wilson (2013); Lucca and Moench (2015)), an effect that is much larger than for other macroeconomic announcements.[44] Furthermore, on FOMC announcement days, the Capital Asset Pricing Model (CAPM) explains the cross-section of stock returns, with a strong positive relation between market beta and returns (Savor and Wilson (2014)).

In this paper we investigate whether there exists an equivalent announcement-day-premium effect for non-U.S. stocks, both for the Federal Reserve (Fed) as well

*This is joint work with Pavel Savor and Mungo Wilson. Pavel Savor is at Fox School of Business, Temple University. Mungo Wilson is at Saïd Business School, University of Oxford, and Oxford-Man Institute, University of Oxford. We are grateful to Charles Jones and Emanuel Moench for insightful comments. We would also like to thank participants at the 2015 Adam Smith Workshop, Deutsche Bank 3rd Annual Global Quantitative Strategy Conference, the Federal Reserve Board, the 2015 Rodney White Conference, and Washington State University for helpful suggestions.

[44]In addition to FOMC announcements, Savor and Wilson (2013) document a positive risk premium in U.S. equity markets for inflation and employment announcements. Jones, Lamont and Lumsdaine (1998), Savor and Wilson (2013), Faust and Wright (2009), and Balduzzi and Moneta (2015) find positive risk premia in fixed income markets for various macroeconomic announcements. Savor and Wilson (2014) and Mueller, Porchia and Vedolin (2014) find high average returns for different forms of the carry trade on various announcement days (just the FOMC in the latter case).

as non-U.S. central banks. We find a very strong FOMC effect (often significantly stronger than in the U.S.) for almost all international stock markets. For example, while the FOMC-day risk premium, measured as the difference between average market excess returns on FOMC and non-announcement days, is 23.5 basis points (bps) in the U.S., the equivalent premium is 28.6 bps for Japan, 21.5 bps for the United Kingdom, 14.5 bps for Germany, 43.8 bps for Brazil, and 28.1 bps for South Africa.[45] The overall global average market excess return is 27.6 bps on FOMC days compared to 2.7 bps on non-announcement days, and the difference is positive (and significant) in 37 (27) out of 38 countries. The effect is substantially stronger in more recent years, with FOMC days accounting for the bulk of the excess return earned by equity investors over the 1998-2013 period. At the same time, we find virtually no effect for the announcements by non-U.S. central banks, either in the U.S. stock market or, more interestingly, in their home markets. Investors in Europe, the United Kingdom, and Japan seem to demand a higher risk premium for risks associated with FOMC decisions than those by their own central banks.

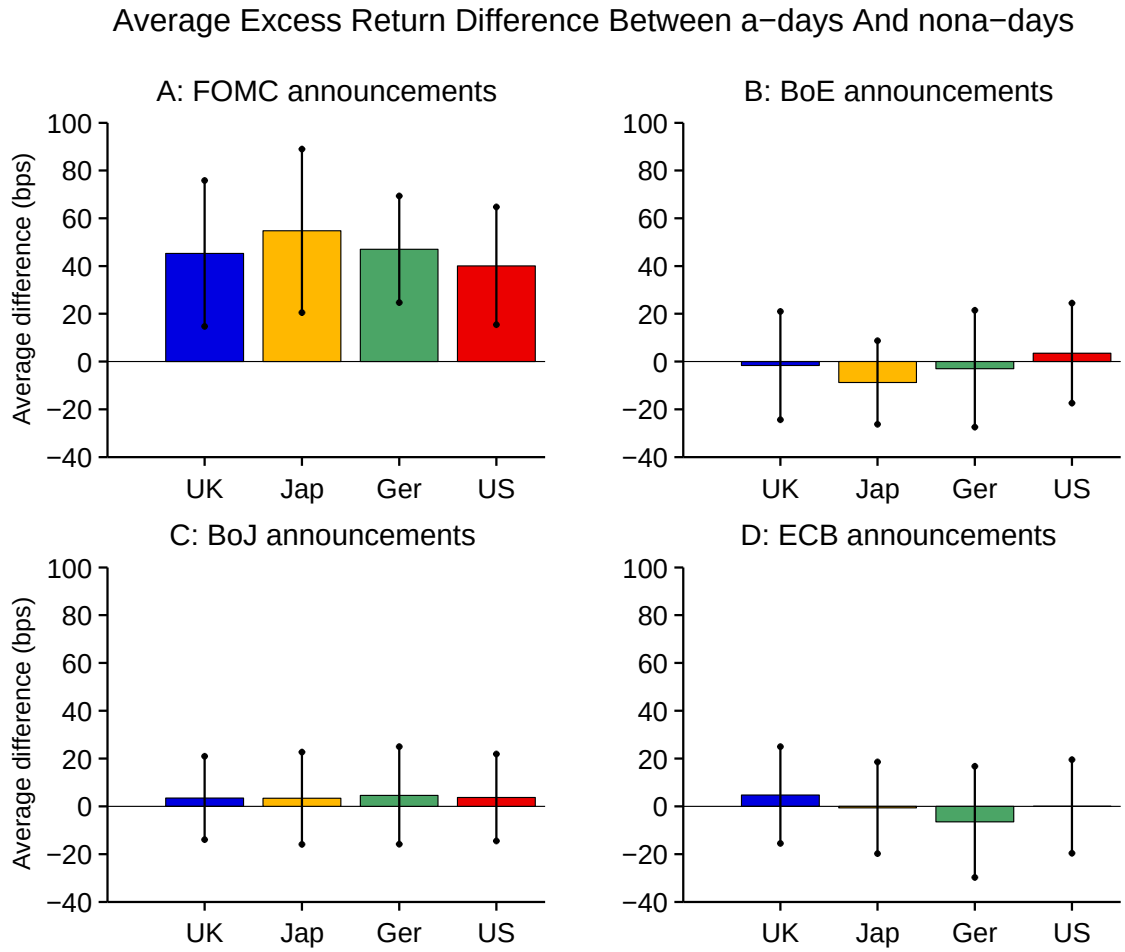
Figure 1 shows our two main results. During the 1999-2013 period, the FOMC, the Bank of Japan (BoJ), the European Central Bank (ECB), and the Bank of England (BoE) all held regular, prescheduled meetings after which they announced their decisions about current monetary policy, including changes to target interest rates, monetary aggregates (in the case of the ECB), and planned asset purchases. All four central banks pursued an active monetary policy over this period, and all four currency zones possess large, liquid, and important equity markets, both by turnover and market capitalization. Each panel shows the difference between average stock market returns (Germany's, in the case of the ECB) on the relevant central banks' scheduled announcement days and non-announcement days for all four currency zones (95% confidence intervals are indicated with black vertical lines).

In the top left panel, we can see a strong FOMC effect on average returns for all four stock markets. U.S. equity returns are on average just under 40 bps higher on FOMC announcement days. This effect is even larger for the other three markets:

[45] We define non-announcement days as days with no scheduled inflation, unemployment, or FOMC announcements.

Figure 1
Summary Chart

This chart reports the difference in daily country average excess returns (in basis points) on announcement (a-days) and non-announcement (nona-days) days in the United States, United Kingdom, Germany, and Japan. Announcement days are those trading days when interest rate decisions of the central bank mentioned at the top of each panel are scheduled for release. The non-announcement sample excludes all the monetary and economic announcements made by each of the four countries. In evaluating the impact of foreign a-days on the Japanese stock market, a-days are shifted to the next day to account for time-differences in trading. Black arrows denote confidence intervals (+/-2 standard deviations). Test assets are Datastream Global Equity Indices denominated in U.S. Dollars. The sample period is January 1999 to December 2013.




over 40 bps for the U.K. and Germany and nearly 60 bps for Japan. In the next section, we document that this effect is also strong for almost every other stock market in the world. As the other three panels show, for the other central banks there is no significant effect, either statistically or economically.

Our results raise an important question: Why is the Fed so powerful compared to other central banks, even those in charge of monetary policy for comparable economies? We discuss a number of potential explanations.

One straightforward possibility is that the domestic economy associated with some central banks is too small for macroeconomic announcements to affect risk premia, even in their own markets. This argument potentially holds for countries such as Switzerland or even the U.K., whose stock market indices contain a large share of multinational firms that are not overly exposed to the local economy. However, it is a less plausible hypothesis for a country like Japan or a whole region like the euro-zone. Furthermore, when we investigate other non-U.S. macroeconomic announcements, such as inflation and employment, we find in many cases a significant effect on average domestic stock market returns, as well as a non-trivial effect on U.S. market returns. We therefore conclude that economy size does not explain our results.

Even if non-U.S. central banks potentially matter to investors, perhaps they choose not to pursue an active monetary policy (or at least have not done so in the past), thereby greatly reducing investor uncertainty regarding banks' decisions. Contrary to this hypothesis, in the period we study all three of the world's non-U.S. major central banks at times exhibit an avowedly activist policy, with large variations in interest rates as well as use of unconventional monetary policy tools. BoE rates ranged from 0.5 to 7.5%, with two tightening and three easing cycles. BoE also engaged in multiple rounds of quantitative easing (QE), purchasing both government and (in smaller amounts) private securities. The interest rates set by the ECB ranged from 0.25% to 4.75%, with three cycles of tightening and easing. While its QE program started after our sample period, the ECB did use various unconventional monetary policy tools, including a number of asset purchase programs and long-term financing operations. The BoJ rates were in a reasonably tight range of 0 to 0.5%, but it engaged in a number of unconventional policies, including an early QE and other asset purchase programs.

Another related explanation is that other central banks actually pursue an active policy, but that their decisions are widely anticipated in advance, whether by design or inadvertently, so that their is little scope for the markets to be surprised by their scheduled announcements.[46] While we were not yet able to collect a comprehensive dataset on forecasts or futures for non-U.S. central banks, anecdotal evidence suggests that BoE, ECB, and BoJ all surprise investors reasonably frequently. Just over the last three years of our sample, we identify through Google searches multiple surprise decisions by all three central banks, which resulted in significant moves in equity, fixed income, and currency markets.

Finally, the Federal Reserve may simply be unique among the world's central banks. One interesting, though at this stage speculative, possibility is that the Federal Reserve sets the global price of money, at least money used to finance securities trading, and that other central banks are just a sideshow.[47] However, if this is the case, it is a puzzling finding. After all, some of these other economies are large and important global players: the combined GDP of the eurozone countries, for example, was actually larger at the end of 2013 than that of the U.S., and both economic regions have open economies and active capital markets. Under this explanation, our results highlight the role of the dollar investor as the marginal investor in global stocks, U.S.- and foreign-traded alike. Another reason for why the Fed is special may be its capabilities. Its researchers and proprietary data sources (see, e.g., Cieslak et al. (2015)) could provide it with better insight into global economic prospects than any of its peers, making its pronouncements much more important to investors locally and globally.

The rest of the paper is organized as follows. Section II presents our results on the impact of U.S. announcements on the rest of the world. Section III gives results for non-U.S. announcements. Section IV concludes and discusses directions for further research. Details about data, such as index construction, announcement

[46]Cieslak et al. (2015) find evidence consistent with the hypothesis that the Fed often releases important information outside its regular meetings.

[47]This may occur because the US dollar's role as the dominant currency for trade (see, e.g., Gopinath (2015)) gives the Fed special freedom of action.

times and dates, and time and date alignment across time zones are only briefly described in the text but are available in the Appendix.

# 2 U.S. Announcements and Global Stock Markets

In this section we examine the effect of U.S. macroeconomic announcements on stock market average excess returns around the world. We first present results for inflation, unemployment, and FOMC announcements separately, and then we focus on FOMC announcements.

Daily total equity return series (denominated in U.S. dollar or domestic currency) are from Datastream Global Equity Indices, with coverage starting in January 1973 for those countries with the longest available time series and ending in December 2013. Where available, we supplement this data with other Datastream return indices for earlier periods (and in one case with data from Global Financial Data).[48] Daily risk-free rates come from Professor Kenneth French's website. Thirty-eight countries are assigned into one of four groups: Europe (17 countries), North America (2 countries), Asia-Pacific (5 countries), and Emerging (14 countries).

U.S. announcement dates are defined as in Savor and Wilson (2013), updated to the end of 2013. Inflation and unemployment announcements occur each month, with dates available from the Bureau of Labor Statistics. The dates of the eight annual scheduled FOMC meetings come from the Federal Reserve. Before February 1994, we assume the FOMC decision became public one day after its meeting (as in Kuttner (2001)). We exclude any unscheduled announcements,

An important issue is to establish which trading day, the same day or the following day, is the first day on which investors can respond to a U.S. announcement in a particular market. To do this, we collect data on time differences relative to the U.S. and on market-open hours for each of the 38 countries in our sample. Because local clocks lose or gain time at various dates in the year (as in, for

[48]The details about when data coverage starts for various countries are given in Table A.2 in the Appendix (only for those countries where it differs from Datastream Global Equity Indices coverage).

example, daylight saving time in the U.S.) and because these timing conventions have changed at various points in our sample period, this is not a trivial task. Table A.1 in the Appendix reports which day (same or next day) represents the effective announcement day for all international stock markets in our sample. Our data on time differences is available in an online Appendix.

## 2.1 Full Sample: 1973-2013

Table 1 shows the impact of the three U.S. macroeconomic announcements on returns in the 38 countries in our sample. Each set of three columns reports the mean announcement-day average excess return, the number of observations, and the difference between average announcement- and non-announcement-day excess returns. The first set does so for FOMC announcements, the second for employment announcements, and the third for inflation announcements.[49] The last column gives the average non-announcement-day excess return (i.e., returns on days other than our U.S. announcement days). $t$-statistics are in brackets, with significant return differences noted in bold font.

Panel B, the North American countries, shows that the FOMC-day average excess return in the U.S. is 24.8 bps, which is 23.5 bps higher than on non-announcement days (n-days) and represents a statistically significantly difference, with a $t$-statistic of 3.75. This result replicates the finding in Savor and Wilson (2013) that FOMC days command a positive risk premium (also in Lucca and Moench (2015)). As in those papers, we find that the n-day average excess return is only 1.3 bps and is not statistically different from zero. Return volatility is also higher on FOMC days than on n-days, but the difference, though statistically significant, is economically marginal (108.7 vs. 106.2 bps). Consequently, Sharpe ratios are much greater on FOMC days (3.6 annualized) than on n-days (0.2 annualized).

[49] A very large literature studies the relation between different macroeconomic surprises and asset prices. See Schwert (1981), Pearce and Roley (1983), Pearce and Roley (1985), Hardouvelis (1987), Cutler et al. (1989), Orphanides (1992), McQueen and Roley (1993), Krueger (1996), Fleming and Remolona (1997), Balduzzi et al. (2001), Flannery and Protopapadakis (2002), Bomfim (2003), Gurkaynak et al. (2005), Boyd et al. (2005), Andersen et al. (2007), and Brenner et al. (2009).

Table 1

Country Daily Stock Market Excess Returns: U.S. Announcements

This table reports daily country average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T). Announcement days are those trading days when either FOMC interest rate decisions, employment numbers, or PPI numbers are scheduled for release in the United States. *t*-statistics are in square brackets. For each country, daily market excess returns are computed as the difference between country market returns denominated in U.S. dollars and the U.S. risk-free rate. Equity series are Datastream Global Country (total return) Indices obtained from Datastream; the daily risk-free rate is obtained from Kenneth French's website. The full sample period is January 1973 to December 2013; data coverage varies across countries. Countries are grouped into four geographical areas. Positive and statistically significant return differences are in bold.

| Country | FOMC | | | Employment | | | Inflation | | | nona-days | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | a-days | | Diff. | a-days | | Diff. | a-days | | Diff. | | |
| | $\mu$ | T | | $\mu$ | T | | $\mu$ | T | | $\mu$ | T |
| A: Europe | | | | | | | | | | | |
| Austria | 15.23 | 309 | **13.01** | -2.06 | 492 | -4.28 | 2.05 | 489 | -0.17 | 2.22 | 9420 |
| | [2.30] | | **[1.93]** | [-0.40] | | [-0.82] | [0.42] | | [-0.03] | [1.78] | |
| Belgium | 16.27 | 307 | **13.88** | -0.09 | 492 | -2.48 | 2.96 | 490 | 0.57 | 2.40 | 9435 |
| | [2.28] | | **[1.92]** | [-0.02] | | [-0.49] | [0.56] | | [0.10] | [2.01] | |
| Denmark | 18.83 | 308 | **15.34** | -3.09 | 491 | -6.57 | 2.29 | 489 | -1.20 | 3.49 | 9426 |
| | [2.49] | | **[2.00]** | [-0.68] | | [-1.38] | [0.48] | | [-0.24] | [2.64] | |
| Finland | 40.73 | 206 | **38.46** | 14.97 | 309 | 12.70 | 0.50 | 308 | -1.77 | 2.27 | 5890 |
| | [3.58] | | **[3.31]** | [1.40] | | [1.16] | [0.05] | | [-0.18] | [0.94] | |
| France | 21.73 | 309 | **18.97** | 0.03 | 492 | -2.73 | 5.38 | 490 | 2.61 | 2.76 | 9429 |
| | [2.92] | | **[2.51]** | [0.01] | | [-0.46] | [0.96] | | [0.45] | [2.00] | |
| Germany | 16.52 | 309 | **14.52** | 1.99 | 491 | -0.01 | 6.64 | 489 | 4.64 | 1.99 | 9435 |
| | [2.63] | | **[2.27]** | [0.38] | | [-0.00] | [1.28] | | [0.87] | [1.56] | |
| Greece | 35.14 | 192 | **34.11** | 22.49 | 288 | **21.47** | -6.43 | 287 | -7.45 | 1.02 | 5494 |
| | [2.49] | | **[2.38]** | [2.20] | | **[2.03]** | [-0.61] | | [-0.68] | [0.39] | |
| Ireland | 14.47 | 309 | 11.28 | -4.77 | 492 | -7.97 | -0.62 | 489 | -3.81 | 3.19 | 9428 |
| | [1.75] | | [1.34] | [-0.87] | | [-1.41] | [-0.10] | | [-0.62] | [2.33] | |
| Italy | 42.33 | 308 | **41.20** | -5.33 | 491 | -6.47 | 0.26 | 489 | -0.88 | 1.14 | 9424 |
| | [4.88] | | **[4.68]** | [-0.83] | | [-0.98] | [0.04] | | [-0.14] | [0.72] | |
| Netherlands | 25.58 | 309 | **23.11** | 0.92 | 490 | -1.55 | 1.17 | 489 | -1.30 | 2.47 | 9442 |
| | [3.84] | | **[3.41]** | [0.18] | | [-0.29] | [0.22] | | [-0.24] | [1.97] | |
| Norway | -1.99 | 284 | -5.84 | 1.04 | 406 | -2.81 | 5.37 | 406 | 1.52 | 3.85 | 7783 |
| | [-0.20] | | [-0.57] | [0.13] | | [-0.34] | [0.71] | | [0.19] | [2.04] | |
| Poland | 15.94 | 156 | 13.35 | 4.86 | 238 | 2.26 | -14.91 | 238 | -17.50 | 2.59 | 4488 |
| | [0.90] | | [0.74] | [0.39] | | [0.18] | [-1.13] | | [-1.30] | [0.84] | |
| Portugal | 32.18 | 191 | **32.16** | 8.08 | 285 | 8.07 | 4.10 | 285 | 4.08 | 0.02 | 5488 |
| | [3.42] | | **[3.36]** | [1.12] | | [1.09] | [0.66] | | [0.63] | [0.01] | |
| Spain | 39.24 | 215 | **37.33** | 6.58 | 322 | 4.67 | 1.36 | 321 | -0.55 | 1.91 | 6135 |
| | [4.15] | | **[3.88]** | [0.87] | | [0.60] | [0.19] | | [-0.07] | [1.04] | |
| Sweden | 37.60 | 258 | **34.16** | 10.65 | 383 | 7.22 | 3.22 | 382 | -0.21 | 3.43 | 7321 |
| | [3.90] | | **[3.48]** | [1.41] | | [0.92] | [0.43] | | [-0.03] | [1.78] | |
| Switzerland | 11.86 | 309 | 9.32 | 3.59 | 492 | 1.05 | 3.50 | 489 | 0.95 | 2.54 | 9433 |
| | [1.96] | | [1.51] | [0.81] | | [0.23] | [0.74] | | [0.20] | [2.27] | |
| United Kingdom | 23.41 | 309 | **21.48** | 7.64 | 492 | 5.71 | 4.27 | 490 | 2.34 | 1.93 | 9446 |
| | [3.28] | | **[2.96]** | [1.41] | | [1.02] | [0.78] | | [0.42] | [1.48] | |

[continued]

| Country | FOMC | | | Employment | | | Inflation | | | nona-days | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | a-days | | Diff. | a-days | | Diff. | a-days | | Diff. | | |
| | $\mu$ | T | | $\mu$ | T | | $\mu$ | T | | $\mu$ | T |
| B: North America | | | | | | | | | | | |
| Canada | 19.39 | 308 | **18.49** | 14.44 | 491 | **13.54** | 5.53 | 487 | 4.62 | 0.90 | 9426 |
| | [2.97] | | **[2.79]** | [3.08] | | **[2.81]** | [1.16] | | [0.94] | [0.82] | |
| United States | 24.79 | 309 | **23.51** | 4.04 | 491 | 2.75 | 8.74 | 490 | 7.45 | 1.28 | 9427 |
| | [4.01] | | **[3.75]** | [0.82] | | [0.55] | [1.68] | | [1.40] | [1.17] | |
| C: Asia Pacific | | | | | | | | | | | |
| Australia | 22.23 | 309 | **19.90** | 7.28 | 492 | 4.95 | -2.93 | 489 | -5.26 | 2.33 | 9443 |
| | [2.68] | | **[2.37]** | [1.13] | | [0.75] | [-0.43] | | [-0.76] | [1.67] | |
| Hong Kong | 26.70 | 309 | **23.33** | 0.17 | 489 | -3.20 | 3.43 | 488 | 0.06 | 3.38 | 9432 |
| | [2.71] | | **[2.33]** | [0.02] | | [-0.31] | [0.40] | | [0.01] | [1.88] | |
| Japan | 29.67 | 309 | **28.64** | -1.96 | 492 | -2.99 | 2.63 | 489 | 1.59 | 1.04 | 9423 |
| | [3.47] | | **[3.32]** | [-0.31] | | [-0.47] | [0.44] | | [0.26] | [0.80] | |
| New Zealand | 26.68 | 208 | **23.44** | -1.15 | 312 | -4.39 | -12.97 | 311 | -16.21 | 3.24 | 5949 |
| | [2.39] | | **[2.08]** | [-0.15] | | [-0.57] | [-1.58] | | [-1.93] | [2.04] | |
| Singapore | 24.15 | 308 | **20.80** | -14.26 | 491 | -17.61 | -16.48 | 489 | -19.84 | 3.35 | 9424 |
| | [2.94] | | **[2.49]** | [-2.01] | | [-2.43] | [-2.19] | | [-2.59] | [2.33] | |
| D: Emerging | | | | | | | | | | | |
| Argentina | 14.10 | 164 | 12.31 | 12.64 | 245 | 10.85 | -3.72 | 243 | -5.51 | 1.79 | 4656 |
| | [1.38] | | [1.16] | [1.14] | | [0.95] | [-0.31] | | [-0.45] | [0.66] | |
| Brazil | 46.44 | 157 | **43.77** | 42.33 | 234 | **39.65** | -8.42 | 233 | -11.10 | 2.68 | 4457 |
| | [2.94] | | **[2.72]** | [2.72] | | **[2.50]** | [-0.57] | | [-0.73] | [0.89] | |
| Chile | 16.44 | 197 | 12.69 | 19.71 | 294 | **15.97** | 7.48 | 294 | 3.73 | 3.75 | 5608 |
| | [1.90] | | [1.44] | [2.98] | | **[2.35]** | [0.99] | | [0.48] | [2.34] | |
| China | 22.83 | 164 | 18.22 | 5.64 | 243 | 1.03 | 5.38 | 245 | 0.76 | 4.62 | 4675 |
| | [1.26] | | [0.99] | [0.36] | | [0.06] | [0.43] | | [0.06] | [1.61] | |
| Indonesia | 66.28 | 190 | **66.15** | -13.34 | 281 | -13.47 | 26.45 | 282 | 26.32 | 0.13 | 5345 |
| | [4.16] | | **[4.06]** | [-1.10] | | [-1.07] | [0.88] | | [0.87] | [0.04] | |
| Malaysia | 41.31 | 224 | **37.42** | -4.24 | 335 | -8.14 | -11.54 | 334 | -15.44 | 3.89 | 6412 |
| | [3.45] | | **[3.08]** | [-0.37] | | [-0.70] | [-1.25] | | [-1.63] | [2.06] | |
| Mexico | 51.89 | 198 | **47.94** | 25.21 | 295 | **21.27** | -2.65 | 295 | -6.60 | 3.94 | 5639 |
| | [3.85] | | **[3.51]** | [2.64] | | **[2.17]** | [-0.27] | | [-0.67] | [1.75] | |
| Philippines | 40.50 | 201 | **37.90** | 15.33 | 300 | 12.73 | -10.26 | 299 | -12.87 | 2.60 | 5745 |
| | [3.81] | | **[3.50]** | [1.29] | | [1.05] | [-1.05] | | [-1.29] | [1.26] | |
| South Africa | 31.45 | 309 | **28.05** | 4.48 | 492 | 1.09 | -8.01 | 489 | -11.40 | 3.40 | 9440 |
| | [3.06] | | **[2.69]** | [0.55] | | [0.13] | [-1.08] | | [-1.49] | [2.02] | |
| South Korea | 56.11 | 212 | **53.94** | 10.70 | 314 | 8.53 | -4.49 | 314 | -6.66 | 2.17 | 5998 |
| | [2.92] | | **[2.78]** | [0.73] | | [0.57] | [-0.30] | | [-0.44] | [0.78] | |
| Taiwan | 5.39 | 205 | 3.06 | 15.81 | 308 | 13.48 | -11.09 | 307 | -13.41 | 2.33 | 5835 |
| | [0.37] | | [0.20] | [1.14] | | [0.96] | [-1.00] | | [-1.18] | [0.95] | |
| Thailand | 39.99 | 217 | **35.39** | 0.85 | 324 | -3.75 | -12.84 | 323 | -17.45 | 4.60 | 6177 |
| | [2.75] | | **[2.40]** | [0.07] | | [-0.32] | [-1.10] | | [-1.46] | [1.89] | |
| Turkey | 30.25 | 197 | 23.97 | 17.96 | 294 | 11.68 | 15.36 | 294 | 9.08 | 6.28 | 5626 |
| | [1.52] | | [1.18] | [1.17] | | [0.74] | [0.99] | | [0.57] | [1.55] | |
| Venezuela | 6.48 | 192 | 0.17 | 35.06 | 287 | **28.75** | 35.71 | 288 | **29.40** | 6.31 | 5483 |
| | [0.36] | | [0.01] | [2.87] | | **[2.26]** | [2.53] | | **[2.02]** | [1.76] | |

Although we estimate a positive premium (defined as the difference between announcement- and non-announcement-day returns) of 2.75 bps for unemployment announcements and one of 7.5 bps for inflation announcements, neither of these are statistically significant. This is consistent with the discussion in Savor and Wilson (2014) that notes the fading away of risk premia associated with inflation and especially unemployment in the U.S. and the gradual increase in the FOMC premium. One potential explanation for these trends is that in the early part of Savor and Wilson's sample (which starts in 1958) market participants used inflation and unemployment numbers to formulate expectations for Federal Reserve policy. After the Fed began to communicate its policy more clearly to the public starting in the mid-1970s and gaining in clarity in the 1990s, the information in unemployment and inflation announcements became less important.

Turning now to our novel results: similar to the American market, the Canadian stock market enjoys a large and significant FOMC premium of 18.5 bps ($t$-statistic=2.79), and its average n-day excess return is low at 0.9 bps and not statistically significant ($t$-statistic=0.82).[50] Panel A reports the same results for our 17 European markets. Of these, all but Norway have a positive FOMC-day premium and all but Ireland, Poland, Switzerland, and Norway a positive and significant one. The significant premia range from 13.0 bps for Austria (Poland's is actually higher, but the Polish market is too volatile for the estimate to be statistically reliable) to 41.2 for Italy. Austria, Belgium, Denmark, France, Germany, the Netherlands, and the U.K. have premia in the range 13-23 bps, while Finland, Greece, Italy, Portugal, Spain, and Sweden have much higher premia in the 34-41 bps range. By contrast, U.S. unemployment and inflation announcement premia are not significant for all but one country (Greece), and are actually frequently negative.[51] As in

[50]Canadian stocks have an insigificant but positive inflation-day premium (4.6, with a $t$-statistic of 0.94) and a strong and significant unemployment-day premium (13.5, with a t-statistic of 2.81).

[51]Many countries (Austria, Belgium, Denmark, France, Germany, Ireland, the Netherlands, Norway, Sweden, Switzerland, and the U.K.) have non-negligible n-day excess returns. There is substantial correlation between a country having a relatively low or absent FOMC premium and a positive n-day average return.

the U.S., return volatility is typically similar on FOMC and n-days, meaning that FOMC-day Sharpe ratios are again orders of magnitude higher.

Panel C shows the results for the developed markets in the Asia-Pacific region. All of these countries have high and significant FOMC announcement premia ranging from 19.9 bps for Australia to 28.6 bps for Japan. None have significantly positive unemployment or inflation premia, with most actually having lower average returns on those days. All of these countries, with the exception of Japan, have significantly positive n-day premia.

Finally, Panel D reports results for emerging markets. Of the 14 countries in this group, a minority (Argentina, Chile, China, Taiwan, Turkey, and Venezuela) do not enjoy statistically significant FOMC premia, but in most of these cases the premia are still quite high (e.g., 24 bps for Turkey) and their lack of significance can be attributed to a shorter sample and higher volatility for those markets. Only Taiwan and Venezuela have economically negligible differences in average excess returns between FOMC and non-announcement days. In those countries where the FOMC premia are statistically significant (Brazil, Indonesia, Malaysia, Mexico, the Philippines, South Africa, South Korea, and Thailand), the magnitude of the premia is greater (and usually much greater) than in the U.S. For example, the FOMC premium in Mexico is 47.9 bps, in South Korea 53.9 bps, and in Indonesia at 66.2 bps. Several countries (Brazil, Chile, Mexico and Venezuela) have significantly higher average returns on U.S. unemployment days, while Venezuela also has higher average returns on U.S. inflation a-days.[52,53]

Overall, the pattern we document is of widespread higher average returns and higher Sharpe ratios on FOMC announcement days. The magnitude of the premium is relatively low in only three countries, all of which enjoy special circumstances: Norway (heavily exposed to oil and gas), Venezuela (transitioning away from a market economy), and Taiwan (a small, isolated, and expensive-to-trade stock

[52]Venezuela is in many ways an anomaly in our sample, probably due to its unique experience. It is likely its returns are significantly mismeasured, given that its currency's official and actual (black market) exchange rates have dramatically diverged during the period we study.

[53]Chile, China, Malaysia, Mexico, South Africa, Thailand, and Venezuela have significantly positive n-day returns.

market). The countries with especially pronounced premia appear to be those that have experienced financial crises during our sample period (Portugal, Italy, Spain and Greece in Europe; Malaysia, Thailand, South Korea, the Philippines, and Indonesia in Asia; and Brazil and Mexico in the Americas).[54]

## 2.2 Recent Sample: 1998-2013

Table 2 presents a similar analysis to Table 1 but just for FOMC announcements for the most recent 15 years of our sample. We focus on this period, since it overlaps with our sample for the three non-U.S. central banks. Although not very long, this period includes many notable financial events, such as the Asian crisis, the LTCM-Russia crisis, the internet boom and bust, the global financial crisis of 2008-2009, and the eurozone crisis.

The FOMC premium in the U.S. (Panel B) is 44.1 bps ($t$-statistic=3.82), which is almost double the 23.5 bps during the full 1978-2013 period. The average n-day excess return is actually negative at -2.1 bps, indicating that the entire 1998-2013 U.S. equity premium was earned on FOMC and other a-days. Even though the Federal Reserve always had a significant impact on the U.S. equity market, its importance increased dramatically in the recent years.

The main take-away from Table 2 is that the FOMC premium is globally very high during the 1998-2013 period, averaging 33.0 bps. Almost all countries show much higher average excess returns on FOMC days than on n-days, the only exceptions being China and Venezuela. Of the rest, the $t$-tests for the difference are not significant for Norway and Poland in Europe, Hong Kong and Singapore in the developed Asia-Pacific, and Argentina, Taiwan and Thailand for emerging markets. However, the magnitudes associated with these countries are all economically large. Norway and Poland's FOMC-day returns are 16.1 and 22.3 bps higher (compared to 1.4 and -1.9 bps on n-days), and Hong Kong and Singapore's are 23.8 and 17.6

[54]Partly this is because we are reporting excess returns in U.S. dollars in this table, and also because for some of these countries our sample period is more recent, when the effect was stronger. However, this pattern partly survives in Table 2, which looks at domestic currency returns in the post-1998 period. Furthermore, the currency effect is interesting in its own right. (See Savor and Wilson (2014) and Mueller et al. (2014).)

Table 2

FOMC Announcements - Post January 1998

This table reports daily country average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T). Announcement days are those trading days when FOMC interest rate decisions are scheduled for release in the United States. Non-announcement days are those days when interest rate decisions, employment numbers and inflation numbers are not scheduled for release in the United States, United Kingdom, Germany, and Japan. In countries denoted by star all announcements except for the Japanese ones are shifted to the next day to account for time-differences in trading. *t*-statistics are in square brackets. Equity series are either Datastream Global Country (total return) Indices or country major stock market indices obtained from Datastream. Countries are grouped into four geographical areas. Positive and statistically significant return differences are in bold. The sample period is January 1998 to December 2013.

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | |
| A: Europe | | | | | |
| Austria | 23.81 | 128 | -0.16 | 2433 | **23.97** |
| | [2.67] | | [-0.06] | | **[2.59]** |
| Belgium | 27.26 | 128 | -0.14 | 2436 | **27.40** |
| | [2.49] | | [-0.06] | | **[2.44]** |
| Denmark | 29.34 | 128 | 1.72 | 2436 | **27.61** |
| | [2.54] | | [0.70] | | **[2.34]** |
| Finland | 53.24 | 127 | -4.72 | 2314 | **57.96** |
| | [3.84] | | [-1.21] | | **[4.03]** |
| France | 46.17 | 128 | -0.44 | 2436 | **46.61** |
| | [4.13] | | [-0.16] | | **[4.05]** |
| Germany | 44.39 | 127 | 0.19 | 2348 | **44.20** |
| | [3.96] | | [0.06] | | **[3.77]** |
| Greece | 33.22 | 125 | -8.18 | 2304 | **41.39** |
| | [1.98] | | [-2.06] | | **[2.41]** |
| Ireland | 32.53 | 128 | -4.43 | 2436 | **36.96** |
| | [2.32] | | [-1.57] | | **[2.59]** |
| Italy | 26.63 | 126 | -3.22 | 2345 | **29.85** |
| | [2.56] | | [-1.23] | | **[2.79]** |
| The Netherlands | 36.65 | 128 | -0.53 | 2436 | **37.18** |
| | [3.55] | | [-0.19] | | **[3.48]** |
| Norway | 17.55 | 127 | 1.41 | 2436 | 16.13 |
| | [1.49] | | [0.46] | | [1.33] |
| Poland | 20.44 | 125 | -1.85 | 2311 | 22.29 |
| | [1.51] | | [-0.60] | | [1.61] |
| Portugal | 27.26 | 126 | -3.47 | 2344 | **30.73** |
| | [2.70] | | [-1.44] | | **[2.96]** |
| Spain | 46.39 | 126 | -2.75 | 2333 | **49.14** |
| | [3.72] | | [-0.88] | | **[3.82]** |
| Sweden | 65.72 | 128 | 0.92 | 2436 | **64.81** |
| | [5.44] | | [0.29] | | **[5.19]** |
| Switzerland | 27.06 | 128 | 0.08 | 2433 | **26.98** |
| | [3.02] | | [0.04] | | **[2.91]** |
| United Kingdom | 33.24 | 127 | -1.42 | 2326 | **34.66** |
| | [3.22] | | [-0.56] | | **[3.26]** |
| B: North America | | | | | |
| Canada | 25.10 | 126 | -3.95 | 2313 | **29.05** |
| | [2.62] | | [-1.59] | | **[2.93]** |
| United States | 41.92 | 128 | -2.13 | 2312 | **44.06** |
| | [3.74] | | [-0.78] | | **[3.82]** |

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | Mean | T | $\mu$ | T | |
| C: Asia Pacific | | | | | |
| Australia* | 24.01 | 128 | 1.58 | 2439 | **22.43** |
| | [2.55] | | [0.79] | | **[2.33]** |
| Hong Kong* | 22.83 | 123 | -1.00 | 2299 | 23.82 |
| | [1.29] | | [-0.28] | | [1.32] |
| Japan* | 47.10 | 123 | -2.21 | 2272 | **49.31** |
| | [3.25] | | [-0.74] | | **[3.33]** |
| New Zealand* | 16.44 | 128 | 0.51 | 2438 | **15.93** |
| | [2.72] | | [0.35] | | **[2.57]** |
| Singapore* | 19.55 | 128 | 1.99 | 2438 | 17.56 |
| | [1.69] | | [0.80] | | [1.48] |
| D: Emerging | | | | | |
| Argentina | 17.56 | 128 | 4.03 | 2416 | 13.53 |
| | [1.57] | | [1.19] | | [1.16] |
| Brazil | 69.02 | 124 | 0.39 | 2294 | **68.63** |
| | [4.14] | | [0.09] | | **[3.97]** |
| Chile | 15.85 | 124 | -0.81 | 2302 | **16.66** |
| | [2.47] | | [-0.47] | | **[2.51]** |
| China* | 4.42 | 113 | 2.84 | 2253 | 1.58 |
| | [0.27] | | [0.83] | | [0.09] |
| Indonesia* | 58.55 | 124 | 3.97 | 2277 | **54.58** |
| | [3.64] | | [1.13] | | **[3.32]** |
| Malaysia* | 34.85 | 123 | 0.71 | 2291 | **34.13** |
| | [3.12] | | [0.26] | | **[2.97]** |
| Mexico | 38.59 | 121 | 1.90 | 2341 | **36.69** |
| | [3.32] | | [0.61] | | **[3.05]** |
| Philippines* | 47.65 | 125 | -0.13 | 2285 | **47.78** |
| | [3.78] | | [-0.04] | | **[3.68]** |
| South Africa | 23.51 | 128 | 3.80 | 2436 | **19.71** |
| | [2.48] | | [1.52] | | **[2.01]** |
| South Korea* | 49.04 | 123 | 0.36 | 2300 | **48.69** |
| | [2.73] | | [0.09] | | **[2.65]** |
| Taiwan* | 26.37 | 117 | -1.74 | 2313 | 28.12 |
| | [1.58] | | [-0.56] | | [1.65] |
| Thailand* | 26.05 | 120 | 2.82 | 2280 | 23.23 |
| | [1.68] | | [0.82] | | [1.46] |
| Turkey | 41.37 | 128 | 1.36 | 2436 | **40.02** |
| | [2.18] | | [0.27] | | **[2.04]** |
| Venezuela | 11.81 | 128 | 10.33 | 2436 | 1.49 |
| | [1.15] | | [3.14] | | [0.14] |

bps higher than their n-day average returns of -1.0 and 2.0 bps. For Argentina, Taiwan, and Thailand, the FOMC-day returns are on average 13.5, 28.1, and 23.2 bps higher than their average n-day returns of 4.0, -1.7, and 2.8 bps respectively. All other countries' returns are significantly higher on FOMC days. Furthermore, on n-days their average excess returns are negligible and even mostly negative: with the sole exception of Venezuela, not a single country has positive n-day average excess returns that are statistically significant.[55]

Not only is the FOMC premium even more widespread in this recent period, but it is also much larger in magnitude. In Europe, the premium ranges from 16.1 bps for Norway to 64.8 bps for Sweden. All countries in Europe, North America, and the developed Asia-Pacific had high and positive FOMC announcement premia. Notably, Japan's FOMC premium is 49.3 bps. The same pattern is observed in emerging markets, with Brazil topping the list at a 68.6 bps increase on FOMC-days, representing a compound annual return of 5.7% just from holding the Brazilian stock market on the eight scheduled FOMC days.

Our estimates imply that a highly disproportionate share of cumulative market returns occurs on announcement days and in particular FOMC days. For example, in the 1998-2013 period, an investor in the German market index would have earned a cumulative excess return of 95.2%, and of this return only 4.9% (i.e., a cumulative return of 4.5%) would be earned on n-days (which account for 88% of all trading days) while 59.2% (i.e., a cumulative return of 56.4%) would be earned on FOMC days, representing just 3.1% of trading days (the remainder are other a-days). In the case of Japan, the figures are even more stark, with a cumulative n-day return that is actually negative, meaning that the cumulative return earned on FOMC days (57.9%) greatly exceeded the total cumulative return (18.8%).

To summarize, over the 1978-2013 period, the FOMC announcement premium is substantial and widespread. In the most recent 15 years, it is very large and almost

[55] 14 out of 17 European countries have average n-day returns that are, to the nearest basis point, zero or negative. Canada, the U.S., Hong Kong, and Japan all have negative average n-day returns. Stock market investors in these markets would mostly have lost money on their investments on n-days.

universal, accounting for the bulk of the excess return earned by equity investors across the globe. The only exceptions are China, which was almost entirely closed to foreign equity market investors in this period, and Venezuela, a near-communist oil producer with a pronounced hostility to foreign investors in general and the U.S. in particular. Our results show that FOMC announcements are a leading driver of stock market returns in every important and investible stock market in the world.

# 3 International Macroeconomic Announcements

In the previous section, we establish that on FOMC announcement days average market returns are near-universally higher across global stock markets, and that this result is especially strong in the more recent 1998-2013 period. By contrast, the premia on days of the other U.S. macroeconomic announcements historically found to be associated with higher average returns (see Savor and Wilson (2013) for stocks and government bonds and Jones et al. (1998) for government bonds) are economically negligible and often negative outside the U.S. (and no longer statistically significant in the U.S.). We now explore the impact of various non-U.S. macroeconomic announcements on global stock returns.

## 3.1 Non-U.S. Central Bank Announcements

We focus on the three major non-U.S. central banks: the Bank of England (BoE), the Bank of Japan (BoJ), and the European Central Bank (ECB) and its predecessor the Bundesbank. By any measure, these three are among the world's most important central banks and are very probably the top three after the Federal Reserve (at least until China's recent emergence). The three central banks have four things in common: they manage interest rates in economies with a freely floating currency; they have independent mandates (politicians cannot directly order them to adopt certain policies, and their senior staff cannot easily be replaced by politicians); these mandates are clearly defined (for example, price stability in the eurozone or the dual mandate of the Federal Reserve) and almost certainly generally understood by market participants (even though they may disagree about what exactly is

meant by these mandates in practice); and they hold regular scheduled meetings to decide policies and reach conclusions about the economy whose outcomes are communicated to the public at a timetable that is published in advance.

Some of these banks have enjoyed independence only recently. The BoE gained independent control of interest rates in June 1997, and the BoJ attained full independence with control of interest rates in 1998. Table A.3 in the Appendix lists the sources for our data on pre-scheduled announcements for these central banks, together with information on their status. This table also details the release dates and data sources for these countries' (or currency zones') other macroeconomic releases. Our sample consists of monthly announcements for the BoJ and BoE starting January 1998 and for the ECB starting January 1999. Prior to January 1999, we have monetary policy announcements by the Bundesbank, which occurred every two weeks beginning in January 1958. Please see the online Appendix for details of the sources and methods for collecting the historical dates of scheduled central banks' announcements.

Obviously, there are many other central banks and monetary authorities beyond the three we study. However, of these others, many do not possess genuine freedom of action, for one reason or another, and so their announcements may be of limited significance to investors.[56] It is also not easy to collect historical announcement dates for many non-U.S. central banks, especially going back more than a few years. It was a non-trivial task to collect the announcement dates for the three central banks we study, and it is likely that the data may be harder to obtain for some other candidates for major central banks (such as the central banks of China or Brazil).

In this section, we define non-announcement days (n-days), for all the markets in our sample, as those trading days on which there are no scheduled central bank,

[56] A few examples: the Hong Kong Monetary Authority manages a currency peg between the Hong Kong and U.S. dollars, and cannot typically offer very different interest rates from those available in the US. The Banque de France in our sample period also managed a currency peg with the German Deutschmark (before the French franc was merged into the euro), and was widely regarded as the less inluential partner in the currency union. The Argentine central bank is widely suspected of suffering from political interference. The People's Bank of China both manages a quasi-peg to the U.S. dollar (or at least did so during most of our sample) and is likely quite exposed to political pressures.

employment, or inflation announcements in the U.S., the U.K., Japan, or Germany (taking the ECB as Germany's central bank). The remaining trading days are announcement days (a-days). As in Tables 1 and 2, we account for time differences across countries, with country market returns coming from Datastream Global Equity Indices. Excess returns are reported in domestic currency.

Table 3 reports results for BoE announcements. In strong contrast to FOMC announcements, average excess returns are not significantly higher on BoE a-days than on n-days for any country in our sample, and for many countries BoE a-day returns are actually lower than on n-days. (Recall that we are excluding all other a-days from n-days, so that both a-day returns for a particular announcement and n-day returns can be negative even if overall cumulative excess returns are not.) Most notably, U.K. stock market average excess returns on BoE a-days are lower than on n-days (-8.9 versus -1.4 bps). Even investors in the U.K. market do not seem to demand a premium to bear risks associated with BoE scheduled announcements.

In Table 4 we see very similar results for the BoJ. BoJ a-days are not associated with significantly higher average returns in 33 out of 38 countries (and are again frequently lower). And even for the five exceptions, the difference is often only marginally significant: Greece (23.2 bps higher, with a $t$-statistic of 1.79), Portugal (11.8 bps higher, with a $t$-statistic of 1.64), Malaysia (18.0 bps higher, with a $t$-statistic of 1.96), New Zealand (12.0 bps higher, with a $t$-statistic of 2.43) and Indonesia (27.8 bps higher, with a $t$-statistic of 2.32). In Japan, BoJ a-day returns are 7.0 bps higher, an economically meaningful magnitude, but the $t$-statistic for the difference is only 0.77. Looking at some of the estimates that are not significant (and with that crucial caveat in mind), the point estimates are quite often higher for Japan's neighbors in the Asia-Pacific (the only exceptions are Taiwan and Hong Kong), which is consistent with the hypothesis that the BoJ may play more of a role in its region. These point estimates are much lower than those for the FOMC announcements in Table 2. If there is any effect for Japan, it is significantly smaller and more muted than the FOMC effect, even in Japan itself. Ironically, Japan

Table 3

Bank of England's Announcements - Post January 1998

This table reports daily country average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T). Announcement days are those trading days when interest rate decisions are scheduled for release at the Bank of England. Non-announcement days are those days when interest rate decisions, employment numbers and inflation numbers are not scheduled for release in the United States, United Kingdom, Germany, and Japan. In countries denoted by star all announcements except for the Japanese ones are shifted to the next day to account for time-differences in trading. *t*-statistics are in square brackets. Equity series are either Datastream Global Country (total return) Indices or country major stock market indices obtained from Datastream. Countries are grouped into four geographical areas. Positive and statistically significant return differences are in bold. The sample period is January 1998 to December 2013.

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | |
| A: Europe | | | | | |
| Austria | -3.13 | 191 | -0.16 | 2433 | -2.97 |
| | [-0.35] | | [-0.06] | | [-0.32] |
| Belgium | -4.39 | 191 | -0.14 | 2436 | -4.25 |
| | [-0.47] | | [-0.06] | | [-0.44] |
| Denmark | -13.46 | 191 | 1.72 | 2436 | -15.18 |
| | [-1.53] | | [0.70] | | [-1.66] |
| Finland | -4.11 | 186 | -4.72 | 2314 | 0.62 |
| | [-0.26] | | [-1.21] | | [0.04] |
| France | -13.31 | 191 | -0.44 | 2436 | -12.87 |
| | [-1.22] | | [-0.16] | | [-1.14] |
| Germany | -15.33 | 191 | 0.19 | 2348 | -15.52 |
| | [-1.25] | | [0.06] | | [-1.22] |
| Greece | 12.05 | 190 | -8.18 | 2304 | 20.22 |
| | [0.91] | | [-2.06] | | [1.46] |
| Ireland | -2.76 | 191 | -4.43 | 2436 | 1.67 |
| | [-0.27] | | [-1.57] | | [0.16] |
| Italy | 0.91 | 191 | -3.22 | 2345 | 4.14 |
| | [0.10] | | [-1.23] | | [0.42] |
| The Netherlands | -12.59 | 191 | -0.53 | 2436 | -12.06 |
| | [-1.15] | | [-0.19] | | [-1.07] |
| Norway | -16.45 | 191 | 1.41 | 2436 | -17.86 |
| | [-1.40] | | [0.46] | | [-1.47] |
| Poland | -0.04 | 188 | -1.85 | 2311 | 1.80 |
| | [-0.00] | | [-0.60] | | [0.16] |
| Portugal | 3.61 | 188 | -3.47 | 2344 | 7.07 |
| | [0.36] | | [-1.44] | | [0.69] |
| Spain | -9.93 | 190 | -2.75 | 2333 | -7.18 |
| | [-0.72] | | [-0.88] | | [-0.50] |
| Sweden | -19.94 | 191 | 0.92 | 2436 | -20.86 |
| | [-1.59] | | [0.29] | | [-1.61] |
| Switzerland | -16.67 | 191 | 0.08 | 2433 | -16.75 |
| | [-2.07] | | [0.04] | | [-2.00] |
| United Kingdom | -8.91 | 191 | -1.42 | 2326 | -7.49 |
| | [-1.01] | | [-0.56] | | [-0.82] |
| B: North America | | | | | |
| Canada | -3.55 | 191 | -3.95 | 2313 | 0.40 |
| | [-0.46] | | [-1.59] | | [0.05] |
| United States | -0.81 | 189 | -2.13 | 2312 | 1.32 |
| | [-0.08] | | [-0.78] | | [0.13] |

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | |
| C: Asia Pacific | | | | | |
| Australia* | -4.88 | 191 | 1.58 | 2439 | -6.46 |
| | [-0.77] | | [0.79] | | [-0.97] |
| Hong Kong* | 4.22 | 180 | -1.00 | 2299 | 5.22 |
| | [0.37] | | [-0.28] | | [0.43] |
| Japan* | -12.76 | 186 | -2.21 | 2272 | -10.55 |
| | [-1.48] | | [-0.74] | | [-1.16] |
| New Zealand* | -0.19 | 190 | 0.51 | 2438 | -0.69 |
| | [-0.04] | | [0.35] | | [-0.14] |
| Singapore* | 3.82 | 190 | 1.99 | 2438 | 1.83 |
| | [0.52] | | [0.80] | | [0.24] |
| D: Emerging | | | | | |
| Argentina | -4.95 | 191 | 4.03 | 2416 | -8.98 |
| | [-0.39] | | [1.19] | | [-0.68] |
| Brazil | 1.08 | 183 | 0.39 | 2294 | 0.70 |
| | [0.07] | | [0.09] | | [0.04] |
| Chile | 2.95 | 189 | -0.81 | 2302 | 3.76 |
| | [0.47] | | [-0.47] | | [0.57] |
| China* | 1.99 | 174 | 2.84 | 2253 | -0.85 |
| | [0.17] | | [0.83] | | [-0.07] |
| Indonesia* | 11.11 | 180 | 3.97 | 2277 | 7.14 |
| | [1.06] | | [1.13] | | [0.64] |
| Malaysia* | 4.34 | 186 | 0.71 | 2291 | 3.62 |
| | [0.55] | | [0.26] | | [0.43] |
| Mexico | -7.95 | 184 | 1.90 | 2341 | -9.85 |
| | [-0.69] | | [0.61] | | [-0.83] |
| Philippines* | 13.64 | 182 | -0.13 | 2285 | 13.78 |
| | [1.38] | | [-0.04] | | [1.33] |
| South Africa | -0.16 | 191 | 3.80 | 2436 | -3.96 |
| | [-0.02] | | [1.52] | | [-0.42] |
| South Korea* | 4.08 | 184 | 0.36 | 2300 | 3.72 |
| | [0.33] | | [0.09] | | [0.29] |
| Taiwan* | 5.78 | 178 | -1.74 | 2313 | 7.52 |
| | [0.58] | | [-0.56] | | [0.72] |
| Thailand* | 12.00 | 178 | 2.82 | 2280 | 9.18 |
| | [1.17] | | [0.82] | | [0.85] |
| Turkey | -1.84 | 191 | 1.36 | 2436 | -3.20 |
| | [-0.10] | | [0.27] | | [-0.16] |
| Venezuela | -4.17 | 191 | 10.33 | 2436 | -14.50 |
| | [-0.39] | | [3.14] | | [-1.28] |

appears to be the land of the setting moon and the U.S. that of the rising sun in the case of central banks and stock markets.

Table 5 presents the equivalent result for the ECB. ECB a-day returns are significantly higher in four out of 38 countries, only one of which is a eurozone country, Finland (32.8 bps higher, with a $t$-statistic for the difference of 1.72), with one additional European country, Poland (20.3 bps higher, with a $t$-statistic of 1.87). There are a few more high point estimates (Greece, Ireland, and Italy), but the associated $t$-statistics are not close to significant levels. In many remaining cases, ECB a-day returns are actually lower than on n-days, including France, Germany, the Netherlands, and Spain. Outside Europe, the ECB a-days are not associated with significantly higher average returns anywhere except Taiwan and the Philippines. Interestingly, unlike Europe itself, the point estimates are usually positive, with the sole exception of Venezuela.

Similar to the impact of BoE and BoJ announcements in their own markets, the announcements of the ECB do not affect risk premia in the central core markets of the eurozone. And even in the few markets where the premium is economically significant, the effect is much smaller than the FOMC effect in the same period. Together, Tables 4, 5, and 6 document an important and potentially puzzling finding: not only do risk premia associated with FOMC announcements dwarf those associated with announcements by major non-U.S. central banks across global markets, they even do so in the non-U.S. banks' home markets.

We also study the effect of Bundesbank announcements earlier in our sample period. Regular Bundesbank monetary policy announcements begin in 1958, even though not many of our stock index return histories extend back so far. Table 6 shows the results for two different sample periods, the entire pre-euro period (ending in December 1998) in Panel I and the 1958-1992 period in Panel II (1992 was the year the Maastricht Treaty establishing the euro was agreed). Again, returns are reported in domestic currency.

Over the full 1958-1998 period, market returns are typically higher on Bundesbank a-days, including for 14 out of 17 European countries, but are not significant

Table 4

Bank of Japan's Announcements - Post January 1998

This table reports daily country average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T). Announcement days are those trading days when interest rate decisions are scheduled for release at the Bank of Japan. Non-announcement days are those days when interest rate decisions, employment numbers and inflation numbers are not scheduled for release in the United States, United Kingdom, Germany, and Japan. $t$-statistics are in square brackets. Equity series are either Datastream Global Country (total return) Indices or country major stock market indices obtained from Datastream. Countries are grouped into four geographical areas. Positive and statistically significant return differences are in bold. The sample period is January 1998 to December 2013.

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | |
| A: Europe | | | | | |
| Austria | -9.71 | 248 | -0.16 | 2433 | -9.55 |
| | [-1.34] | | [-0.06] | | [-1.25] |
| Belgium | 3.80 | 248 | -0.14 | 2436 | 3.95 |
| | [0.50] | | [-0.06] | | [0.50] |
| Denmark | -1.41 | 248 | 1.72 | 2436 | -3.13 |
| | [-0.19] | | [0.70] | | [-0.41] |
| Finland | 1.72 | 242 | -4.72 | 2314 | 6.44 |
| | [0.15] | | [-1.21] | | [0.52] |
| France | -1.70 | 248 | -0.44 | 2436 | -1.26 |
| | [-0.22] | | [-0.16] | | [-0.15] |
| Germany | -1.65 | 246 | 0.19 | 2348 | -1.84 |
| | [-0.17] | | [0.06] | | [-0.18] |
| Greece | 15.02 | 241 | -8.18 | 2304 | **23.20** |
| | [1.22] | | [-2.06] | | **[1.79]** |
| Ireland | 6.36 | 248 | -4.43 | 2436 | 10.79 |
| | [0.67] | | [-1.57] | | [1.09] |
| Italy | -0.49 | 246 | -3.22 | 2345 | 2.73 |
| | [-0.07] | | [-1.23] | | [0.35] |
| The Netherlands | -2.61 | 248 | -0.53 | 2436 | -2.07 |
| | [-0.33] | | [-0.19] | | [-0.25] |
| Norway | -2.92 | 248 | 1.41 | 2436 | -4.33 |
| | [-0.33] | | [0.46] | | [-0.46] |
| Poland | 8.62 | 243 | -1.85 | 2311 | 10.46 |
| | [0.95] | | [-0.60] | | [1.09] |
| Portugal | 8.36 | 242 | -3.47 | 2344 | **11.83** |
| | [1.23] | | [-1.44] | | **[1.64]** |
| Spain | -2.62 | 244 | -2.75 | 2333 | 0.13 |
| | [-0.31] | | [-0.88] | | [0.01] |
| Sweden | -5.65 | 248 | 0.92 | 2436 | -6.56 |
| | [-0.60] | | [0.29] | | [-0.66] |
| Switzerland | -5.01 | 248 | 0.08 | 2433 | -5.09 |
| | [-0.73] | | [0.04] | | [-0.70] |
| United Kingdom | -8.55 | 246 | -1.42 | 2326 | -7.13 |
| | [-1.25] | | [-0.56] | | [-0.98] |
| B: North America | | | | | |
| Canada | 5.81 | 246 | -3.95 | 2313 | 9.76 |
| | [0.79] | | [-1.59] | | [1.26] |
| United States | 1.21 | 244 | -2.13 | 2312 | 3.34 |
| | [0.15] | | [-0.78] | | [0.38] |

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | |
| C: Asia Pacific | | | | | |
| Australia* | 7.64 | 248 | 1.93 | 2436 | 5.70 |
| | [1.22] | | [0.94] | | [0.86] |
| Hong Kong* | 0.61 | 240 | 0.82 | 2288 | -0.21 |
| | [0.06] | | [0.22] | | [-0.02] |
| Japan* | 6.02 | 248 | -0.95 | 2245 | 6.97 |
| | [0.71] | | [-0.32] | | [0.77] |
| New Zealand* | 11.90 | 247 | -0.07 | 2433 | **11.97** |
| | [2.53] | | [-0.05] | | **[2.43]** |
| Singapore* | 8.45 | 248 | -2.91 | 2434 | 11.36 |
| | [1.23] | | [-1.15] | | [1.55] |
| D: Emerging | | | | | |
| Argentina | 0.90 | 246 | 4.03 | 2416 | -3.13 |
| | [0.10] | | [1.19] | | [-0.33] |
| Brazil | -5.04 | 236 | 0.39 | 2294 | -5.43 |
| | [-0.39] | | [0.09] | | [-0.40] |
| Chile | 1.09 | 239 | -0.81 | 2302 | 1.90 |
| | [0.20] | | [-0.47] | | [0.34] |
| China* | 4.87 | 233 | 2.93 | 2262 | 1.94 |
| | [0.48] | | [0.83] | | [0.18] |
| Indonesia* | 29.22 | 235 | 1.39 | 2260 | **27.83** |
| | [2.56] | | [0.39] | | **[2.32]** |
| Malaysia* | 18.99 | 239 | 0.96 | 2287 | **18.03** |
| | [2.17] | | [0.33] | | **[1.96]** |
| Mexico | -11.78 | 244 | 1.90 | 2341 | -13.68 |
| | [-1.20] | | [0.61] | | [-1.33] |
| Philippines* | 3.57 | 239 | -0.94 | 2270 | 4.50 |
| | [0.39] | | [-0.29] | | [0.47] |
| South Africa | 4.13 | 248 | 3.80 | 2436 | 0.33 |
| | [0.58] | | [1.52] | | [0.04] |
| South Korea* | 13.12 | 242 | -0.92 | 2292 | 14.03 |
| | [1.13] | | [-0.23] | | [1.15] |
| Taiwan | -8.21 | 232 | -3.59 | 2304 | -4.62 |
| | [-0.82] | | [-1.13] | | [-0.44] |
| Thailand* | 11.57 | 243 | 0.11 | 2246 | 11.46 |
| | [0.95] | | [0.03] | | [0.90] |
| Turkey | -6.39 | 248 | 1.36 | 2436 | -7.74 |
| | [-0.43] | | [0.27] | | [-0.50] |
| Venezuela | 21.99 | 248 | 10.33 | 2436 | 11.66 |
| | [2.24] | | [3.14] | | [1.13] |

Table 5

ECB Announcements - Post January 1999

This table reports daily country average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T). Announcement days are those trading days when ECB interest rate decisions are scheduled for release. Non-announcement days are those days when interest rate decisions, employment numbers and inflation numbers are not scheduled for release in the United States, United Kingdom, Germany, and Japan. In countries denoted by star all announcements except for the Japanese ones are shifted to the next day to account for time-differences in trading. *t*-statistics are in square brackets. Equity series are either Datastream Global Country (total return) Indices or country major stock market indices obtained from Datastream. Countries are grouped into four geographical areas. Positive and statistically significant return differences are in bold.

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | |
| **A: Europe** | | | | | |
| Austria | -0.88 | 215 | -0.18 | 2293 | -0.71 |
| | [-0.12] | | [-0.07] | | [-0.09] |
| Belgium | 3.81 | 215 | -1.16 | 2296 | 4.97 |
| | [0.45] | | [-0.46] | | [0.57] |
| Denmark | 5.75 | 215 | 1.84 | 2296 | 3.91 |
| | [0.69] | | [0.73] | | [0.45] |
| Finland | 25.78 | 212 | -7.04 | 2182 | **32.82** |
| | [1.38] | | [-1.77] | | **[1.72]** |
| France | -4.83 | 215 | -0.66 | 2296 | -4.17 |
| | [-0.49] | | [-0.23] | | [-0.41] |
| Germany | -8.13 | 215 | -0.73 | 2216 | -7.40 |
| | [-0.71] | | [-0.21] | | [-0.62] |
| Greece | 7.16 | 213 | -9.86 | 2170 | 17.02 |
| | [0.61] | | [-2.45] | | [1.37] |
| Ireland | 5.86 | 215 | -5.53 | 2296 | 11.39 |
| | [0.65] | | [-1.91] | | [1.20] |
| Italy | 6.59 | 215 | -4.36 | 2210 | 10.96 |
| | [0.83] | | [-1.67] | | [1.31] |
| The Netherlands | -6.95 | 215 | -1.03 | 2296 | -5.91 |
| | [-0.74] | | [-0.36] | | [-0.60] |
| Norway | -3.92 | 215 | 1.81 | 2296 | -5.73 |
| | [-0.35] | | [0.58] | | [-0.49] |
| Poland | 17.86 | 214 | -2.46 | 2179 | **20.31** |
| | [1.71] | | [-0.82] | | **[1.87]** |
| Portugal | -0.14 | 212 | -3.90 | 2213 | 3.77 |
| | [-0.02] | | [-1.61] | | [0.47] |
| Spain | -4.90 | 214 | -3.56 | 2203 | -1.34 |
| | [-0.46] | | [-1.13] | | [-0.12] |
| Sweden | -1.74 | 215 | 0.40 | 2296 | -2.14 |
| | [-0.15] | | [0.12] | | [-0.18] |
| Switzerland | 2.48 | 215 | -0.06 | 2293 | 2.54 |
| | [0.33] | | [-0.03] | | [0.33] |
| United Kingdom | -1.18 | 215 | -1.74 | 2194 | 0.56 |
| | [-0.14] | | [-0.66] | | [0.06] |
| **B: North America** | | | | | |
| Canada | 2.57 | 213 | -3.94 | 2180 | 6.51 |
| | [0.32] | | [-1.54] | | [0.77] |
| United States | -1.61 | 212 | -2.46 | 2180 | 0.85 |
| | [-0.17] | | [-0.86] | | [0.09] |

| Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | |
| **C: Asia Pacific** | | | | | |
| Australia* | 2.82 | 215 | 1.39 | 2301 | 1.43 |
| | [0.46] | | [0.67] | | [0.22] |
| Hong Kong* | 17.11 | 207 | 0.18 | 2168 | 16.93 |
| | [1.49] | | [0.05] | | [1.41] |
| Japan* | 2.43 | 206 | -1.99 | 2145 | 4.42 |
| | [0.28] | | [-0.64] | | [0.47] |
| New Zealand* | 5.94 | 213 | 0.90 | 2300 | 5.04 |
| | [1.12] | | [0.64] | | [0.92] |
| Singapore* | 9.28 | 214 | 2.76 | 2300 | 6.52 |
| | [1.30] | | [1.14] | | [0.87] |
| **D: Emerging** | | | | | |
| Argentina | 11.75 | 215 | 4.07 | 2276 | 7.68 |
| | [1.12] | | [1.19] | | [0.70] |
| Brazil | 6.26 | 213 | 0.56 | 2164 | 5.70 |
| | [0.45] | | [0.13] | | [0.39] |
| Chile | 6.60 | 214 | -0.05 | 2169 | 6.65 |
| | [1.17] | | [-0.03] | | [1.13] |
| China* | 15.87 | 193 | 3.60 | 2125 | 12.27 |
| | [1.50] | | [1.01] | | [1.10] |
| Indonesia* | 15.00 | 202 | 5.15 | 2147 | 9.85 |
| | [1.52] | | [1.57] | | [0.95] |
| Malaysia* | 3.54 | 208 | 1.41 | 2163 | 2.14 |
| | [0.57] | | [0.70] | | [0.33] |
| Mexico | 8.86 | 210 | 2.71 | 2207 | 6.16 |
| | [0.83] | | [0.86] | | [0.55] |
| Philippines* | 14.89 | 207 | 0.80 | 2153 | **14.09** |
| | [1.91] | | [0.26] | | **[1.68]** |
| South Africa | 4.71 | 215 | 4.27 | 2296 | 0.44 |
| | [0.53] | | [1.75] | | [0.05] |
| South Korea* | 1.67 | 202 | 0.17 | 2174 | 1.50 |
| | [0.13] | | [0.05] | | [0.11] |
| Taiwan* | 22.04 | 200 | -0.14 | 2181 | **22.17** |
| | [2.06] | | [-0.04] | | **[1.98]** |
| Thailand* | 6.01 | 206 | 3.55 | 2153 | 2.46 |
| | [0.61] | | [1.08] | | [0.24] |
| Turkey | 18.86 | 215 | -0.24 | 2296 | 19.10 |
| | [0.97] | | [-0.05] | | [0.95] |
| Venezuela | 9.76 | 215 | 11.31 | 2296 | -1.55 |
| | [1.10] | | [3.64] | | [-0.17] |

except for Italy. There are also no significant return differentials outside of Europe, apart from Malaysia and Thailand. However, during the pre-Maastricht 1958-1992 period, returns are economically and statistically higher in the core eurozone countries, including Germany (9.7 bps, with a $t$-statistic of 2.51), France (9.9 bps, with a $t$-statistic of 1.76), and Italy (15.4 bps, with a $t$-statistic of 2.45). Globally, there are six countries where average market returns are not higher on Bundesbank a-days, and the only such country in Europe is Greece, with a negligible difference of -1.5 bps ($t$-statistic=-0.08).

Comparison of these estimates with those of the previous tables is complicated by the fact that Bundesbank announcements occurred twice as frequently as other announcements, and thus were presumably roughly half as important. As a rough rule of thumb, doubling the point estimates in Panel II of Table 6 makes the Bundesbank announcement premium in most European markets comparable in magnitude to the FOMC announcement premium over the same period. This makes the non-existence of the ECB announcement premium since 1999 in many ways all the more striking.

Bundesbank decisions were typically announced in the afternoon or the next morning, and our data does not provide the exact time. Thus, it is possible that markets could sometimes react to its decisions only on the next trading day. As a robustness test, Table A.4 in the Appendix shows our results under the assumption that the market reaction occurred on the first trading day following the scheduled meeting. All our main results remain the same.

## 3.2 Other non-U.S. Macroeconomic Announcements

The irrelevance of non-U.S. central bank announcements for their domestic market risk premia raises the question of whether other macroeconomic announcements matter in domestic markets.

Table 6

Bundesbank Announcements - Various Sample Periods

This table reports daily country average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T). Announcement days are those trading days when interest rate decisions are scheduled for release at the Bundesbank. Non-announcement days are those days when interest rate decisions, employment numbers and inflation numbers are not scheduled for release in the United States, United Kingdom, Germany, and Japan. In countries denoted by star all announcements except for the Japanese ones are shifted to the next day to account for time-differences in trading. $t$-statistics are in square brackets. Equity series are either Datastream Global Country (total return) Indices or country major stock market indices obtained from Datastream. Countries are grouped into four geographical areas. Positive and statistically significant return differences are in bold.

I: January 1958 to December 1998

| Country | a-days | | nona-days | | Diff. | Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | | | $\mu$ | T | $\mu$ | T | |
| A: Europe | | | | | | C: Asia Pacific | | | | | |
| Austria | 1.33 | 608 | 0.33 | 4261 | 1.00 | Australia* | 2.85 | 624 | 1.73 | 4497 | 1.12 |
| | [0.42] | | [0.26] | | [0.29] | | [0.65] | | [1.02] | | [0.24] |
| Belgium | 5.89 | 629 | 0.76 | 4433 | 5.13 | Hong Kong* | 16.92 | 681 | 4.50 | 4870 | 12.42 |
| | [1.89] | | [0.63] | | [1.54] | | [2.44] | | [1.60] | | [1.66] |
| Denmark | 3.85 | 620 | 2.77 | 4335 | 1.08 | Japan* | 1.15 | 958 | 1.86 | 7143 | -0.71 |
| | [0.91] | | [1.92] | | [0.24] | | [0.42] | | [1.73] | | [-0.24] |
| Finland | 8.07 | 287 | 0.66 | 1826 | 7.41 | New Zealand* | 1.37 | 266 | 3.15 | 1788 | -1.77 |
| | [1.13] | | [0.22] | | [0.96] | | [0.21] | | [1.12] | | [-0.25] |
| France | 8.00 | 629 | 1.97 | 4444 | 6.03 | Singapore* | -1.73 | 627 | 1.44 | 4466 | -3.17 |
| | [1.73] | | [1.20] | | [1.23] | | [-0.34] | | [0.67] | | [-0.58] |
| Germany | 1.25 | 823 | -1.06 | 5744 | 2.31 | D: Emerging | | | | | |
| | [0.37] | | [-0.76] | | [0.63] | Argentina | -36.36 | 131 | 7.92 | 846 | -44.28 |
| Greece | 1.88 | 239 | 9.08 | 1510 | -7.20 | | [-2.05] | | [1.22] | | [-2.34] |
| | [0.17] | | [1.84] | | [-0.59] | Brazil | 40.96 | 179 | 71.98 | 1082 | -31.02 |
| Ireland | 8.29 | 626 | 4.32 | 4441 | 3.96 | | [1.31] | | [5.82] | | [-0.92] |
| | [2.03] | | [2.72] | | [0.90] | Chile | 8.27 | 278 | 2.93 | 1806 | 5.35 |
| Italy | 9.23 | 698 | -3.30 | 4905 | **12.53** | | [1.14] | | [1.22] | | [0.70] |
| | [1.79] | | [-1.68] | | **[2.27]** | China* | 44.78 | 164 | 1.88 | 1071 | 42.90 |
| The Netherlands | -0.06 | 629 | 2.40 | 4443 | -2.46 | | [1.95] | | [0.12] | | [1.55] |
| | [-0.02] | | [1.69] | | [-0.62] | Indonesia* | 22.77 | 337 | 0.29 | 2266 | 22.48 |
| Norway | 4.90 | 459 | 2.85 | 3105 | 2.05 | | [1.41] | | [0.09] | | [1.37] |
| | [0.70] | | [1.08] | | [0.28] | Malaysia* | 20.09 | 442 | -0.32 | 2997 | **20.41** |
| Poland | -12.10 | 165 | 31.96 | 874 | -44.06 | | [2.55] | | [-0.11] | | **[2.42]** |
| | [-0.57] | | [3.06] | | [-1.87] | Mexico | 23.21 | 257 | 4.50 | 1655 | 18.71 |
| Portugal | 9.87 | 255 | -1.08 | 1585 | 10.95 | | [2.17] | | [1.01] | | [1.61] |
| | [1.46] | | [-0.46] | | [1.53] | Philippines* | 10.59 | 298 | 4.82 | 1988 | 5.77 |
| Spain | 1.50 | 572 | -1.13 | 3461 | 2.64 | | [1.00] | | [1.07] | | [0.50] |
| | [0.33] | | [-0.59] | | [0.53] | South Africa | 3.74 | 628 | 3.84 | 4448 | -0.10 |
| Sweden | 8.80 | 411 | 3.37 | 2742 | 5.43 | | [0.76] | | [1.81] | | [-0.02] |
| | [1.47] | | [1.38] | | [0.84] | South Korea* | -4.50 | 555 | 0.54 | 3840 | -5.04 |
| Switzerland | 2.87 | 626 | -1.01 | 4418 | 3.88 | | [-0.74] | | [0.24] | | [-0.78] |
| | [0.88] | | [-0.80] | | [1.11] | Taiwan* | 5.37 | 637 | 3.23 | 4559 | 2.14 |
| United Kingdom | 0.54 | 712 | 0.33 | 5000 | 0.21 | | [0.83] | | [1.20] | | [0.31] |
| | [0.15] | | [0.22] | | [0.05] | Thailand* | 13.11 | 536 | -1.72 | 3769 | **14.83** |
| B: North America | | | | | | | [2.62] | | [-0.74] | | **[2.69]** |
| Canada | -2.42 | 715 | -1.11 | 4967 | -1.31 | Turkey | 37.71 | 265 | 24.79 | 1744 | 12.92 |
| | [-0.77] | | [-0.97] | | [-0.39] | | [2.32] | | [3.59] | | [0.73] |
| United States | 0.82 | 982 | 1.90 | 7207 | -1.08 | Venezuela | -16.25 | 219 | 9.90 | 1423 | -26.15 |
| | [0.34] | | [2.00] | | [-0.42] | | [-1.01] | | [1.36] | | [-1.48] |

II: January 1958 to December 1992

| Country | a-days | | nona-days | | Diff. | Country | a-days | | nona-days | | Diff. |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\mu$ | T | $\mu$ | T | | | $\mu$ | T | $\mu$ | T | |
| A: Europe | | | | | | C: Asia Pacific | | | | | |
| Austria | 5.04 | 463 | -0.08 | 3319 | 5.12 | Australia | 6.89 | 478 | 0.34 | 3540 | 6.55 |
| | [1.44] | | [-0.05] | | [1.35] | | [1.30] | | [0.17] | | [1.16] |
| Belgium | 7.21 | 483 | -0.76 | 3491 | **7.97** | Hong Kong | 24.65 | 540 | 4.92 | 3960 | **19.73** |
| | [2.01] | | [-0.55] | | **[2.07]** | | [3.06] | | [1.58] | | **[2.28]** |
| Denmark | 5.40 | 474 | 1.72 | 3394 | 3.68 | Japan | 1.76 | 818 | 2.24 | 6250 | -0.49 |
| | [1.07] | | [1.04] | | [0.69] | | [0.62] | | [2.04] | | [-0.16] |
| Finland | 5.18 | 142 | -6.53 | 930 | 11.71 | New Zealand | 16.67 | 120 | -0.93 | 831 | 17.59 |
| | [0.73] | | [-1.86] | | [1.48] | | [1.44] | | [-0.20] | | [1.41] |
| France | 11.62 | 483 | 1.68 | 3503 | **9.94** | Singapore | 0.16 | 481 | 2.40 | 3509 | -2.24 |
| | [2.19] | | [0.89] | | **[1.76]** | | [0.03] | | [0.95] | | [-0.34] |
| Germany | 6.42 | 677 | -3.28 | 4848 | **9.71** | D: Emerging | | | | | |
| | [1.80] | | [-2.24] | | **[2.51]** | Argentina | - | - | - | - | - |
| Greece | 6.58 | 99 | 8.13 | 638 | -1.54 | | | | | | |
| | [0.36] | | [0.93] | | [-0.08] | Brazil | - | - | - | - | - |
| Ireland | 8.02 | 480 | 3.25 | 3500 | 4.77 | | - | | - | | - |
| | [1.64] | | [1.77] | | [0.91] | Chile | 15.55 | 137 | 5.63 | 918 | 9.92 |
| Italy | 10.66 | 554 | -4.78 | 4003 | **15.44** | | [1.23] | | [1.51] | | [0.75] |
| | [1.81] | | [-2.19] | | **[2.45]** | China | - | - | - | - | - |
| The Netherlands | 1.91 | 483 | 0.70 | 3501 | 1.21 | | - | | - | | - |
| | [0.45] | | [0.44] | | [0.27] | Indonesia | 27.26 | 199 | 0.58 | 1360 | 26.67 |
| Norway | 11.58 | 313 | 1.41 | 2164 | 10.18 | | [1.07] | | [0.16] | | [1.04] |
| | [1.24] | | [0.41] | | [1.02] | Malaysia | 16.68 | 303 | 2.84 | 2091 | **13.84** |
| Poland | - | - | - | - | - | | [2.41] | | [0.94] | | **[1.83]** |
| | [0.81] | | [-0.64] | | [1.03] | Mexico | 58.37 | 116 | 6.92 | 757 | **51.46** |
| Portugal | 12.33 | 113 | -10.19 | 704 | **22.52** | | [3.52] | | [1.07] | | **[2.89]** |
| | [1.10] | | [-2.81] | | **[1.90]** | Philippines | 19.45 | 155 | 8.38 | 1069 | 11.07 |
| Spain | 0.71 | 432 | -3.72 | 2570 | 4.44 | | [1.28] | | [1.20] | | [0.66] |
| | [0.14] | | [-1.68] | | [0.80] | South Africa | 6.07 | 482 | 3.72 | 3506 | 2.35 |
| Sweden | 12.38 | 265 | 1.16 | 1800 | 11.22 | | [1.05] | | [1.51] | | [0.37] |
| | [1.62] | | [0.37] | | [1.36] | South Korea | -0.39 | 416 | 2.67 | 2944 | -3.06 |
| Switzerland | 3.64 | 480 | -2.52 | 3476 | 6.17 | | [-0.06] | | [1.19] | | [-0.44] |
| | [1.05] | | [-1.86] | | [1.66] | Taiwan | 11.37 | 502 | 3.72 | 3661 | 7.65 |
| United Kingdom | 0.97 | 566 | 0.16 | 4095 | 0.81 | | [1.56] | | [1.20] | | [0.96] |
| | [0.22] | | [0.09] | | [0.17] | Thailand | 16.13 | 399 | 0.91 | 2870 | **15.22** |
| B: North America | | | | | | | [3.20] | | [0.39] | | **[2.74]** |
| Canada | -0.17 | 570 | -1.79 | 4071 | 1.62 | Turkey | -16.41 | 119 | 11.66 | 802 | -28.07 |
| | [-0.05] | | [-1.42] | | [0.42] | | [-0.82] | | [1.38] | | [-1.30] |
| United States | 2.55 | 837 | 1.36 | 6307 | 1.19 | Venezuela | 10.09 | 73 | 15.50 | 481 | -5.42 |
| | [0.99] | | [1.34] | | [0.43] | | [0.33] | | [1.04] | | [-0.16] |

To address this question, we look at announcement premia for Japan, Germany, and the U.K. (as well as the U.S.) on not just domestic central bank a-days, but also domestic inflation and employment a-days.[57] Non-announcement days are again

[57]Please see the online Appendix for details of the sources and methods for collecting the historical dates of these scheduled announcements.

defined as days on which there are neither U.S. announcements (FOMC, inflation, or employment) nor domestic ones (central bank, inflation, or employment). The four panels of Table 7 report a-day average excess returns (in domestic currency) and their difference relative to n-day average excess returns for each of the four major stock markets, broken out by announcement type and then all together. Each panel is also divided into domestic a-days in the first row and U.S. a-days in the second row.

Starting with Japan in the first panel, we see there is no statistically significant increase in the risk premium for either domestic monetary policy or employment announcements, although the difference between a- and n-days is positive in both cases. However, average market returns are significantly higher for Japanese inflation announcements (12.2 bps, with a $t$-statistic of 2.64), and the return differential is also positive (5.7 bps) and significant ($t$-statistic=1.68) when we aggregate all three announcements. The first column of the second row shows the premium for FOMC announcements in Japan. As we already document in previous sections, there is a large premium of 25.8 bps with a $t$-statistic of 3.29 over the 1958-2013 period. In contrast, the next two columns show that U.S. employment and inflation announcements are not associated with higher average returns on Japanese stocks in the full 1958-2013 period.

Turning to Germany and the Bundesbank-ECB announcements in the second panel, again we find no evidence of a domestic central bank a-day effect and strong evidence of an FOMC a-day effect. For other macroeconomic announcements, the pattern looks the opposite, with a strong domestic a-day and a weak U.S. a-day effect. The German stock market enjoys excess returns that are on average 14.7 bps higher ($t$-statistic=2.19) on German employment a-days compared to n-days. While the announcement premium is not quite significant ($t$-statistic=1.49) for German inflation, its economic magnitude is quite meaningful at 9.2 bps. The premia in the German market associated with U.S. inflation and employment announcements are positive, but much lower in magnitude and not statistically significant.

Finally, in the U.K. (third panel), as usual we find no evidence for a BoE announcement premium (it's actually negative) and document a strong FOMC

Table 7

Domestic (D) versus U.S. Announcements: Major Findings

This table reports daily country average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T) for four major countries (Japan, Germany, United Kingdom, and the Unites States). Announcement days are those trading days when either interest rate decisions (Monetary Policy), employment numbers, or inflation numbers are scheduled for release in the country of interest or in the United States. Non-announcement days are those trading days when domestic (D) and U.S. announcements are not scheduled for release. Test assets are country major stock market (price) indices denominated in domestic currency (Column 1). The sample period and the country where the announcements are made (a-day Country) are defined, respectively, in Columns 2 and 3. Column 4 accommodates time-differences in trading (a-day Timing) stating whether a-days are shifted to the next-day (Next) or not (Same). $t$-statistics are in square brackets. For each country, daily market excess returns are computed as the difference between the return on the test asset and the U.S. risk-free rate. Equity series are from Datastream; the daily risk-free rate is obtained from Kenneth French's website. Positive and statistically significant return differences are in bold.

| Test Asset | Sample | a-day Country | a-day Timing | a-days | | | | | | | | | | | | nona-days | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | Monetary Policy | | | Employment | | | Inflation | | | ALL | | | | |
| | | | | $\mu$ | T | Diff. | $\mu$ | T | Diff. | $\mu$ | T | Diff. | $\mu$ | T | Diff. | $\mu$ | T |
| TOPIX (Japan) | Jan 1958 Dec 2013 | D | US: Next D: Same | 6.99 [0.88] | 291 | 6.50 [0.82] | 4.57 [1.11] | 653 | 4.08 [0.96] | 12.69 [2.82] | 528 | **12.20 [2.64]** | 6.19 [1.91] | 1181 | **5.70 [1.68]** | 0.49 [0.49] | 11225 |
| | | US | | 26.32 [3.38] | 300 | **25.83 [3.29]** | -2.44 [-0.52] | 619 | -2.93 [-0.61] | -1.04 [-0.23] | 633 | -1.53 [-0.33] | 2.89 [0.93] | 1500 | 2.40 [0.74] | 0.49 [0.49] | 11225 |
| DAX 30 (Germany) | Jan 1965 Dec 2013 | D | Same | -0.69 [-0.19] | 1038 | -0.33 [-0.09] | 14.31 [2.18] | 371 | **14.67 [2.19]** | 8.85 [1.46] | 442 | 9.21 [1.49] | 5.65 [1.94] | 1748 | **6.01 [1.89]** | -0.36 [-0.28] | 9287 |
| | | US | | 15.64 [2.26] | 307 | **16.00 [2.27]** | 7.13 [1.40] | 577 | 7.49 [1.43] | 3.99 [0.81] | 573 | 4.34 [0.86] | 7.79 [2.42] | 1406 | **8.15 [2.35]** | -0.36 [-0.28] | 9287 |
| FTSE ALL SHARE (UK) | July 1981 Dec 2013 | D | Same | -6.60 [-0.93] | 249 | -7.34 [-1.02] | 4.31 [0.80] | 357 | 3.57 [0.65] | 9.10 [1.82] | 388 | 8.36 [1.61] | 3.24 [0.97] | 974 | 2.50 [0.70] | 0.74 [0.57] | 6357 |
| | | US | | 16.43 [2.64] | 261 | **15.66 [2.46]** | 10.66 [2.10] | 377 | **9.89 [1.88]** | -2.64 [-0.49] | 384 | -3.41 [-0.62] | 6.89 [2.16] | 1011 | **6.12 [1.78]** | 0.77 [0.59] | 6386 |
| CRSP (US) | Jan 1958 Dec 2013 | US | Same | 24.76 [4.11] | 309 | **23.31 [3.83]** | 8.70 [2.31] | 652 | **7.26 [1.88]** | 8.12 [2.09] | 664 | **6.68 [1.68]** | 11.39 [4.55] | 1574 | **9.94 [3.76]** | 1.45 [1.69] | 12523 |

announcement premium. Domestic inflation and employment announcement premia are positive at 8.4 and 3.6 bps, respectively, though not quite statistically significant.[58] The last panel replicates the results in Savor and Wilson (2013) showing the effects of FOMC, employment, and inflation announcements since 1958 (1974 for the FOMC). These are all positive and significant, although as we document above, the employment and inflation effect have faded over time as the FOMC has become more prominent.

Taken together, these results cast doubt on one potential explanation for our earlier results, namely that our three non-U.S. central banks do not matter to diversified equity investors because the economies they represent are not systemically important. For example, it would not be surprising if Croatian stocks do not exhibit elevated risk premia in response to Croatian inflation (or central bank) announcements, as the idiosyncratic component of Croatian inflation is very likely completely diversifiable to international investors holding Croatian stocks. This is much less likely for holders of Japanese or German stocks, and consistent with this conjecture we show that those markets have higher average returns on domestic a-days. The "small domestic economy" explanation is more likely to apply for the U.K., since its stock market contains many large, globally active companies with low exposure to the U.K. economy. However, even in the U.K., we document positive, although not significant, domestic a-day premia.

## 3.3 Impact on the U.S. Stock Market

Results in the previous section suggest that domestic macroeconomic announcements, unlike domestic central bank announcements, are sometimes associated with positive risk premia, especially in the case of inflation. A natural follow-on question is to explore whether these announcements also command higher returns in foreign markets. In Table 8, we examine the relation between U.S. stock market returns and announcements in U.K., Japan, and Germany.

[58] There is also a positive U.S. employment a-day premium (9.9 bps, with a $t$-statistic of 1.88).

Panel I of Table 8 shows average U.S. returns on a-days in each of the three other countries and the difference between a-days and n-days. We see a similar pattern to the one we already observed in domestic markets: in contrast to central bank announcements, U.S. market returns are higher on inflation and employment a-days for all the three countries. While the difference is only statistically significant in the case of U.K. inflation, the magnitudes involved are economically meaningful. Furthermore, when we combine the announcements across countries in Panels III and IV, inflation announcements are associated with robustly positive return differentials, ranging from 6.7 to 9.3 bps (depending on the exact grouping) and with $t$-statistics that are typically well above 2. This result holds even when we explicitly exclude all U.S. a-days from the a-day classification (U.S. announcements sometimes coincide with foreign ones).

The findings here provide support for the hypothesis that the economic news in the three countries we study is systemically important to global investors, and that consequently investors demand a risk premium for exposure to this news, even in other markets. This is an intuitive and not overly surprising result, but it makes our previous results on non-U.S. central banks even more puzzling. If all four economic regions are systematically important, why do global investors demand a very high risk premium to bear risks associated with FOMC announcements and not demand any risk premium for exposure to risks associated with BoE, BoJ, and ECB decisions?

## 3.4 Inert Central Banks?

One potential explanation for why non-U.S. central banks do not appear to matter to investors is that these central banks do not pursue active monetary policies (or at least have not done so during our sample period). In this case, investors would face little uncertainty regarding banks' decisions, and therefore exposure to such decisions would not command a risk premium. However, contrary to this hypothesis, in the period we study the BoE, ECB, and BoJ all exhibit activist monetary policies, with large variations in interest rates as well as use of unconventional monetary policy tools. For example, BoE rates ranged from 0.5 to 7.5%, with two tightening

Table 8

How much do U.S. investors care about foreign announcements?

This table reports daily U.S. average excess returns ($\mu$, in basis points) on announcement (a-days) and non-announcement (nona-days) days, their difference (*Diff.*) and the number of observations in each sample (T). Announcement days are those trading days when interest rate decisions, employment numbers and/or inflation numbers are scheduled for release in Germany (Ger), Japan (Jap) and/or the United Kingdom (UK). Column 3 and 4 report the set of countries entering the announcement and nonannouncement sample, respectively. $t$-statistics are in square brackets. Positive and statistically significant return differences are in bold.

| Test | Sample | a-day | nona-day | a-days | | | | | | | | | | | | nona-days | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | Monetary Policy | | | Employment | | | Inflation | | | ALL | | | | |
| Asset | | countries | countries | $\mu$ | T | Diff. | $\mu$ | T | Diff. | $\mu$ | T | Diff. | $\mu$ | T | Diff. | $\mu$ | T |
| Panel I: Country-i a-days | | | | | | | | | | | | | | | | | |
| CRSP | July 1981 | UK | UK+US | -0.95 | 196 | -2.63 | 2.36 | 359 | 0.68 | 12.08 | 386 | **10.40** | 5.32 | 927 | 3.64 | 1.68 | 6433 |
| | Dec 2013 | | | [-0.10] | | [-0.27] | [0.39] | | [0.11] | [2.17] | | **[1.81]** | [1.38] | | [0.89] | [1.24] | |
| CRSP | Jan 1958 | Jap | Jap+US | 2.34 | 284 | 1.00 | 6.07 | 634 | 4.73 | 6.36 | 512 | 5.02 | 4.94 | 1149 | 3.60 | 1.34 | 11495 |
| | Dec 2013 | | | [0.32] | | [0.14] | [1.71] | | [1.29] | [1.71] | | [1.31] | [1.73] | | [1.21] | [1.50] | |
| CRSP | Jan 1965 | Ger | Ger+US | -0.80 | 1030 | -1.58 | 6.81 | 361 | 6.03 | 8.53 | 437 | 7.75 | 3.44 | 1726 | 2.67 | 0.78 | 9343 |
| | Dec 2013 | | | [-0.28] | | [-0.51] | [1.24] | | [1.08] | [1.83] | | [1.62] | [1.46] | | [1.04] | [0.75] | |
| Panel II: Country-i a-days that are not U.S. a-days | | | | | | | | | | | | | | | | | |
| CRSP | July 1981 | UK | UK+US | -2.50 | 189 | -4.18 | 2.73 | 315 | 1.05 | 7.44 | 258 | 5.76 | 2.56 | 748 | 0.88 | 1.68 | 6433 |
| | Dec 2013 | | | [-0.26] | | [-0.42] | [0.46] | | [0.17] | [1.16] | | [0.88] | [0.62] | | [0.20] | [1.24] | |
| CRSP | Jan 1958 | Jap | Jap+US | -4.94 | 233 | -6.27 | 5.16 | 562 | 3.83 | 6.12 | 509 | 4.78 | 2.69 | 1028 | 1.35 | 1.34 | 11495 |
| | Dec 2013 | | | [-0.60] | | [-0.76] | [1.37] | | [0.99] | [1.64] | | [1.25] | [0.90] | | [0.43] | [1.50] | |
| CRSP | Jan 1965 | Ger | Ger+US | -1.63 | 949 | -2.40 | 6.91 | 307 | 6.14 | 7.72 | 423 | 6.95 | 2.91 | 1582 | 2.13 | 0.78 | 9343 |
| | Dec 2013 | | | [-0.53] | | [-0.75] | [1.20] | | [1.05] | [1.64] | | [1.44] | [1.19] | | [0.81] | [0.75] | |

[continued]

| Test Asset | Sample | a-day countries | nona-day countries | a-days: Monetary Policy $\mu$ | Monetary Policy T | Monetary Policy Diff. | Employment $\mu$ | Employment T | Employment Diff. | Inflation $\mu$ | Inflation T | Inflation Diff. | ALL $\mu$ | ALL T | ALL Diff. | nona-days $\mu$ | nona-days T |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Panel III: All foreign a-days that are not U.S. a-days | | | | | | | | | | | | | | | | | |
| CRSP | Jan 1958 | UK, Ger | ALL | 1.23 | 1443 | 0.35 | 4.65 | 1278 | 3.76 | 8.81 | 1248 | **7.92** | 4.42 | 3425 | **3.53** | 0.89 | 9387 |
| | Dec 2013 | Jap | | [0.48] | | [0.12] | [1.68] | | [1.28] | [3.22] | | **[2.73]** | [2.62] | | **[1.81]** | [0.90] | |
| CRSP | Jan 1970 | UK, Jap | ALL | -2.17 | 418 | -2.68 | 4.16 | 814 | 3.65 | 8.86 | 894 | **8.36** | 4.70 | 1802 | 4.20 | 0.50 | 7011 |
| | Dec 2013 | | | [-0.35] | | [-0.42] | [1.15] | | [0.95] | [2.77] | | **[2.43]** | [1.86] | | [1.49] | [0.40] | |
| CRSP | Jan 1970 | UK, Ger | ALL | 0.01 | 921 | -0.49 | 4.01 | 663 | 3.50 | 9.79 | 807 | **9.28** | 4.60 | 2206 | 4.10 | 0.50 | 7011 |
| | Dec 2013 | | | [0.00] | | [-0.14] | [0.96] | | [0.80] | [2.69] | | **[2.42]** | [2.05] | | [1.60] | [0.40] | |
| CRSP | Jan 1970 | Jap, Ger | ALL | -0.48 | 1113 | -0.98 | 5.79 | 803 | 5.29 | 7.18 | 865 | **6.68** | 3.27 | 2372 | 2.77 | 0.50 | 7011 |
| | Dec 2013 | | | [-0.15] | | [-0.29] | [1.71] | | [1.47] | [2.34] | | **[2.02]** | [1.62] | | [1.17] | [0.40] | |
| Panel IV: All foreign a-days that are not U.S. a-days or a-days of other countries | | | | | | | | | | | | | | | | | |
| CRSP | Jan 1970 | UK, Jap | ALL | -1.18 | 407 | -1.69 | 4.17 | 812 | 3.67 | 9.03 | 893 | **8.53** | 5.49 | 1754 | **4.98** | 0.50 | 7011 |
| | Dec 2013 | | | [-0.19] | | [-0.26] | [1.15] | | [0.96] | [2.82] | | **[2.48]** | [2.17] | | **[1.77]** | [0.40] | |
| CRSP | Jan 1970 | UK, Ger | ALL | 2.13 | 803 | 1.62 | 3.55 | 660 | 3.05 | 9.77 | 805 | **9.27** | 5.34 | 2009 | **4.84** | 0.50 | 7011 |
| | Dec 2013 | | | [0.65] | | [0.46] | [0.85] | | [0.70] | [2.68] | | **[2.41]** | [2.31] | | **[1.84]** | [0.40] | |
| CRSP | Jan 1970 | Jap, Ger | ALL | -0.20 | 1099 | -0.70 | 5.39 | 787 | 4.89 | 7.43 | 795 | **6.92** | 2.73 | 2241 | 2.23 | 0.50 | 7011 |
| | Dec 2013 | | | [-0.06] | | [-0.21] | [1.60] | | [1.36] | [2.32] | | **[2.02]** | [1.33] | | [0.93] | [0.40] | |

and three easing cycles. BoE also engaged in multiple rounds of QE, in which it purchased both government and (in smaller amounts) private securities. The interest rates set by the ECB ranged from 0.25% to 4.75%, with three cycles of tightening and easing. The ECB also used various unconventional monetary policy tools, including a number of asset purchase programs and long-term financing operations.[59] The BoJ rates moved only in a tight range of 0 to 0.5%, but BoJ did engage in a number of unconventional policies, including an early QE (started in 2001) and other asset purchase programs.

Another possibility is that the absence of an announcement-day stock market premium for the major non-U.S. central banks is due to the fact that their decisions are widely anticipated, either by design or inadvertently. Decisions by a central bank might be unsurprising if it is inactive, which we argue above is not the case, or if it is active but its actions are common knowledge to investors in advance of its announcements. We describe such banks as inert.[60]

If a central bank is inert, its announcements should not be surprising to investors, and consequently there should be no discernible effect of its announcements on the volatility of price changes for the affected assets. To explore this hypothesis, we look at the market most directly impacted by central bank decisions, the one for short-run interest rates in the relevant currencies. In the U.S., there exists a dedicated Federal Funds futures market, in which participants can speculate on (or hedge) the actual Federal Reserve interest rate announcements. Bernanke and Kuttner (2005) show that there are surprises in that market when Fed announcements occur and that stock market prices respond accordingly (although Savor and Wilson (2013) find that these surprises appear only to affect stock market returns for unscheduled announcements). Because this phenomenon is already well explored in the U.S., we study only the three other central banks, the ECB, BoJ, and BoE.

[59] ECB's QE program started a year after our sample ends.

[60] Central banks might be inert because they choose not to pursue an active monetary policy, because control of monetary policy resides elsewhere (as it did in the U.K. and Japan before 1997 and 1998, respectively) so that the banks' announcements do not reveal relevant new information, or because the banks manage investors' expectations of monetary policy so effectively that their announcements are common knowledge before they occur.

In the euro, yen, and sterling currency zones, there are no equivalent exchange-traded interest rate futures to the Federal Funds futures. However, all three currency zones have an active interbank lending market, with their own versions of the London Interbank Offered Rate (Libor), the average of interbank lending rate quotes. Futures contracts on 3-month Euro Interbank Offered Rate (Euribor), Euroyen Tokyo Interbank Offered Rate (Euroyen Tibor), and Sterling London Interbank Offered Rate (Sterling Libor) are all traded on large and liquid exchanges in either London or Chicago.[61] Because these interbank rates are typically set with reference to central bank rates, we assume that central bank announcement surprises are reflected in the futures-implied rates. Thus, we use the change in the rates implied by interest rate futures prices as our proxy for surprises associated with announcements by a particular central bank.

We collect 3-month Euribor, Euroyen Tibor, and Sterling Libor futures prices. (Although 1-month futures contracts exist, they have much lower volume and open interest.) We then create a continuous time series of implied rate changes that is based on the nearest-to-maturity contract. For example, for an announcement in April, we first use the 3-month futures contract expiring in March up to its expiration date, and then switch to the 3-month contract expiring in June. To deal with any effects of switches in the front contract, we create a dummy variable, 'Contract Switch', which equals 1 on the first date of a new contract.

Table 9 shows the results of OLS regressions of daily changes (or absolute changes) in the implied interest rates on the relevant central bank announcement dummy, as well as the Contract Switch dummy as a control, for the 1999-2013 period (the period during which all three central banks are independent).[62] The specification where the dependent variable is the absolute change is of special interest, as it tests whether (implied) interest rate expectations experience larger changes

[61] Euro and sterling interbank futures are traded in London on the London International Financial Futures and Options Exchange (LIFFE) and yen in Chicago on the Chicago Mercantile Exchange (CME).

[62] In the case of the euro and sterling, no announcement day ever coincides with a switching day, while in the case of the yen there is one BoJ announcement that occurs on a switching day.

on announcement days. If a given central bank is inert, we should find no relation between the absolute change in implied rates and its announcement indicator.

Table 9
Central Bank Announcements and Interest Rate Changes

This table presents the results of OLS regressions of daily interbank rate futures on announcement-day dummmies for the European Central Bank, the Bank of Japan, and the Bank of England. In Panel A, the dependent variable is the change or absolute change in the 3-month futures-implied Euribor rate (in basis points), and the announcement day dummy is set to 1 on days of scheduled ECB announcements and to 0 otherwise. In Panel B, the dependent variable is the change or absolute change in the 3-month futures-implied Euroyen Tibor rate (in basis points), and the announcement day dummy is set to 1 on days of scheduled BoJ announcements and to 0 otherwise. In Panel C, the dependent variable is the change or absolute change in the 3-month futures-implied Sterling Libor rate (in basis points), and the announcement day dummy is set to 1 on days of scheduled BoJ announcements and to 0 otherwise. Contract switch is a dummy variable for days of contract changes, which occur when we roll the futures contract to the next available one. *t*-statistics are in brackets.

| | Intercept | A-day | Contract Switch | $R^2$ (%) |
|---|---|---|---|---|
| Panel A: 3-month futures-implied Euribor | | | | |
| $\Delta(Euribor)$ | -0.09 | 0.29 | 0.24 | 0.03 |
| | (1.44) | (1.06) | (0.47) | |
| $\lvert\Delta(Euribor)\rvert$ | 1.71 | 1.08 | -1.20 | 0.71 |
| | (29.53) | (4.42) | (-2.66) | |
| Panel B: 3-month futures-implied Euroyen Tibor | | | | |
| $\Delta(Tibor)$ | -0.06 | 0.11 | 2.01 | 0.82 |
| | (1.09) | (0.58) | (5.24) | |
| $\lvert\Delta(Tibor)\rvert$ | 0.61 | 0.26 | 1.53 | 0.55 |
| | (12.59) | (1.33) | (4.11) | |
| Panel C: 3-month futures-implied Sterling Libor | | | | |
| $\Delta(SterlingLibor)$ | -0.11 | -0.41 | -0.05 | 0.03 |
| | (1.37) | (1.10) | (0.07) | |
| $\lvert\Delta(SterlingLibor)\rvert$ | 2.13 | 1.63 | -0.86 | 0.69 |
| | (29.04) | (4.87) | (1.50) | |

Panel A presents the results for Euribor. The first row shows no relation, either economically or statistically, between 3-month Euribor changes implied by futures prices and the ECB announcement dummy. This is not all surprising: there is no reason to think that euro interest rates should on average move in a particular direction when ECB announces its policy decisions. However, the second row shows that the absolute value of Euribor changes is significantly higher on ECB

announcement days, with a coefficient of 1.08 ($t$-statistic = 4.42), which represents a very meaningful effect relative to the mean of 1.71 bps on other days. Since interest rates on average experience significantly larger moves on ECB a-days, we conclude that the ECB is not inert.

Panel B shows that futures-implied yen interest rates are neither affected (first row) by BoJ announcements nor significantly more volatile (second row) on BoJ a-days. Although the absolute value of rate changes is higher on announcement days than on other days, the coefficient on the BoJ a-day dummy is not statistically significant (0.26 bps with a $t$-statistic of 1.33). Thus, based on this test, we cannot rule out the hypothesis that the BoJ is inert. This is not a surprising finding, since Japanese interest rates were already almost zero in 1999 and have remained low ever since. The innovations in BoJ policies in our sample period occur not through (short-term) interest rate changes but through its use of unconventional monetary policy tools.

Finally, panel C contains results for sterling futures-implied rates. Similarly to the euro, the trend for sterling interest rates is downward in this period, as shown in the first row. The key result is again in the second row, where we see that the absolute change in futures-implied sterling interbank rates is 1.64 bps ($t$-statistic = 4.87) higher an BoE a-days (relative to the mean of 2.13 on other days). Like those by the ECB, BoE announcements do not appear to be fully anticipated by investors. Overall, our results suggest that the inertness hypothesis cannot explain the lack of a significant market risk premium for ECB and BoE announcements, while for the BoJ we cannot dismiss this hypothesis.

# 4 Conclusion

In this paper we show that the high average excess returns previously observed for U.S. stocks on days with scheduled FOMC announcements also exist in almost all other stock markets across the globe. Just as in the U.S., the FOMC effect is substantially stronger (roughly doubling) in recent years and, probably combined

with an analogous currency effect, appears to be even stronger for countries that have experienced financial crises in the past two decades.

By contrast, hardly any stock market indices display a similar effect for non-U.S. central bank announcements, not even the domestic stock markets of the announcing banks. This is a puzzling finding, especially since these markets do exhibit domestic inflation or unemployment announcement effects, and since non-U.S. announcements of inflation and employment appear to be associated with (marginally) higher returns even in the U.S. Given that Japan, Germany, and the U.K. are large and systemically important economies, and given that they have independent central banks that purse an active monetary policy, our preliminary conclusion is that the Federal Reserve is unique in its importance to global investors, and that this uniqueness does not simply stem from the size of the U.S. economy and its securities markets. An important direction for future research is to establish why, and exactly in what way, the Fed is special.

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# Appendix

Table A.1

U.S. Announcements: Timing Around the Globe

This table reports country details about the timing of announcement days (a-days) and data coverage. Countries are grouped into four geographical regions and are listed in Column 1. Column 2 and 3 report, respectively, the minimum and maximum time difference between the country local time and the Eastern Time, where the time difference is measured in number of hours $h$. Column 4 reports country trading hours (in local time). Column 5 and 6 present the timing of announcement days around the globe. Announcement days are those trading days when FOMC interest rate decisions (Column 5) and employment/inflation numbers (Column 6) are scheduled for release in the United States. According to the time zone, country-i a-days either coincide with U.S. a-days (Same) or are led by one period (Next). The last column reports the starting date of the return series on each Datastream Country Global Equity Index.

| Country | Time Difference (h) | | Trading Hours | A-days Timing | | Data Coverage |
|---|---|---|---|---|---|---|
| | Min | Max | (Local Time) | FOMC | Emp/Infl | |
| **Europe** | | | | | | |
| Austria | 5 | 7 | 08:55-17:35 | Same | Same | 02-Jan-73 |
| Belgium | 5 | 7 | 09:00-17:40 | Same | Same | 02-Jan-73 |
| Denmark | 5 | 7 | 9:00-17:00 | Same | Same | 03-Jan-73 |
| Finland | 6 | 8 | 10:00-18:30 | Same | Same | 28-Mar-88 |
| France | 5 | 7 | 9:00-17:35 | Same | Same | 02-Jan-73 |
| Germany | 5 | 7 | 09:00-17:30 | Same | Same | 02-Jan-73 |
| Greece | 6 | 8 | 10:30-17:00 | Same | Same | 02-Jan-90 |
| Ireland | 4 | 6 | 08:00-16:30 | Same | Same | 02-Jan-73 |
| Italy | 5 | 7 | 9:05-17:35 | Same | Same | 02-Jan-73 |
| Netherlands | 5 | 7 | 09:00-17:30 | Same | Same | 02-Jan-73 |
| Norway | 5 | 7 | 9:00-16:20 | Same | Same | 03-Jan-80 |
| Poland | 5 | 7 | 09:00-16:50 | Same | Same | 02-Mar-94 |
| Portugal | 4 | 7 | 08:00-16:30 | Same | Same | 03-Jan-90 |
| Spain | 5 | 7 | 09:00-17:30 | Same | Same | 03-Mar-87 |
| Sweden | 5 | 7 | 09:00-17:30 | Same | Same | 05-Jan-82 |
| Switzerland | 5 | 7 | 09:00-17:30 | Same | Same | 02-Jan-73 |
| United Kingdom | 4 | 6 | 8:00-16:30 | Same | Same | 02-Jan-73 |
| **North America** | | | | | | |
| Canada | -1 | 0 | 9:30-16:00 | Same | Same | 02-Jan-73 |
| United States | 0 | 0 | 9:30-16:00 | Same | Same | 02-Jan-73 |

[continued]

| Country | Time Difference (h) | | Trading Hours | A-days Timing | | Data Coverage |
|---|---|---|---|---|---|---|
| | Min | Max | (Local Time) | FOMC | Emp/Infl | |
| **Asia Pacific** | | | | | | |
| Australia | 14 | 16 | 10:00-16:00 | Next | Next | 02-Jan-73 |
| Hong Kong | 12 | 14 | 9:30-12:30, 14:30-16:00 | Next | Next | 02-Jan-73 |
| Japan | 13 | 14 | 9:00-11:30, 12:30-15:00 | Next | Next | 02-Jan-73 |
| New Zealand | 17 | 18 | 10:00-16:45 | Next | Next | 05-Jan-88 |
| Singapore | 11.5 | 13 | 9:00-12:30, 14:00:-17:00 | Next | Next | 02-Jan-73 |
| **Emerging** | | | | | | |
| Argentina | 1 | 3 | 11:00-17:00 | Same | Same | 03-Aug-93 |
| Brazil | 1 | 3 | 11:00-18:00 | Same | Same | 05-Jul-94 |
| Chile | 0 | 2 | 09:00-17:30 | Same | Same | 04-Jul-89 |
| China | 12 | 13 | 09:30-11:30, 13.00-15.00 | Next | Next | 27-Jul-93 |
| Indonesia | 11 | 12 | 09:30-12:00, 13.30-16.00 | Next | Next | 03-Apr-90 |
| Malaysia | 11.5 | 13 | 09:00-12:30, 14.30-17.00 | Next | Next | 03-Jan-86 |
| Mexico | -2 | -1 | 08:30-15:00 | Same | Same | 11-May-89 |
| Philippines | 12 | 14 | 09:30-12:10 | Next | Next | 09-Nov-88 |
| South Africa | 6 | 7 | 09:00-17:00 | Next | Next | 02-Jan-73 |
| South Korea | 13 | 14 | 09:00-15:00 | Next | Next | 10-Sep-87 |
| Taiwan | 12 | 13 | 9:55-12:30, 14:45-16:40 | Next | Next | 03-May-88 |
| Thailand | 11 | 12 | 09:00-13:30 | Next | Next | 05-Jan-87 |
| Turkey | 6 | 9 | 09:30-12:30, 14:00-17:30 | Same | Same | 13-Jun-89 |
| Venezuela | -0.5 | 1 | | Same | Same | 03-Jan-90 |

Table A.2
Country Major Stock Market Indices

This table reports the name and the data coverage (i.e. first observation available) of the major stock market index of each country listed in Column 1. All indices are denominated in the domestic currency and are obtained from Datastream.

| Country | Equity Index | Data Coverage |
|---|---|---|
| Brazil | Brazil Bovespa | 22-Dec-89 |
| Canada | S&P/TSX Composite Index | 31-Dec-68 |
| Chile | IGPA Index | 05-Jan-87 |
| France | FTSE France | 31-Dec-86 |
| Germany | DAX 30 Performance | 01-Jan-65 |
| Hong Kong | Hang Seng | 24-Nov-69 |
| Indonesia | IDX Composite | 05-Apr-83 |
| Japan | Topix | 02-Jan-58 |
| Japan | Nikkei 225 Stock Average | 02-Jan-58 |
| Malaysia | FTSE BURSA MALAYSIA KLCI | 03-Jan-80 |
| Mexico | IPC Bolsa | 05-Jan-88 |
| Philippines | Philippine Stock Exchange index (PSEi) | 03-Jan-86 |
| Poland | Warsaw General Index | 17-Apr-91 |
| South Korea | KOSPI Index | 01-Jan-75 |
| Spain | Madrid Stock Exchange General Index (IGBM) | 03-Jan-74 |
| Taiwan | TAIEX | 06-Jan-71 |
| United Kingdom | FTSE All Share | 26-Dec-68 |
| Venezuela | Venezuela Stock Exchange General Index | 02-Apr-93 |

Table A.3
Scheduled Releases of Domestic Announcements

This table reports details about domestic scheduled releases of interest rate decisions, employment numbers or inflation numbers in three major countries, namely Japan, Germany and the United Kingdom (Column 1). Announcements are defined in Column 2. For each announcement the total number of releases, the sample period between the first and last release and details about the available data are reported, respectively, in Columns 3, 4 and 5. Data sources are in Column 6.

| | Announcement | Obs | Period | Scheduled Releases | Source |
|---|---|---|---|---|---|
| Japan | Monetary Policy | 295 | 1973-1995 | Sparse. BoJ was independent, but very weak in legal status. | Bank of Japan |
| | | | 1996-1997 | Not available. | |
| | | | 1998-2013 | Monthly. Independent Bank of Japan. | |
| | Inflation | 528 | 1970-2013 | Monthly. Consumer Price Index. | Statistics Bureau of Japan |
| | Employment | 528 | 1950-2013 | Monthly. From "Labour Force Survey". | Statistics Bureau of Japan |
| Germany | Monetary Policy | 825 | 1948-2013 | Announcements were made every two weeks. | Bundesbank |
| | Inflation | 479 | 1974-2013 | Final CPI from Jan 1974, preliminary CPI from Sep 1976. | Federal Statistical Office |
| | Employment | 371 | Feb1983-Dec2013 | | Federal Employment Agency |
| UK | Monetary Policy | 249 | Prior to 1992 | No formal meetings. | UK Treasury |
| | | | Dec 1992-May 1997 | Monthly. Regular meetings between the Chancellor and the Governor (Monthly Monetary Meetings). The former was not obliged to accept the bank's advice; financial markets were able to see the decision with a certain time lag. | Bank of England |
| | | | June 1997-2013 | Monthly. Independent Bank of England. | |
| | Inflation | 389 | 1981-2013 | Monthly. Retail Price Index (prior to November 2003), Consumer Price Index (afterwards). | ONS |
| | Employment | 350 | Feb 1984-Jun 2001 | Monthly. Only unemployment releases are available prior to Dec-1984. Both unemployment and employment releases are available afterwards. They were announced independently from Dec-1984 to May 1986 and jointly from June 1986. Releases are from: "Employment Gazette" (Feb 1984-Oct 1995) and "Labour Market Trends" (Nov 1995-Jun 2001). | Hard copies |
| | | | Jul 2001-Dec 2013 | Monthly. From: "Labour market trends" (July 2001-Dec 2006), "Economic and Labour Market Review" (Jan 2007-Apr 2011) and "Labour Market Statistics" (Jan 2008-onwards). | ONS website |

Table A.4
Bundesbank decisions were announced in the afternoon or the next morning.
Assumption here: All Bundesbank announcements were made the next morning.

| | (1) Jan1958-Dec1998 | | | | | (2) Jan1958-Dec1996 | | | | | (3) Jan1958-Dec1992 | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | A-days | | Nona-days | | | A-days | | Nona-days | | | A-days | | Nona-days | | |
| Country | Mean | Obs | Mean | Obs | Diff. | Mean | Obs | Mean | Obs | Diff. | Mean | Obs | Mean | Obs | Diff. |
| A: Europe | | | | | | | | | | | | | | | |
| Austria | 0.97 | 599 | 0.33 | 4261 | 0.63 | 1.26 | 550 | 0.20 | 3961 | 1.06 | 2.15 | 453 | -0.08 | 3319 | 2.23 |
| | [0.30] | | [0.26] | | [0.18] | [0.39] | | [0.16] | | [0.30] | [0.58] | | [-0.05] | | [0.56] |
| Belgium | 4.18 | 628 | 0.76 | 4433 | 3.42 | 4.61 | 579 | -0.28 | 4133 | 4.89 | 5.04 | 482 | -0.76 | 3491 | 5.81 |
| | [1.29] | | [0.63] | | [0.99] | [1.41] | | [-0.23] | | [1.40] | [1.35] | | [-0.55] | | [1.46] |
| Denmark | 3.23 | 613 | 2.77 | 4335 | 0.46 | 4.43 | 564 | 2.24 | 4035 | 2.19 | 6.48 | 467 | 1.72 | 3394 | 4.76 |
| | [0.76] | | [1.92] | | [0.10] | [0.99] | | [1.54] | | [0.46] | [1.24] | | [1.04] | | [0.87] |
| Finland | -1.96 | 279 | 0.66 | 1826 | -2.62 | -2.71 | 232 | -2.57 | 1544 | -0.14 | -5.77 | 141 | -6.53 | 930 | 0.76 |
| | [-0.28] | | [0.22] | | [-0.34] | [-0.38] | | [-0.88] | | [-0.02] | [-0.78] | | [-1.86] | | [0.09] |
| France | 9.17 | 628 | 1.97 | 4444 | 7.20 | 11.39 | 579 | 1.67 | 4144 | **9.72** | 14.25 | 482 | 1.68 | 3503 | **12.57** |
| | [2.18] | | [1.20] | | [1.60] | [2.63] | | [1.00] | | **[2.09]** | [2.95] | | [0.89] | | **[2.42]** |
| Germany | 3.91 | 779 | -1.06 | 5744 | 4.97 | 5.31 | 734 | -2.08 | 5461 | **7.39** | 8.41 | 641 | -3.28 | 4848 | **11.70** |
| | [1.06] | | [-0.76] | | [1.26] | [1.52] | | [-1.51] | | **[1.97]** | [2.29] | | [-2.24] | | **[2.96]** |
| Greece | 34.07 | 225 | 9.08 | 1510 | **24.99** | 42.79 | 179 | 4.65 | 1225 | **38.14** | 36.80 | 92 | 8.13 | 638 | **28.68** |
| | [3.45] | | [1.84] | | **[2.26]** | [4.24] | | [0.88] | | **[3.35]** | [2.36] | | [0.93] | | **[1.60]** |
| Ireland | 1.74 | 626 | 4.32 | 4441 | -2.59 | 3.09 | 577 | 3.58 | 4141 | -0.50 | 3.08 | 481 | 3.25 | 3500 | -0.17 |
| | [0.48] | | [2.72] | | [-0.65] | [0.86] | | [2.21] | | [-0.13] | [0.76] | | [1.77] | | [-0.04] |
| Italy | 16.89 | 683 | -3.30 | 4905 | **20.20** | 17.01 | 636 | -4.42 | 4618 | **21.43** | 22.82 | 540 | -4.78 | 4003 | **27.60** |
| | [3.71] | | [-1.68] | | **[4.07]** | [3.68] | | [-2.21] | | **[4.25]** | [4.53] | | [-2.19] | | **[5.03]** |
| The Netherlands | 2.47 | 627 | 2.40 | 4443 | 0.07 | 2.13 | 578 | 1.76 | 4143 | 0.37 | 3.42 | 481 | 0.70 | 3501 | 2.72 |
| | [0.71] | | [1.69] | | [0.02] | [0.66] | | [1.27] | | [0.11] | [0.94] | | [0.44] | | [0.68] |
| Norway | -3.38 | 459 | 2.85 | 3105 | -6.23 | -2.31 | 410 | 2.73 | 2805 | -5.04 | -2.35 | 313 | 1.41 | 2164 | -3.76 |
| | [-0.57] | | [1.08] | | [-0.96] | [-0.37] | | [0.98] | | [-0.73] | [-0.30] | | [0.41] | | [-0.44] |
| Portugal | 3.54 | 250 | -1.08 | 1585 | 4.62 | 4.16 | 206 | -3.95 | 1305 | **8.11** | -2.49 | 112 | -10.19 | 704 | 7.70 |

[continued]

| | (1) Jan1958-Dec1998 | | | | | (2) Jan1958-Dec1996 | | | | | (3) Jan1958-Dec1992 | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | A-days | | Nona-days | | | A-days | | Nona-days | | | A-days | | Nona-days | | |
| Country | Mean | Obs | Mean | Obs | Diff. | Mean | Obs | Mean | Obs | Diff. | Mean | Obs | Mean | Obs | Diff. |
| Spain | -1.36 | 555 | -1.13 | 3461 | -0.23 | -1.54 | 510 | -2.21 | 3179 | 0.68 | -1.40 | 416 | -3.72 | 2570 | 2.32 |
| | [-0.31] | | [-0.59] | | [-0.05] | [-0.35] | | [-1.15] | | [0.14] | [-0.28] | | [-1.68] | | [0.42] |
| Sweden | 1.19 | 410 | 3.37 | 2742 | -2.17 | 3.91 | 361 | 2.76 | 2442 | 1.15 | 5.02 | 264 | 1.16 | 1800 | 3.87 |
| | [0.22] | | [1.38] | | [-0.37] | [0.70] | | [1.11] | | [0.19] | [0.73] | | [0.37] | | [0.52] |
| Switzerland | 8.25 | 621 | -1.01 | 4418 | **9.26** | 9.26 | 572 | -1.68 | 4118 | **10.94** | 11.59 | 475 | -2.52 | 3476 | **14.12** |
| | [2.55] | | [-0.80] | | **[2.67]** | [2.94] | | [-1.37] | | **[3.23]** | [3.29] | | [-1.86] | | **[3.73]** |
| United Kingdom | 0.46 | 718 | 0.33 | 5000 | 0.13 | 1.88 | 669 | 0.17 | 4712 | 1.71 | 3.18 | 573 | 0.16 | 4095 | 3.02 |
| | [0.12] | | [0.22] | | [0.03] | [0.47] | | [0.11] | | [0.40] | [0.70] | | [0.09] | | [0.62] |
| B: North America | | | | | | | | | | | | | | | |
| Canada | 1.72 | 717 | -1.11 | 4967 | 2.83 | 1.38 | 668 | -1.08 | 4681 | 2.46 | 1.38 | 571 | -1.79 | 4071 | 3.17 |
| | [0.64] | | [-0.97] | | [0.97] | [0.50] | | [-0.94] | | [0.83] | [0.45] | | [-1.42] | | [0.95] |
| United States | 4.40 | 977 | 1.90 | 7207 | 2.51 | 4.35 | 928 | 1.70 | 6921 | 2.65 | 4.78 | 831 | 1.36 | 6307 | 3.43 |
| | [2.00] | | [2.00] | | [1.04] | [1.96] | | [1.80] | | [1.10] | [1.99] | | [1.34] | | [1.31] |

# 5 | Human Capital, Unemployment Risk, and Asset Prices*

## 1 Introduction

Empirical studies find that human wealth, the present discounted value of future labour income, accounts for about 90% of total wealth of U.S. households (Lustig, Van Nieuwerburgh and Verdelhan, 2013). Despite the importance of embedding this component in both empirical and theoretical asset pricing work has been recognised since Mayers (1972) and Roll (1977), the literature has focused mainly on financial rather than human wealth over time. A major challenge is posed by human capital being a non-tradable asset. This makes the treatment of human wealth cumbersome, as cash-flows are observable (wages), but discount rates are not. To the best of my knowledge, researchers have taken two approaches to address this issue over time. This paper proposes an alternate perspective: It sheds lights on the riskiness of human capital by empirically pinning down its determinants in the labour market.

Early empirical asset pricing work imposes ad-hoc restrictions on human capital discount rates (Campbell, 1996; Shiller, 1995; Jagannathan and Wang, 1996).[63] Recent work adopts VAR techniques to recover the human capital discount rate

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[63]The conditional expected human return is assumed to be constant (Shiller, 1995), equal to the conditional expected return on financial wealth (Campbell, 1996), constant and labour income growth as unforecastable (Jagannathan and Wang, 1996).

from aggregate consumption data conditioning on a set of standard predictors for consumption growth and stock returns. Under this approach, earlier measures are shown to be inconsistent with the data (Lustig and Van Nieuwerburgh, 2008), and reliable economic inference depends on appropriately incorporating a measure of macroeconomic uncertainty (Bansal, Kiku, Shaliastovich and Yaron, 2014).[64]

This paper presents a new approach based on the idea that human capital is a risky asset, owned by workers and invested in the labour market. Under these premises, human capital risk can be mapped into labour market risk and empirically measured. My starting point is a straightforward insight from labour macroeconomics: At each point in time the labour market equilibrium outcome is jointly determined by wage and unemployment dynamics. I derive a simple measure of human capital that accounts for both macro-variables and use U.S. micro-level data to quantify this measure at aggregate level. I work with representative agents who differ in their endowment of human capital.

Abstracting from governmental transfers, the measure I construct adjusts the growth rate of wages by the unemployment risk borne by employed workers over the same horizon. This measure can also be interpreted as growth in unemployment-(U-) adjusted labour income. To the best of my knowledge, existent work in asset pricing accounts for wages (human capital cash-flows), but ignores unemployment dynamics. This paper contributes to the literature on human wealth by documenting a key role for the unemployment channel beyond the more traditional wage channel.[65] I interpret unemployment as the bad state of the world faced by any employed worker and unemployment risk as the downside risk he bears. I take the probability of being

[64] A separate strand of literature in finance focuses on the consumption-wealth ratio (Campbell and Mankiw, 1989; Lettau and Ludvigson, 2001; Lustig, Van Nieuwerburgh and Verdelhan, 2013). Santos and Veronesi (2006) develop a general equilibrium model that can rationalize the predictability of stock returns by the labour income to consumption ratio. Most recently, Palacios (2014) incorporates investment in human capital in a general equilibrium model and derive a three-factor asset pricing model that depends on the human investment.

[65] Standard asset pricing measures allow unemployment to affect human cash-flows, via unemployment benefits included in national labour income data. Recent human capital measures derived from life-cycle models allow unemployment to affect the human discount rate (e.g. Brown, Fang and Gomes (2012)). However, since these models concern the choices of individual households over the life-cycle, a substantial share of the risks they capture is idiosyncratic rather than systematic. These measures are therefore not useful in an asset pricing framework.

unemployed as a proxy of the worker's exposure to unemployment risk, and propose the unemployment rate as an empirical proxy for this probability at aggregate level.

Figure 1 compares the cumulated quarterly growth in U-adjusted labour income experienced by a representative diploma holder (in blue) and a representative post-college holder (in green) in the United States over the period 1979 to 2014. Both measures are computed using micro-level (CPS) survey data and generalised to the U.S. population via population sampling weights. To improve precision of estimates, I adopt quasi-experimental methods (propensity score weighting) to estimate their wage-related component. For each agent, the graph compares a pure wage-based measure (dashed line) with its U-adjusted counterpart (solid line).

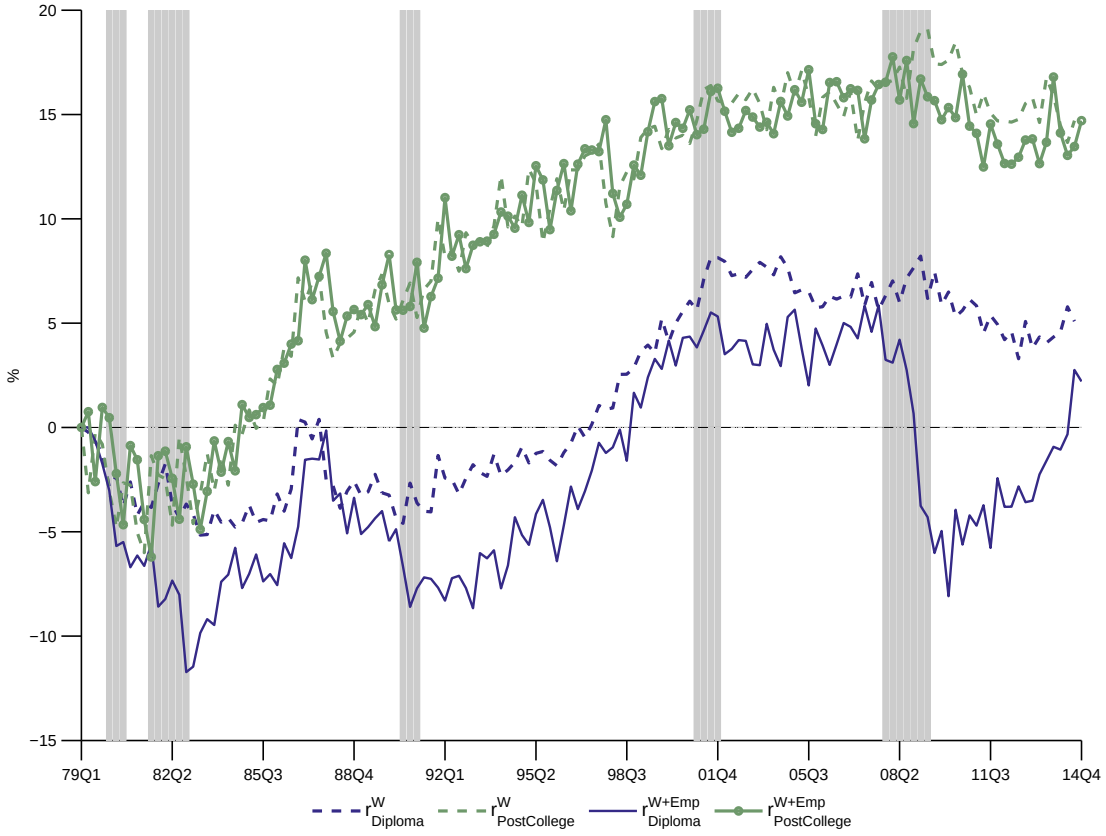
The graph shows that highly educated workers historically experience a less severe drop in U-adjusted labour income in bad times than low educated workers. By definition, a risky asset pays-off poorly in bad times. Since labour income is the pay-off of the human capital asset, the graph points to human capital being historically riskier for low educated agents than highly educated ones. Accounting for unemployment is key to draw this conclusion: Wages are sticky, while unemployment captures a deterioration of labour market conditions in real time. In the graph, the unemployment-adjustment visibly crashes in recessions (shaded areas). This figure is also consistent with the well documented fact that highly educated agents earn historically higher wages than low educated ones on average. Taken together, these findings have implications for human wealth, the present discounted value of future labour income. This evidence points to inequality in human wealth being worse than the well established inequality in labour income.[66] In other words, accounting properly for the riskiness of human capital magnifies inequality in human capital cash-flows and reveals even worse inequality in human capital prices.

I argue that these findings should imply heterogeneous financial investment behaviour. Agents who face high exposure to changes in labour market conditions

[66]Work on trends in the wage structure points to a sharp rise in income inequality in the United States Piketty and Saez (2003). See Autor, Katz and Kearney (2008) for a review of the debate on the forces driving income inequality and weak evidence in support of the fall in the real value of the minimum wage as one-time driver ("revisionists" view (Card and DiNardo, 2002).

Figure 1
Cumulated Human Capital Measures

This figure compares the time-series of cumulated quarterly growth in a pure wage-based measure of human capital ($r_e^W$, dashed line) with its unemployment-adjusted counterpart ($r_e^{W+Emp}$, solid line). Measures are computed for two U.S. representative human capital holders $e$: A diploma holder (blue line) and a post college (green line) holder. The unemployment-adjusted measure is obtained as the sum of growth in wages and growth in employment shares. Measures are in percentage. Shaded areas denote NBER recessions. Data are from monthly samples of the CPS Outgoing Rotation Group. The sample period is 1979:Q1-2014:Q4.




(low educated) should invest in financial markets mainly to hedge their labour income risk – investing in assets that perform well when they are more likely to lose their jobs. On the contrary, agents who face mild exposure (highly educated) should be willing to bear the risk of a poor performance of their financial assets in bad labour market states – investing in risky financial assets – but demand a premium for bearing it.

To this end, I build a macroeconomic factor that captures differential exposure to unemployment risk in the United States as the difference between the growth

in quarterly employment rates of highly educated workers (post college holders) and low educated workers (diploma holders). I show that sorting industry-level portfolios by their time-varying loadings on this factor − controlling for aggregate market risk − yields a novel cross-section of average excess returns. A standard long-short strategy, long in high-$\hat{\beta}^{Emp}$ industries and short in low-$\hat{\beta}^{Emp}$ industries, yields on average a statistically significant premium of about 4% over the last 35 years. Risk premia are driven by bad labour market states (6.5% in bad times vs. 2% in good times). In these quarters unemployment risk is systematic risk, as all types of agents face a higher probability of being unemployed. I also argue that these risk premia are consistent with heterogeneous financial investment behavior, and driven by the demand of highly educated workers.

Finally, I provide some preliminary evidence about the relationship between this novel risk premium and the well-known value premium. Sorting the same set of test assets by time-varying exposure to the HML factor reveals that the two sources of risk are closely related, although they do not fully overlap.

The paper is organised as follows. In the next section I derive a novel unemployment-adjusted measure of human capital and extend it to a related measure of returns to education. In Section 3 I present data and the propensity score methodology. In Section 4, I document patterns in unemployment probabilities of heterogeneous agents and discuss empirical results from U.S. labour and financial markets. Section 5 concludes highlighting avenues for further research.

# 2 Unemployment-Adjusted Human Capital Measures

In this section I derive a novel unemployment-adjusted measure of human capital from the outcome equilibrium in the labour market and extend it to obtain a related measure of returns to education.

**Unemployment-adjusted labour income - $\mathbf{R^e_{t+1}}$.** Workers in the economy differ in their endowment of educational capital $e$. The investment in education is made prior to entering the labour market: Subject to their preferences, workers

study up to a diploma, college or post college degree. At any point in time, the labour force amounts to the set of all full-time employed (E) and unemployed (U) workers in the economy endowed with one of these degrees. The population P is restricted to the labour force, that is $P_t = \sum_{e=1}^{3}(E_t^e + U_t^e)$.

Wages $W_t$ and the employment rate $\Theta_t$, where $\Theta_t = \frac{E_t}{E_t+U_t}$, are determined each period by bargaining between firms and workers subject to the realisation of labour market shocks. Any worker is characterised by educational capital $e$, a labour market status (employed/unemployed) and a pay-off.

The labour market outcome equilibrium is the aggregate payoff for the labour market - the sum of payoffs of employed and unemployed workers - at a given point in time. Under the assumption that pay-offs are earned by a representative investor, the labour market outcome equilibrium is obtained as the sum of the share of employed workers times their aggregate payoff and the share of unemployed workers times their aggregate payoff. Abstracting for governmental transfers, I attach a zero pay-off to any unemployed worker and an aggregate wage to the representative worker[67]. Figure 2 - Panel A provides a graphical description of this setup.

There are three representative workers (i.e., $e$ -types) in the economy: A diploma holder, a college holder and a post-college holder. I define growth in unemployment-adjusted labour income as the ratio of labour market outcome equilibria at two different points in time. For any given representative worker $e$, this measure is equal to:

$$R_{t+1}^e = \frac{\Theta_{t+1}^e \times W_{t+1}^e + (1-\Theta_{t+1}^e) \times 0}{\Theta_t^e \times W_t^e + (1-\Theta_t^e) * 0} = \frac{\Theta_{t+1}^e}{\Theta_t^e} \times \frac{W_{t+1}^e}{W_t^e}, \quad (1)$$

where $\Theta_t^e$ denotes the share of workers with endowment $e$ who are employed at time $t$ and $W_t^e$ is the wage of the representative worker $e$ in the same period. In log form (lowercase):

$$r_{t+1}^e = (\theta_{t+1}^e - \theta_t^e) + (w_{t+1}^e - w_t^e). \quad (2)$$

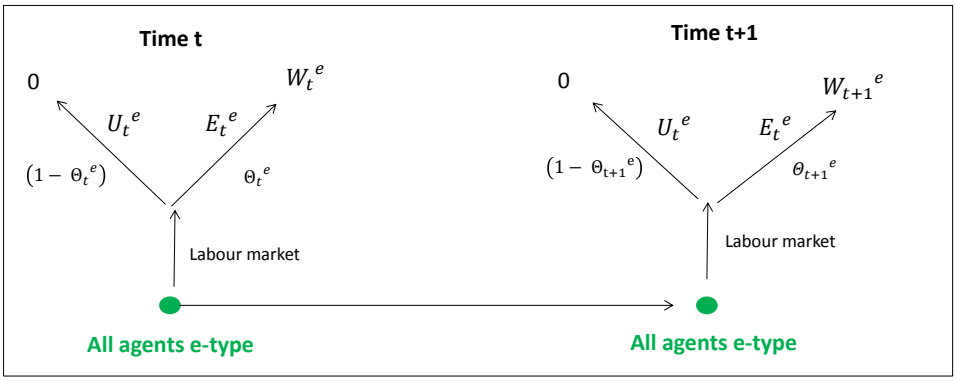
[67] Abstracting for governmental transfers is an over-simplifying assumption that I plan to address in future work. Households in the ORG are not asked to disclose unemployment benefits.

Figure 2
Human Capital Measures

This figure presents a graphical representation of the set up presented in the main text that provides foundation for two novel human capital measures derived from the outcome equilibrium in the labour market. These measures are defined in Equation (1) and Equation (4) in the main text and are reported at the top of each panel, respectively. Panel B presents also two existent measures of returns to education, namely the "skill premium" and its risk-adjusted version examined by Palacios-Huerta (2003). $E_t$: Full-time employed workers at time $t$ ; $U_t$: Full-time unemployed workers at time $t$ ; W: Wage; $\Theta_t^e = \frac{E_t^e}{E_t^e + U_t^e}$: employment rate of workers with educational attainment $e$ at time $t$ .

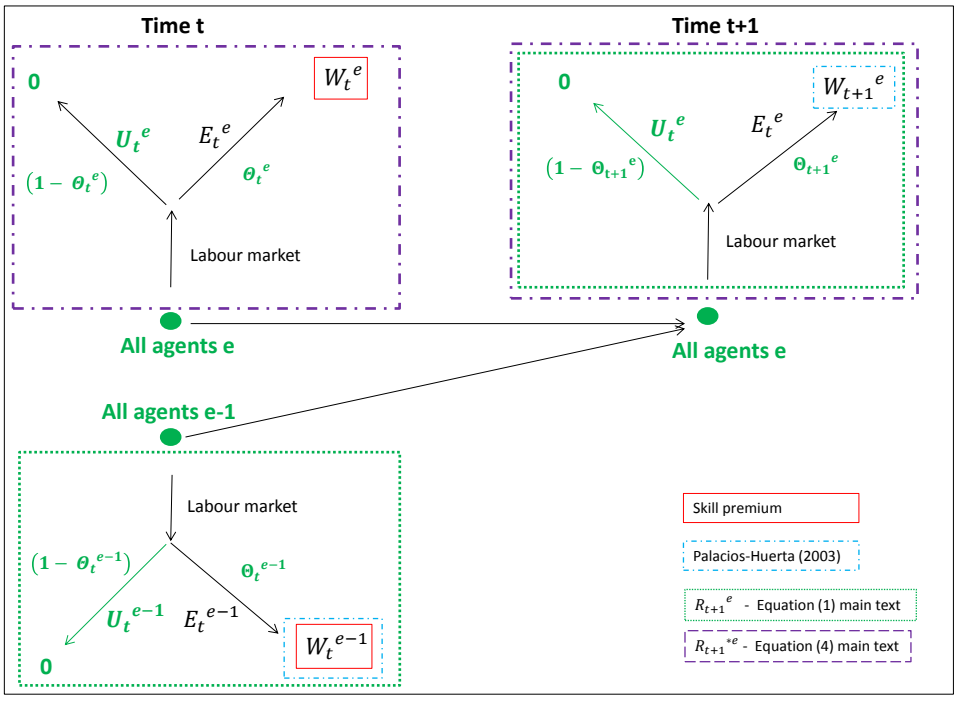
Panel A

$$R_{t+1}^e = \frac{\Theta_{t+1}^e}{\Theta_t^e} * \frac{W_{t+1}^e}{W_t^e}$$




Panel B

$$R_{t+1}^{*e} = \frac{\Theta_t^e}{\Theta_t^{e-1}} * \frac{\Theta_{t+1}^e}{\Theta_t^e} * \frac{W_t^e}{W_t^{e-1}} * \frac{W_{t+1}^e}{W_t^e} = R_{t+1}^e * \frac{\Theta_t^e}{\Theta_t^{e-1}} * \frac{W_t^e}{W_t^{e-1}}$$



This human capital measure adjusts the growth rate of wages of worker $e$ (second term) by the growth rate of his employment probability (first term). I interpret the probability of being unemployed as the worker's exposure to unemployment risk and the unemployment rate as an empirical proxy for this probability. At individual level the unemployment probability is subjective, and likely to depend on the own past history of the worker as well. Under the assumption that past histories of workers are idiosyncratic and average out in the aggregate, the unemployment probability of the representative investor $e$ approximates to $1 - \Theta^e$, the unemployment rate conditional on being endowed with educational human capital $e$.

I interpret unemployment risk as the downside risk any employed worker faces to become unemployed. Unemployment is the bad state for any human capital holder, as the human capital asset does not pay a full dividend and its price crashes. Intuitively, the longer the time spent in the unemployment state, the less attractive the worker becomes for firms and the more his barganing position weakens. In turn, the "price" at which he can sell themselves in the labour market decreases and the value of his human capital asset falls.

Since the unemployment rate is always positive, unemployment risk matters for everyone. The human capital measure in Equation (2) accounts for this phenomenon by adjusting the growth rate of human capital cash-flows over a period by the exposure to unemployment risk borne by the investor over the same horizon. When the unemployment rate of worker $e$ increases, the employment share $\theta^e$ decreases. Adverse changes in labour market conditions are captured by the unemployment-adjustment term $(\theta^e_{t+1} - \theta^e_t)$ being negative.

The human capital measure in Equation (2) can be interpreted as a key-state variable for human capital, as it contains the two main determinants of human wealth, wages and unemployment. The unemployment-adjustment term empirically captures the unemployment channel, that is unemployment having an impact on marginal utility that goes beyond the lost-wage impact. Accounting for unemployment is important, as wage-earners are by definition employed workers. This means that the second term in Equation (2) only captures fluctuations in actual wages in the

labour market. But, unemployment, not wages, is a key indicator of labour market conditions. Unemployment probabilities capture the share of the labour force that is in the bad human capital state and can therefore be interpreted as an indirect way to measure empirically changes in the price of human capital in the economy.

**Return to education - $\mathbf{R^{*e}_{t+1}}$.** The set up above can be extended to explore the relative gains in the economy by comparing payoffs of agents who differ in their human capital endowment. The return to an additional unit of educational human capital earned by worker $e$ is:

$$R^{*e}_{t+1} = \frac{\Theta^e_{t+1} \times W^e_{t+1} + (1 - \Theta^e_{t+1}) \times 0}{\Theta^{e-1}_t \times W^{e-1}_t + (1 - \Theta^{e-1}_t) \times 0} = \frac{\Theta^e_{t+1}}{\Theta^{e-1}_t} \times \frac{W^e_{t+1}}{W^{e-1}_t}, \quad (3)$$

where $e$ and *e-1* denote two educational endowment - say post-college and college degree - and the star differentiates returns to education from those defined in Equation (2).

Multiplying and dividing the first term in Equation (3) by $\Theta^e_t$ and the second term by $W^e_t$ yield:

$$R^{*e}_{t+1} = \underbrace{\frac{\Theta^e_t}{\Theta^{e-1}_t}}_{\text{(a)}} \times \underbrace{\frac{\Theta^e_{t+1}}{\Theta^e_t}}_{\text{(b)}} \times \underbrace{\frac{W^e_t}{W^{e-1}_t}}_{\text{(c)}} \times \underbrace{\frac{W^e_{t+1}}{W^e_t}}_{\text{(d)}} = R^e_{t+1} \times \frac{\Theta^e_t}{\Theta^{e-1}_t} \times \frac{W^e_t}{W^{e-1}_t}. \quad (4)$$

Terms (b) and (d) are again $R^e_{t+1}$ and are driven by changes in the labour market conditions over time. Terms (a) and (c) quantify relative gains between workers who differ in their educational human capital at a given point in time.

Term (c) is the traditional object of study in the labour literature on the returns to education and is well known as the "wage premium" or "skill premium". The term "premium" stems from the evidence that wages are increasing in the educational attainment. This definition has foundation in the human capital theory of Becker (1964), who derived the optimal investment in education by maximising the expected present value of the lifetime stream of wages earned by a representative investor. As emphasised by Becker himself later on, this measure has the drawback of not accounting for the riskiness of the investment in education. Wage premia are not a risk-adjusted measure of human capital returns, as they simply amount to a comparison of average levels, disregarding the role of risk.

Terms (c) and (d) jointly adjust the “skill premium" to account for earnings risk. The risk properties of these two terms are jointly studied in Palacios-Huerta (2003a).[68] Henceforth, I refer to this measure as a cash-flow measure or wage measure of return to human capital.

Equation (4) is an extension of this measure that incorporates unemployment dynamics. It is an enriched measure of returns to education, as it accounts for two extra terms. Term (a) refers to gains/losses attributable to differences in the exposure to unemployment risk due to a distinct endowment of educational human capital. Term (b) adjusts this quantity to account for the change in the exposure to unemployed risk experienced by the worker who owns the additional human capital unit over the period. Figure 2 - Panel B graphically summarises these return measures by extending the setup presented in Panel A.

Contrary to previous contributions in the labour literature, returns in Equation (4) are derived taking into account the entire labour force, rather than being conditional on wage earners (i.e., a full-time employed workers).

The notation and semantic used henceforth are summarised in the following equation (log-form):

$$r^{*e}_{t+1} = \underbrace{(\theta^{e}_{t} - \theta^{e-1}_{t})}_{\substack{\text{Employment Premium}\\ \text{(a)}}} + \underbrace{(\theta^{e}_{t+1} - \theta^{e}_{t})}_{\substack{\text{Employment Growth}\\ \text{(b)}}} + \underbrace{(w^{e}_{t} - w^{e-1}_{t})}_{\substack{\text{Wage Premium}\\ \text{(c)}}} + \underbrace{(w^{e}_{t+1} - w^{e}_{t}).}_{\substack{\text{Wage Growth}\\ \text{(d)}}} \quad (5)$$

# 3 Data and Methodology

Labour market data are from the Earner Study, also known as the CPS (Current Population Survey) Outgoing Rotation Group (ORG). The CPS is the major source of labour market statistics in the United States: This survey covers about 60,000 households (approximately 110,000 individuals) conducted monthly by the Bureau

[68] In evaluating the robustness of the conditional CAPM by Jagannathan and Wang (1996), Palacios-Huerta (2003a) finds that an aggregate version of the return measure examined in Palacios-Huerta (2003b) explains about 70% of the cross-sectional variation in 100 beta-size portfolios of U.S. stocks.

of Labor Statistics (BLS) since 1979.[69] The ORG is the monthly subset of these households (about one-fourth) asked to report details about earnings in addition to employment status and demographics.[70] All monthly micro-level extracts used in this paper are from the NBER website and cover the period from 1979 to 2014.[71]

For the purpose of this paper, this dataset is ideal as data on both wages and labour market status are collected jointly at the monthly frequency along with demographics and educational attainment. I adopt CPS population sampling weights to generalize survey data to the U.S. labour force and recover a proxy of the U.S. labour market outcome equilibrium at any point in time.

The set of employed workers in my dataset is the one traditionally examined in labour market studies on the wage premia. For data cleanings, I follow Katz and Murphy (1992). I retain individuals who declare themselves as full-time employed workers with potential work experience ranging from 0-40 years. The actual number of years of work experience (*Exp*) is not available in the CPS dataset. Consistent with previous work, I construct this variable combining age and schooling years: *Exp* = min {age - years of schooling - 7, age - 17}. *Exp* is a proxy and does not account for periods spent unemployed. Full time or part time students, self-employed workers, workers without pay and workers with missing weekly earnings are excluded.

Nominal wages are defined as earnings per week (in dollars) and are converted into real terms (2009 dollars) using the Personal Consumption Expenditure (PCE) deflator published by the Bureau of Economic Analysis. CPS nominal wages are top-coded to protect the privacy of respondents and top-coding has been revised twice by BLS over time.[72] I follow Morgan and Cha (2007) and impose the lowest top-code in real terms on all months. To deal with outliers at the bottom of the wage distribution, I remove all individuals that declare nominal weekly earnings below 50 dollars.

[69]The annual March CPS survey was initiated in 1964. Only monthly data are used in this paper, as point-in-time data are needed to infer human capital returns from labour market equilibria at the business cycle frequency.

[70]Each CPS household is interviewed for four consecutive months, ignored for the following eight months and interviewed again for other four consecutive months. The ORG comprises all CPS households in either their fourth or sixteenth month.

[71]Monthly micro-level data are available at: `http://www.nber.org/data/morg.html`.

[72]CPS nominal top-codes are defined as follows: 999 U.S. dollars (1979-1988), 1923 U.S. dollars (1989-1997) and 2884 U.S. dollars (1998-2014).

I extend this standard dataset to also include unemployed workers, that is, all individuals declaring to be full time unemployed and match the requirements above.

Each individual in the final dataset is characterised by the following set of variables: race (white/black/other), sex (male/female), age, potential years of work experience, labour market status (employed/unemployed), education, industry, geographical area, CPS sampling weights (population-based and earnings-based weights) and a real weekly wage (if employed).

Education is measured in schooling years (0-18). It is well known that this classification is not fully consistent over time due to two major changes in the CPS survey questions (in 1992 and 1998, respectively). Before 1992, I define the educational attainment as the highest grade attended and completed. Since 1992, I follow the two classifications proposed by (Jaeger, 1997, 2003) who map the highest degree received into schooling years. I define diploma, college and post college holders as workers with 12, 16 or more than 16 years of schooling, respectively.

I map CPS industries into 10 divisions using SIC (NAICS)-based codes before (after) 2002. I map Census state codes into nine divisions using the first digit of state code as the division code. Details about these two classifications are provided in the Online Appendix.

Macroeconomic variables, namely the U.S. aggregate unemployment rate and the U.S. industrial production index, are from the BLS and FRED Louis websites, respectively.

Financial variables, namely 30 U.S. industry portfolios, the risk-free rate and the excess return on the aggregate U.S. stock market, are from Kenneth French website. These series are monthly and are compounded at the quarterly frequency.

I identify good and bad times in the labour market using the the monthly job posting index compiled by Barnichon (2010), who accounts for both help-wanted advertisements appeared in major newspapers (i.e., “print" measure of vacancies) and those posted over the internet (i.e., “online" measure of vacancies). The latter component ensures the indicator is representative of the recent past. The number of job openings, or vacancies, is a standard pro-cyclical indicator of labour market

conditions: Good times are periods in which firms actively search for new workers by posting vacancies at a high rate. I construct a quarterly series taking the last month of each quarter. I define bad (good) states as quarters in which this series is below (above or equal to) its standardized mean. There are 144 quarters and five NBER recessions in the sample, 66 quarters fall in the bad states and 23 of these quarters are recession periods.

Finally, I define human capital as an asset that is not tradeable but fluctuates in value over time. Human capital holdings are jointly determined by innate abilities (unobservable), educational attainment and years of work experience. I assume that individuals invest in education in the early stages of their life according to their preferences and then enter the labour market endowed with educational human capital. The compensation they get in the labour market each period (if employed) is a joint reward for both the educational investment they have made and their work experience.

I build a cross-section of workers sorting along two characteristics, education and work experience.[73] I restrict the U.S. labour force to workers holding a diploma, a college degree or a post-college degree. These workers are classified as having 1-5, 6-15 or 16-40 experience years. Table 1 shows that the combination of these characteristics results in nine education-experience cells, or equivalently, a cross-section of nine representative human capital holders:

Workers with the lowest human capital endowment in the labour force have a diploma degree and few years of work experience (1-5); workers with the highest human capital endowment in the labour force have a post-college degree and many years of work experience (16-40). I examine human capital measures either at this level of aggregation or by sorting by education only.

[73]Labour market studies on the wage premia usually adopt a finer classification by accounting for demographics (i.e., gender-race-education-experience groups). Palacios-Huerta (2003a) studies the risk properties of cash-flow risk adjusted returns by allocating workers into 90 gender-race-education-experience groups.

Table 1
Human Capital Endowment Matrix

| | | | Working Experience (years) | | |
|---|---|---|---|---|---|
| | | | L (1-5) | M (6-15) | H (16-40) |
| Education | L | Diploma Holders | L-L | L-M | L-H |
| | M | College Holders | M-L | M-M | M-H |
| | H | Post College Holders | H-L | H-M | H-H |

## 3.1 Propensity Score Approach to Wage Premia

A leading approach to estimate the "skill premium" (Equation 5, Terms c) is the Mincer (1974)-Becker (1964) approach that quantifies the average internal rate of human return via panel regressions. Another strand of the literature has relied on the same dataset to examine trends in the wage structure over time. Researchers in this area work in time-series and aggregate wage data in three steps. Differences in average log wages are obtained by bundling together agents with similar demographic characteristics. In each period, (i) agents are allocated to different sex-race-education-experience subgroups,[74] (ii) mean log real weekly wages are computed for each of these subgroups, and (iii) mean wages for broader groups are obtained by taking fixed-weighted averages of the means computed in step (ii). Adjusted wage premia are then studied at this level of aggregation.

There are at least two limitations in using this aggregation procedure: (i) All workers are treated as homogeneous within each sub-group; (ii) the sorting scheme depends on a small set of characteristics due to the standard curse-of-dimensionality problem.

I overcome these two issues by treating the estimation of these quantities as an evaluation problem. For a given educational attainment $e$, I define the treatment group T as the set of full-time employed workers in the dataset at time *t+1*, $T = E^e_{t+1}$, and the treatment variable as their educational attainment $e$. The counterfactual of T is $E^{e-1}_t$, that is the same set of workers of group T but holding

[74]For instance, Palacios-Huerta (2003a) considers 2880 subgroups.

lower educational human capital, *e-1* rather than *e*, in the previous period. This counterfactual is not observable, hence the evaluation problem. I define a control group C as a proxy for this counterfactual, where C is the set of full-time employed workers in the dataset at time *t*, declaring educational attainment *e-1*.

In each quarter, I estimate wage premia and log wage changes, separately, as the average treatment effect on the treated (ATT) using the propensity score methodology (Rosenbaum and Rubin, 1983). This technique is a two-step procedure that uses the propensity scores as weights to create a balanced sample of treated and control observations. The propensity score is the conditional probability of receiving the treatment conditional on a set of observable characteristics. The advantages of this methodology are twofold: (i) It mitigates the curse of dimensionality allowing a larger set of characteristics to be accounted for and (ii) treats all individuals as heterogeneous by assigning individual weights.

The first step of the procedure involves the estimation of a propensity score for each individual in the sample by running a logistic regression conditioning on a set of observable characteristics. I define the following set of predictors: Years of working experience, sex, race (Whites/Blacks), industry dummies, geographical dummies, dummies for broad working experience groups, an interaction term for years of working experience and sex, an interaction term for sex and "Whites", CPS earnings-based sampling weights.

In the second step of the procedure, wage premia/log wage changes are estimated as the average difference in outcomes between the treatment and the control group. Controls are weighted by the inverse of the probability of receiving treatment, that is

$$\hat{\delta}_{ATT} = \underbrace{\frac{1}{N_T} \sum_{i:d_i=1} y_i}_{\text{T}} - \underbrace{\frac{\sum_{i:d_i=0} \hat{r}_i y_i}{\sum_{i:d_i=0} \hat{r}_i}}_{\text{C}}, \tag{6}$$

where $i$ denotes an agent, $y$ is the outcome variable (i.e. the log wage), $d$ is a treatment indicator taking value of 1 if $i$ is treated and 0 otherwise, $N_T$ denotes the number of treated individuals and $\hat{r}_i = \frac{\hat{p}_i}{1-\hat{p}_i}$, where $\hat{p}_i$ is the estimated propensity score of $i$ from the first step. This estimator is unbiased and efficient (Imbens et al., 2003).

Following DuGoff, Schuler and Stuart (2014) I combine the propensity score method and the CPS sampling weights to obtain unbiased treatment effects that are generalizable to the original survey target population. In the first step, I include the CPS earnings-based sampling weights as a predictor in the propensity score model. In the second step, I weight both treated and controls by their CPS population-based sampling weights. For controls, I create combined weights by multiplying their propensity score-based weights by their CPS weights.

Estimates are robust to the trimming of all individuals whose propensity score does not fall within a common support (lower/upper bound given by the minimum/maximum propensity score) and also to discarding all individuals whose propensity score falls in the bottom/top 1% of the propensity score distribution.

# 4 Evidence From Labour and Financial Markets

In this section, I document cross-sectional variation in the exposure to unemployment risk, present time-series evidence on each component of the human capital measure in Equation (5) and examine differences between wage measures and unemployment-adjusted measures. I then derive a macroeconomic factor that captures time-varying heterogeneity in the exposure to unemployment risk in the U.S. labour market. I rely on this factor to examine the risk-return trade off faced by heterogenous and constrained human capital holders when investing in financial markets.

## 4.1 Downside Risk to Human Capital

I interpret unemployment risk as the downside risk faced by any employed worker and the unemployment rate as an aggregate proxy of the probability of being unemployed. In this section, I examine if and the extent to which there is cross-sectional variation in the exposure to this risk in the U.S. economy by computing conditional unemployment rates. Percentage quarterly averages are reported in Table 2. Unemployment rates are obtained by conditioning jointly on a human capital endowment (education in Panel I and education and work experience in Panel II)

and a labour market state. Cross-sectional patterns are in rows, time-series patterns are in columns. Quarterly series are obtained as the last month of each quarter.

Table 2
Average Conditional Unemployment Rates

This table reports average unemployment rates (in percentage) conditioning on states (all periods/bad/good; Columns 1-3) and human capital endowment (education in Panel I; education and work experience in Panel II). Education is measured as Diploma, College or Post College degree. Working experience is measured in years: 1-5 years, 6-15 years or 16-40 years. The last two columns report time-series mean difference tests for the bad over the good state, that is the mean difference between the two states (Column 4) and the standard error (*s.e.*) of this difference (Column 5). The unemployment rate is defined as the ratio of the unemployed workers to the labour force. The bad (good) state refers to quarters in which the job posting index by Barnichon (2010) is below (above or equal to) its standardized mean. Data are from monthly samples of the CPS Outgoing Rotation Group. Quarterly series are obtained as the last month of each quarter. The sample period is 1979:Q1-2014:Q4.

| Cross-Section | Time-Series | | | | |
|---|---|---|---|---|---|
| | All (1) | Bad (2) | Good (3) | Bad−Good (4) | (*s.e.*) (5) |
| | Panel I: By Education | | | | |
| Diploma | 7.13 | 8.68 | 5.82 | 2.86 | (0.39) |
| College | 3.21 | 4.07 | 2.49 | 1.58 | (0.21) |
| Post college | 2.39 | 2.88 | 1.98 | 0.89 | (0.16) |
| | Panel II: By Education And Work Experience | | | | |
| Exp: 1-5 years | | | | | |
| Diploma | 12.29 | 14.45 | 10.46 | 3.99 | (0.76) |
| College | 3.44 | 4.39 | 2.63 | 1.77 | (0.27) |
| Post college | 2.66 | 3.19 | 2.22 | 0.97 | (0.31) |
| Exp: 6-15 years | | | | | |
| Diploma | 8.11 | 9.90 | 6.60 | 3.30 | (0.46) |
| College | 2.95 | 3.73 | 2.29 | 1.44 | (0.21) |
| Post college | 2.27 | 2.71 | 1.90 | 0.81 | (0.16) |
| Exp: 16-40 years | | | | | |
| Diploma | 5.46 | 6.74 | 4.37 | 2.37 | (0.35) |
| College | 3.26 | 4.10 | 2.55 | 1.55 | (0.25) |
| Post college | 2.34 | 2.84 | 1.93 | 0.91 | (0.21) |

Not surprisingly, all rates are positive and statistically different from zero at the 1% level confirming that at any point in time any representative human capital holder faces the risk to end up in the bad state of the world. More interestingly, Column 1 documents cross-sectional variation in the exposure to this risk: i) Full-sample unemployment probabilities are a decreasing function of the

educational attainment (Panel I), ii) these monotonic differences are statistically different from each other (Table 3 - Panel A), and (iii) this pattern is robust when controlling for work experience (Panel II). Over the 1979-2014 period, the average U.S. unemployment rate is 7.2% for diploma holders, about 3.2% for college holders and 2.4% for post-college holders. The average cross-sectional spread in the economy - the difference between junior diploma holders and the most experienced post-college holders - is about 10%.

Bad times in the economy are periods in which labour market condition deteriorates: Firms post less vacancies and aggregate unemployment raises, as predicted by the Beveridge curve. Columns 2, 3 and 4 report, respectively, conditional unemployment rates in bad times, good times and their difference. State-spreads are also monotonically decreasing in the educational attainment - about 3% for diploma holders and less than 1% for post-college holders - statistically significant at the 1% level and patterns are robust when controlling for work experience.[75]

Taken together, this evidence indicates that longer investments in education made before entering the U.S. labour market provide an hedge to workers against unemployment risk. This has twofold implications. On the one hand, at any point in time the human capital holdings of highly educated workers are less exposed to adverse changes in labour market conditions than those of low educated ones. On the other hand, it is in bad times that highly educated workers gain the most relative to low educated workers, as the increase in the riskiness of human capital they experience is less pronounced.

Table 3 reports cross-sectional mean difference tests, that is tests for differences in means along rows of Table 2. Mean differences are in percentage, associated standard errors are in parenthesis. Panel I tests educational mean spreads conditioning on work experience (Column 1: Diploma versus College; Column 2: College versus Post College); Panel II tests experience mean spreads conditioning on education (Column 1: 1-5 versus 6-15 years; Column 2: 6-15 versus 16-40 years); all tests

[75]Standard errors of these time-series mean difference tests are reported in parenthesis in Column 5.

are conditional on a time-dimension (i.e., bad/good state, all periods). Spreads in Panel I correspond to the "employment premium" (Equation 5, Term a).[76]

Table 3
Average Unemployment Rates: Cross-Sectional Mean Difference Tests

This table reports cross-sectional (along rows) mean ($\mu$) difference tests for average unemployment rates reported in Table 2. For each test, the table shows the difference in means (in percentage) and the standard error of this difference (in parenthesis). Tests in Panel I are for Diploma over College holders (Column 1) and College over Post College holders (Column 2) and are conditional on years of working experience (all years/1-5 years/6-15 years/16-40 years). Tests in Panel II are for workers with 1-5 over 6-15 years of work experience (Column 1) and 6-15 over 16-40 years of working experience (Column 2) and are conditional on the educational attainment (Diploma, College, Post College). The unemployment rate is defined as the ratio of the unemployed workers to the labour force. The bad (good) state refers to quarters in which the job posting index by Barnichon (2010) is below (above or equal to) its standardized mean. Data are from monthly samples of the CPS Outgoing Rotation Group. Quarterly series are obtained as the last month of each quarter. The sample period is 1979:Q1-2014:Q4.

Panel I: Conditioning On Work Experience

| Work Experience | (1) $\mu^{Diploma} - \mu^{College}$ | | | (2) $\mu^{College} - \mu^{PostCollege}$ | | |
|---|---|---|---|---|---|---|
| | Bad | Good | All | Bad | Good | All |
| All years | 4.61 | 3.34 | 3.92 | 1.19 | 0.50 | 0.82 |
| | (0.40) | (0.20) | (0.25) | (0.23) | (0.13) | (0.15) |
| 1-5 years | 10.06 | 7.83 | 8.85 | 1.20 | 0.41 | 0.77 |
| | (0.71) | (0.40) | (0.43) | (0.34) | (0.24) | (0.22) |
| 6-15 years | 6.16 | 4.30 | 5.16 | 1.02 | 0.39 | 0.68 |
| | (0.44) | (0.26) | (0.29) | (0.22) | (0.15) | (0.14) |
| 16-40 years | 2.64 | 1.82 | 2.20 | 1.26 | 0.62 | 0.91 |
| | (0.39) | (0.18) | (0.24) | (0.29) | (0.16) | (0.17) |

Panel II: Conditioning on Education

| Education | (1) $\mu^{(1-5)} - \mu^{(6-15)}$ | | | (2) $\mu^{(6-15)} - \mu^{(16-40)}$ | | |
|---|---|---|---|---|---|---|
| | Bad | Good | All | Bad | Good | All |
| Diploma | 4.55 | 3.86 | 4.18 | 3.16 | 2.22 | 2.65 |
| | (0.78) | (0.44) | (0.48) | (0.51) | (0.27) | (0.32) |
| College | 0.66 | 0.34 | 0.48 | -0.36 | -0.26 | -0.31 |
| | (0.29) | (0.18) | (0.19) | (0.29) | (0.16) | (0.18) |
| Post College | 0.48 | 0.32 | 0.39 | -0.13 | -0.02 | -0.07 |
| | (0.28) | (0.21) | (0.18) | (0.22) | (0.15) | (0.14) |

[76]Panel I tests the "employment premium", as an increase in the unemployment rate of diploma holders relative to college holders is equivalent to an increase in the employment rate of college holders relative to diploma holders, that is $(1 - \theta_t^{e-1}) - (1 - \theta_t^e) = \theta_t^e - \theta_t^{e-1}$.

All differences in Panel I except one are statistically significant at the 1% level, indicating that highly educated workers are always better off relative to low educated workers with similar work experience.[77] Findings in Panel II depend on the educational attainment of the worker. For workers who enter the labour market with a diploma the exposure to unemployment risk is decreasing in work experience, while for highly educated ones most of the insurance comes from their prior educational investment.[78]

Although the educational spread - the difference between diploma and post-college holders - is diminishing in work experience, the labour market "employment premium" is mainly compensation for holding human capital in the form of education rather than work experience. As Table 2 shows, the average unemployment rate for a worker with high experience and low education is always greater than the average unemployment rate for a worker with low experience and high education.

The take-away of this exercise is twofold. First, U.S. highly educated workers are always less constrained than U.S. low educated workers, as their human capital holdings are less risky. Second, the U.S. labour market rewards educational investments in proportion to the investment made: The longer the time spent in education, the greater the hedge against unemployment risk.

## 4.2 Dissecting Human Returns Over Time

For each representative investor (diploma/college/post college holder), I construct quarterly time-series of the four return components in Equation 5. In the following, I present time-series evidence on each term. I start dissecting the returns to human capital, $r^e_{t+1}$. The first term, log changes in wages (Term d), captures the growth rate of human capital cash-flows over time (Figure 3 - Panel A). The second term, log changes in employment shares (Term b), quantifies the extent to which the unemployment risk associated to those gains varies over time (Figure 3 - Panel B).

[77] The difference in average unemployment rates between junior diploma and college holder is only marginally significant.

[78] For both college and post-college workers the decrease in unemployment probability is very small in magnitude (around 0.5%) when their work experience increase from 1-5 years to 6-15 years, and not statistically significant in their maturity.

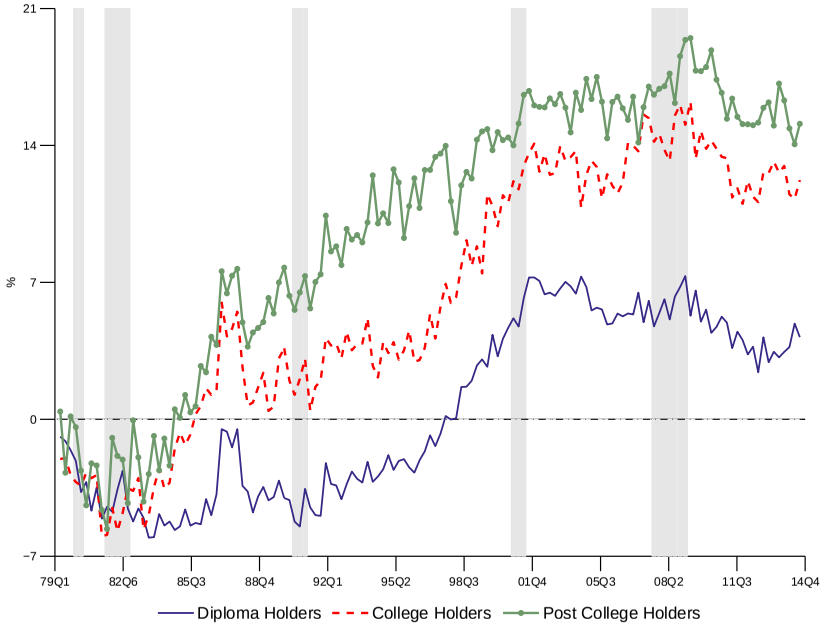
Figure 3

Cumulated Log Changes In Wages And Employment Rates By Educational Attainment

This figure shows the time-series of the cumulated percentage growth in wages (left graph) and employment rates (right graph) for diploma (blue line), college (red dashed line) and post-college holders (green circle line). The employment rate is the ratio of employed workers to the labour force conditional on a given educational attainment $e$ and is expressed in log-form. The wage growth is estimated via propensity score weighting at the quarterly frequency. Wages are expressed in real terms. Shaded areas represent NBER recessions. Data are from monthly samples of the CPS Outgoing Rotation Group. Quarterly series for employment rates are obtained as the last month of each quarter. The sample period is 1979:Q1-2014:Q4.

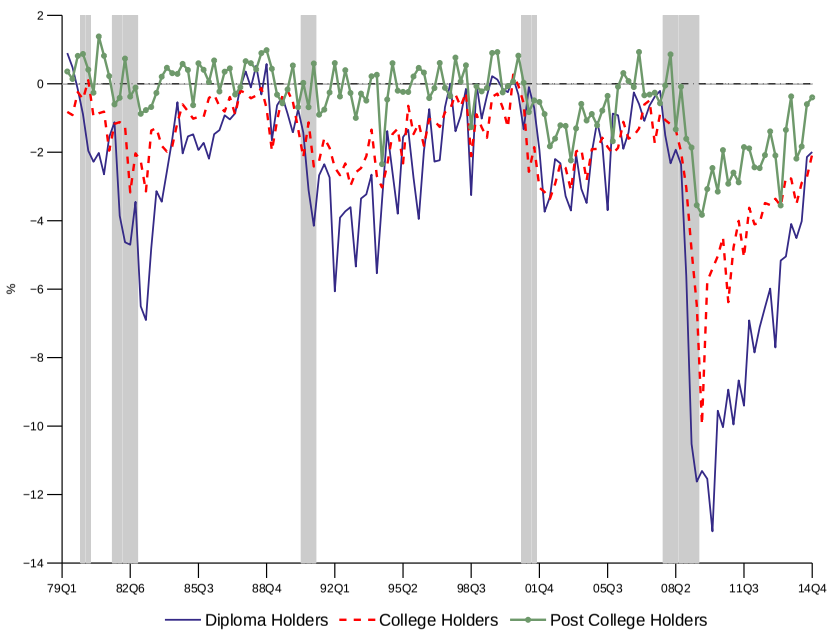
Panel A - Wages

$\sum_{t=1}^{T-1}(w_{t+1}^{e} - w_{t}^{e})$




Panel B - Employment

$\sum_{t=1}^{T-1}(\theta_{t+1}^{e} - \theta_{t}^{e})$



**Log changes in wages (d).** Over the last 35 years, the monetary cumulated returns to human capital amount to about 15% for post college holders, 12% for college holders and only 4% for diploma holders. Series are highly correlated to each other - the pairwise correlations is about 0.85 for diploma and post-college holders and above 0.9 for all other workers - indicating that labour market shocks tend to affect all wages simultaneously, regardless the education of the worker. All series show a downward pattern at the beginning of the sample due to negative shocks to wages stemming from the 1980 recession, a steady increase in the '90s and another downward trend in the final part of the sample, though more pronounced for low educated workers. Cycles reveal that fluctuations have been historically more marked for highly educated workers. This indicates that their wages are higher, but also more volatile.

**Log changes in employment rates (b).** As pointed out by Table 2, swings in unemployment probabilities are more severe the less the worker is educated. Table 4 corroborates this finding by regressing quarterly log changes in educational employment rates on a constant and the log change in either the aggregate U.S. unemployment rate (Panel I) or the U.S. industrial production index (Panel II). A 1% increase in the former is associated to a more than proportional decline in the employment rate of diploma holders (-1.53%), but a less than proportional decline for both college and post-college holders (-0.60% and -0.44%, respectively). Similarly, a 1% quarterly increase in the latter is associated to a 0.35% (0.11%) increase in the employment rate of diploma (college) holders and a positive but not significant increase for post-college holders (0.06%). Overall, this indicates that highly educated workers have been historically less exposed to business-cycle fluctuations, though Figure 3 points out that in the last two recessions they have also suffered more than in the past.

Log changes in the employment rate of diploma holders anticipate those of other workers, but not vice versa. A 1% quarterly increase in the former is followed by a significant 0.2% increase experienced by both college and post-college workers in the next quarter. This points to a structure in the transmission of labour

Table 4
Business Cycle Exposure

This table shows coefficients and associated Newey-West t-statistics obtained by regressing returns to human capital earned by diploma, college and post college holders on a constant and either the log change in the U.S. aggregate unemployment rate (Panel I) or the log change in the U.S. industrial production index (Panel II). Returns are defined as the sum of log changes in employment rates ($\theta^e_{t+1} - \theta^e_t$) and log changes in wages ($w^e_{t+1} - w^e_t$). These components are examined, separately, in the second and third column. Series are quarterly.

$$r^e_{t+1} = \alpha + \beta X_{t+1} + \epsilon_{t+1}$$

| | $r^e_{t+1}$ | | | | $\theta^e_{t+1} - \theta^e_t$ | | | | $w^e_{t+1} - w^e_t$ | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | $\beta$ | $t(\alpha)$ | $t(\beta)$ | $\alpha$ | $\beta$ | $t(\alpha)$ | $t(\beta)$ | a | $\beta$ | $t(\alpha)$ | $t(\beta)$ |
| | I) $X_{t+1} = \Delta U_{t+1}$ | | | | | | | | | | | |
| Diploma | 0.01 | -1.59 | 0.21 | -5.55 | -0.02 | -1.53 | -0.41 | -4.71 | 0.04 | -0.06 | 0.59 | -0.29 |
| College | 0.09 | -0.38 | 1.12 | -1.75 | -0.01 | -0.60 | -0.24 | -2.52 | 0.10 | 0.22 | 1.35 | 1.14 |
| Post College | 0.10 | -0.13 | 1.23 | -0.34 | -0.01 | -0.44 | -0.19 | -3.12 | 0.10 | 0.31 | 1.32 | 0.86 |
| | II) $X_{t+1} = \Delta IP_{t+1}$ | | | | | | | | | | | |
| Diploma | -0.11 | 0.33 | -1.10 | 4.08 | -0.16 | 0.35 | -1.89 | 3.84 | 0.05 | -0.03 | 0.72 | -0.48 |
| College | 0.05 | 0.11 | 0.62 | 1.49 | -0.05 | 0.11 | -1.05 | 2.55 | 0.10 | -0.01 | 1.43 | -0.11 |
| Post College | 0.09 | 0.03 | 0.89 | 0.27 | -0.03 | 0.06 | -0.76 | 1.41 | 0.12 | -0.03 | 1.21 | -0.33 |

market shocks to employment: Low educated workers are hit at first and shocks subsequently reach the highly educated ones.[79]

I now present evidence on the other two components of the returns to education defined in Equation 5 - employment premia (Term a) and wage premia (Term c) - that capture relative differences between the three representative workers at a given point in time.

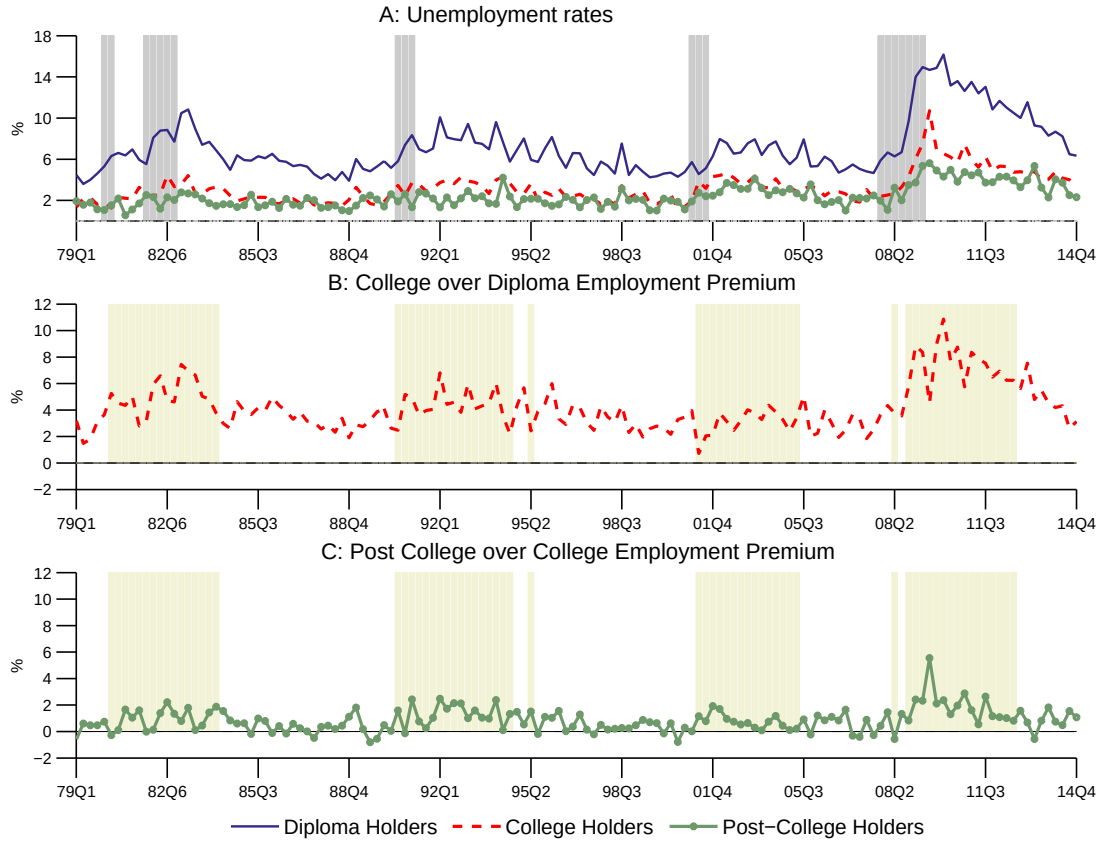
**Employment premia (a).** Figure 4 - Panel A shows the time-series of the unemployment rates of diploma, college and post-college holders. The full sample pairwise correlation is 0.86 for diploma-college holders, 0.82 for college-post college holders and 0.73 for diploma-post college holders. Co-movements are stronger in bad than good labour market states indicating that labour market shocks are aggregate shocks in those states.[80] Put differently, unemployment risk is aggregate risk in the bad state.

[79] Changes in the employment rate of college holders do not anticipate changes in the employment rates of post college holders.

[80] Pairwise correlations between diploma-college, college-post college and diploma-post college holders are, respectively, 0.82, 0.81 and 0.69 in the bad labour market state and 0.77, 0.72 and 0.58 in the good state.

Figure 4
College and Post-College Employment Premia

This figure shows the time-series of the educational unemployment rates (Panel A) and the time-series of the employment premia for college and post college holders (Panel B and C, respectively). The educational unemployment rate is the unemployment rate conditional on the educational attainment $e$, that is diploma, college or post-college degree. The employment premium, $\theta_t^e - \theta_t^{e-1}$, is the log difference in employment rates between workers holding a college degree (post college degree) and those holding a diploma (college degree). All series are in percentage. Shaded areas represent NBER recessions in Panel A and bad labour market times in Panels B and C. Bad times for the labour market are periods in which the job posting index by Barnichon (2010) is above its standardized mean. Data are from monthly samples of the CPS Outgoing Rotation Group. Quarterly series are obtained as the last month of each quarter. The sample period is 1979:Q1-2014:Q4.



Panels B and C present the time-series of the employment premium for college and post college holders, respectively. The employment premium is the log difference in quarterly employment rates between workers holding a college degree (post college degree) and those holding a diploma (college degree) in a given quarter. Shaded areas represent bad times for the labour market.

Employment premia are time-varying: They widen in bad times and shrink in good times. The college employment premium is about 4% on average - 4.5% in bad times and 3.3% in good times - and positive in all quarters of the sample, indicating that college holders have been systematically better off relative to diploma holders.

The post-college employment premium is smaller, about 0.8% on average - 1.2% in bad times and 0.5% in good times - consistent with the hypothesis of diminishing marginal returns to scale to further education. Like wages, also the employment insurance is an increasing function of the educational attainment, but is decreasing in magnitude at the margin (i.e. for any additional unit of education). Although post-college workers are historically better off relative to diploma workers, the post-college premium is zero or even negative (up to -0.8%) in 14% of the quarters indicating that post-college workers can be worse off relative to college workers. This evidence, however, does not affect the novel informational content of the post-college premium. Firstly, the post-college premium in the good state is statistically different from the post college premium in the bad state and both premia are statistically different from zero (Table 3). Secondly, 98% of those observations fall in the good labour market state that is quarters in which unemployment risk is not aggregate risk.

Employment premia predict each other. A 1% increase in the college (post college) premium anticipates a 0.25% (0.85%) increase in the post college (college) premium in the next quarter. Both effects are statistically significant (Newey-West standard errors).

**Wage premia (c).** For descriptive purposes, I start defining a representative worker as a median wage earner. Figure 5 - Panel A shows the time-series of median real weekly earnings in U.S. dollars for representative diploma, college and post college holders. As widely documented, wages are historically an increasing

function of the educational attainment. Earnings of diploma holders have been relatively stable over time, while those of both college and post college holders have been rising. These gains are not a free lunch, though, as they have come at the cost of greater volatility: Median earnings of post college holders are historically about 3.5 times more volatile than those of diploma holders.[81] From an economic perspective, this suggests that labour market shocks transmitted to wages are more severe for workers earning wages at the top rather than at the bottom of the wage distribution. Expected human capital payoffs of highly educated workers are not only higher, but also riskier.

Panel B shows the time-series of the college and post college wage premia estimated via propensity score weighting. As discussed in Section 3.1, this technique accounts for the heterogeneity of the individuals in the sample by assigning individual weights. Both wage premia are positive consistent with the hypothesis of diminishing marginal returns to education. On average, a college (post college) holder has earned 44% (16%) more than a diploma (college) holder over the last 35 years. Overall, both premia steadily increased in the '80s, dropped in the '90s and rose again in the next 15 years, though the college premium at a much lower pace than in the past. Trends are very pronounced for college earners - their premium rose from about 35% to about 50% just in the first part of the sample - while swings are moderate for post college earners - their variation is historically bounded between 10% and 20%. Though the contemporaneous correlation in wage premia is moderate, about 0.37, they predict each other: A 1% increase in the college (post-college) premium anticipates a 0.15% (1.15%) increase in the post college (college) premium next period.[82]

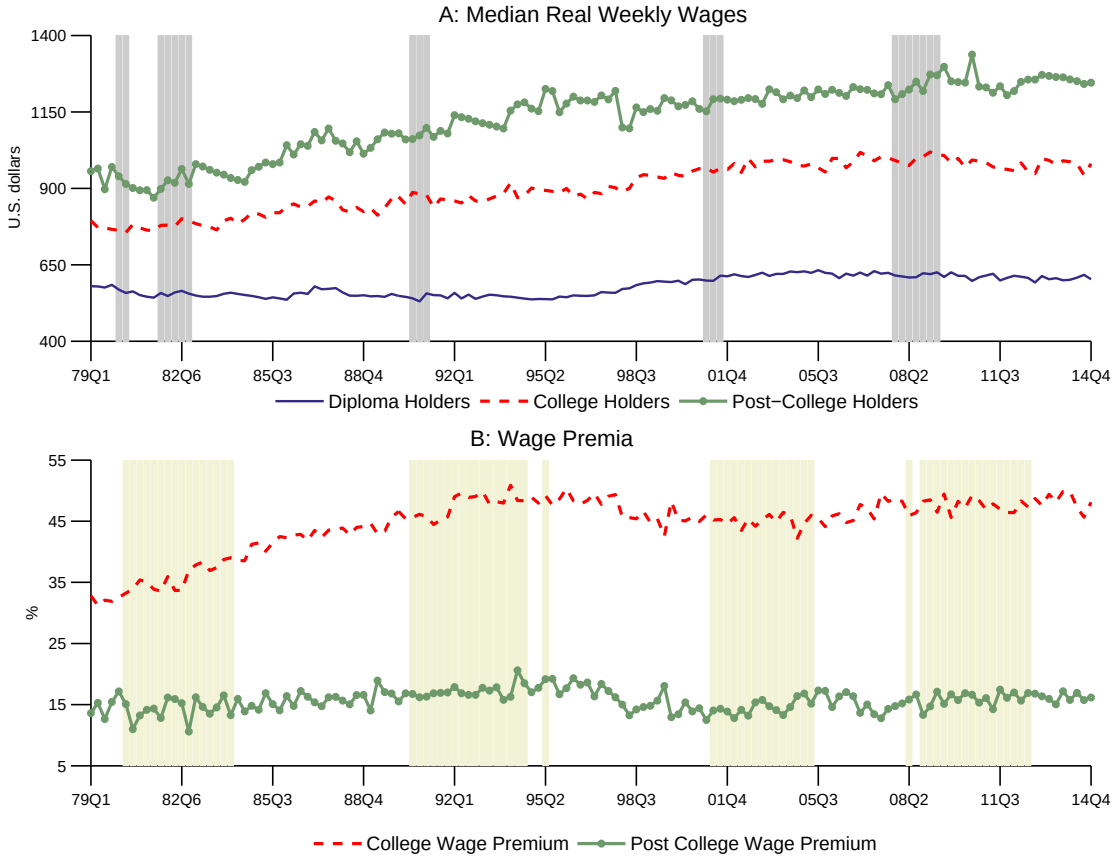
I obtain quarterly time-series of returns to human capital, $r^{e}_{t+1}$ (Equation 2), and returns to education, $r^{*e}_{t+1}$ (Equation 5), by aggregating the time-series of these four components. I discuss evidence in the next section.

[81] The full sample standard deviations of diploma, college and post college holders are 31.55, 78.37 and 113.55 U.S. dollars, respectively.

[82] Predictive coefficients are statistically significant at the 1% level (Newey-West standard errors).

Figure 5
College and Post-College Wage Premia

This figure shows the time-series of median real weekly earnings (in U.S. dollars) for diploma, college and post college holders (Panel A) and the time-series of the estimated wage premia (i.e., $w_t^e - w_t^{e-1}$, in percentage) for college and post college holders (Panel B). Wage premia, namely log differences in real wages between workers holding a college degree (post college degree) and those holding a diploma (college degree), are estimated via propensity score weighting. Shaded areas represent NBER recessions in Panel A and bad labour market times in Panels B. Bad times for the labour market are periods in which the job posting index by Barnichon (2010) is above its standardized mean. Data are from monthly samples of the CPS Outgoing Rotation Group. All series are computed at the quarterly frequency. The sample period is 1979:Q1-2014:Q4.



## 4.3 Cash-Flow Versus Unemployment-Adjusted Human Capital Measures

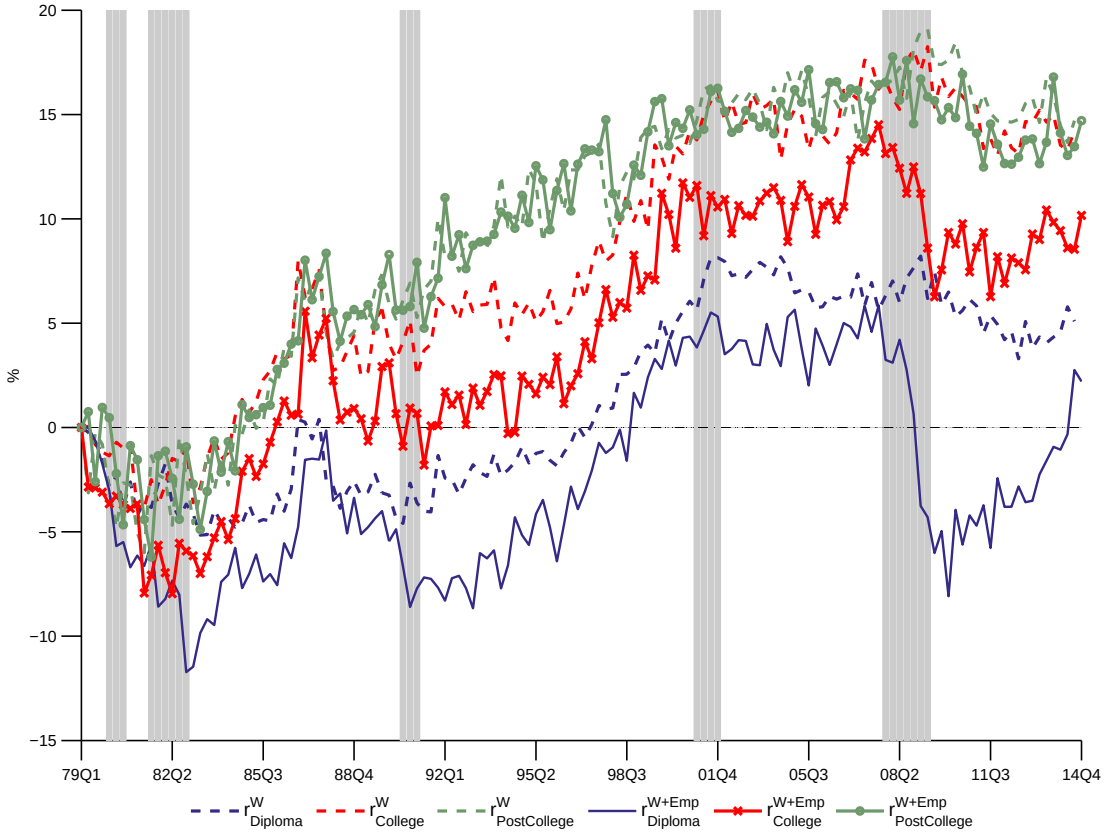
Figure 6 compares the quarterly time-series of the cumulated human capital measures for three representative investors. Measures are defined either as the growth rate of wages or by augmenting this term by the growth rate of employment shares. The former is a wage-based measure of human capital that captures fluctuations in wages of employed workers. The latter is an unemployment-adjusted measure which adjusts those fluctuations by the unemployment risk the employed workers bear over the same horizon. Since employment shares are a decreasing function of the unemployment shares, an increase in unemployment risk translates into a lower value of this measure.

Both measures are increasing in the educational attainment, but accounting for unemployment alters the inference about gains/losses human capital holders experience over time. Differences between the two measures are less pronounced the more the worker is educated and widen in bad times. The spread is tiny for highly educated workers, due to minimal historical exposure to business cycle conditions.

In bad times, highly educated workers historically experience a less severe drop in U-adjusted labour income than low educated workers. This finding points to human capital being historically riskier for low educated workers than highly educated ones. By definition, risky assets are assets that pay-off poorly in bad times for the investor. Since labour income is the pay-off of the human capital asset, human capital is risky if labour income drops in bad times. Accounting for unemployment dynamics is key to draw this conclusion: wages are sticky, while unemployment captures a deterioration of labour market conditions in real time. As revealed by the data, the unemployment component of U-adjusted labour income visibly crashes in recessions, while there is no comparable evidence for its wage component. Cash-flow measures account for negative labour market shocks only with a delay. The correlation between the two measures is negative after recessions as wages are falling while employment is rising relative to the recent past. These findings, coupled with the fact that average wages are historically higher for highly educated

Figure 6
Cumulated Human Capital Measures

This figure compares the time-series of cumulated quarterly growth in wages ($r_e^W$, dashed lines) with their unemployment-adjusted counterparts ($r_e^{W+Emp}$, solid line). Both measures are computed for three U.S. representative human capital holders $e$: A diploma holder (blue line), a college holder (red line) and a post college holder (green line). The unemployment-adjusted measure is obtained as the sum of growth in wages and growth in employment shares. Measures are in percentage. Shaded areas denote NBER recessions. Data are from monthly samples of the CPS Outgoing Rotation Group. The sample period is 1979:Q1-2014:Q4.




agents than low educated ones, have implications for human wealth. Inequality in human wealth has to be worse than the well established inequality in labour income. Taken together, these empirical facts suggest that unemployment may be the main source of the human capital risk premium and it may have an impact on marginal utility going beyond its lost-wage impact. This points to unemployment being a key driver of human capital discount rates. Unemployment is therefore a key-state variable for human capital, and potentially an omitted variables in VAR-based studies like Bansal, Kiku, Shaliastovich and Yaron (2014).

Figure 7 shows the quarterly time-series of returns to education earned by

a college (post-college) holder relative to a diploma (college) holder. Cash-flow (dashed line) and unemployment-adjusted (solid line) measures are compared as above.[83] Human capital return measures indicate that all workers are worse off in bad times relative to good time, as higher exposure to unemployment risk translates into a drop in returns. Return to education measures point out that there are relative gains in the economy at any point in time, as highly educated workers are always better off relative to low educated ones.

In the United States, education provides benefits through two channels: Higher cash-flows (wages) and lower unemployment probabilities. This implies that the traditional economic measure of human capital returns, the wage premia, underestimates the overall gains in the U.S. economy, regardless the estimation technique used to quantify it. From an economic perspective, the traditional definition of "skill premium" should also be richer in content, as benefits come from two major sources. Overall, agents should invest in education not only to maximize their expected earnings, but also to minimise their unemployment risk. The former is a cash-flow story for an educational investment, the latter is a discount rate one. When educational human capital is invested in the U.S. labour market, investing in education is like shopping for hedging instruments. The longer the time spent in education, the more expensive the instrument, but the greater the insurance provided by the labour market.

In the next section, I investigate the implications of these findings for asset prices by relating differential exposure to unemployment risk to heterogeneous financial investment behaviour.

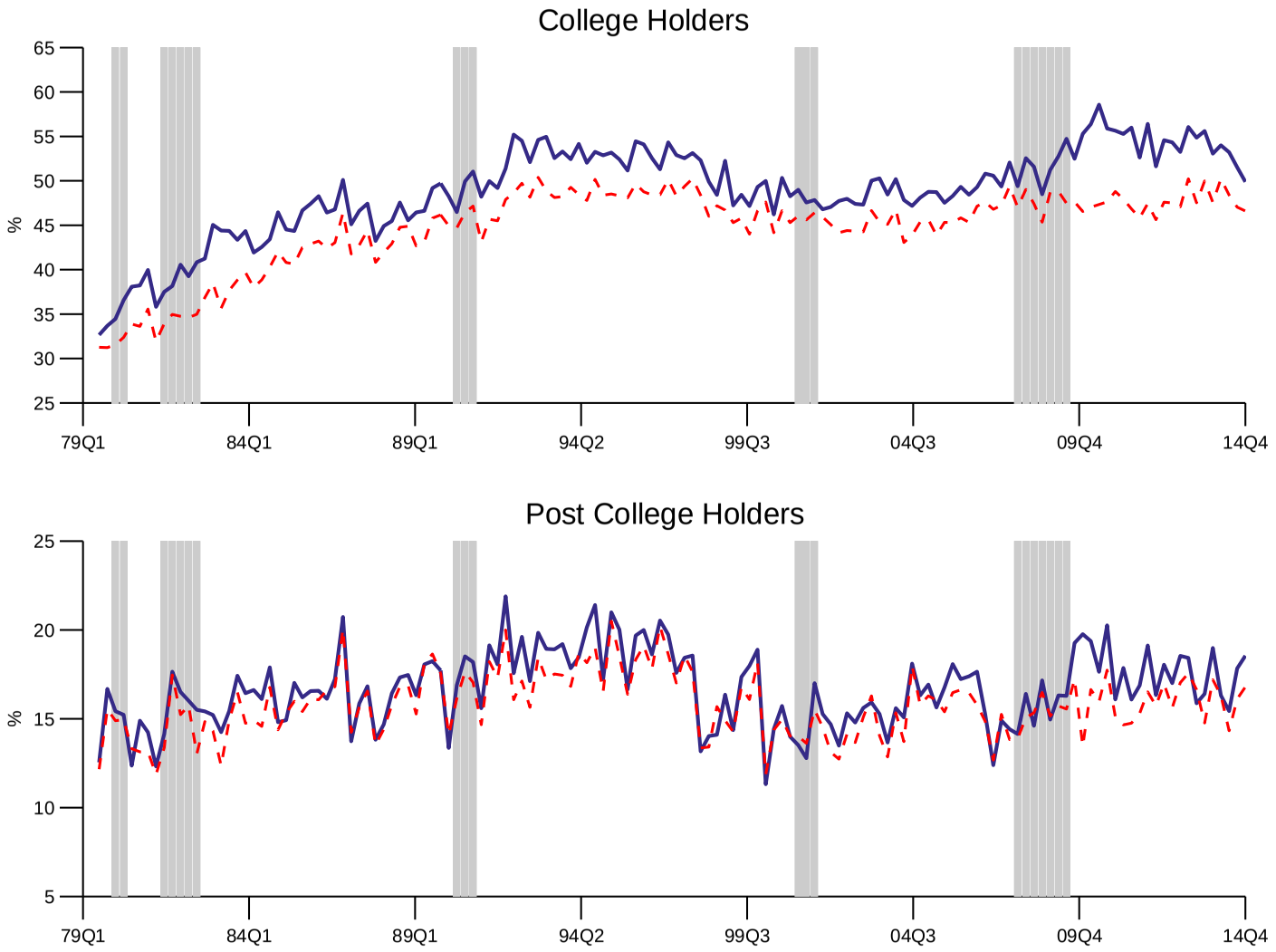
## 4.4 Unemployment Risk And Industry-Level Asset Prices

The evidence discussed in the previous section points to human capital being riskier for low educated workers than highly educated ones due to a more pronounced drop in labour income in bad times. The riskiness of human capital, at least

[83]Mean difference tests reported in the Online Appendix indicate that cash-flow and unemployment-adjusted return measures are statistically different from each other confirming that the employment premium has informational content even after adjusting for log changes in employment shares.

Figure 7
Cash-Flow Versus Unemployment-Adjusted Measures Of Returns To Education

This figure compares the time-series of cash-flow (dashed line) and unemployment-adjusted (solid line) measures of returns to education earned by a college holder relative to a diploma holder (top graph) and by a post-college holder relative to a college holder (bottom graph). The cash-flow measure is growth rate of relative wages. The unemployment-adjusted measure is sum of growth rate of relative wages and growth rate of relative employment shares. Shaded areas represent NBER recessions. Returns are quarterly and in percentage. The sample period is 1979:Q1-2014:Q4.



at aggregate level, is mainly attributable to the unemployment component of U-adjusted labour income rather than the wage component. These findings have at least two implications for financial assets.

First, if unemployment is a source of systematic risk, this risk should be priced in financial markets.

Second, differential exposure to unemployment risk points to heterogeneous financial investment behaviour. Agents who are highly exposed to changes in labour market conditions should primarily invest in financial markets to hedge their exposure to labour income risk, that is picking assets that perform well when they are more likely to lose their jobs. Agents facing mild exposure to labour market cycles, instead, should be willing to invest beyond hedging purpose, bearing the risk of a poor performance of their financial assets in bad labour market states. Using the terminology by Bodie, Merton and Samuelson (1992), human capital holders are income-constrained financial investors. Since agents cannot sell their human capital in bad times, exposure to labour income risk "constrains" their financial investment decisions. The evidence discussed above points to low (highly) educated workers being "highly" ("weakly") income-constrained financial investors and intuitively closer to the former (latter) type of financial investment behaviour.

I investigate the extent to which these two implications hold empirically.

**Employment factor, $F_t^{Emp}$.** As a first step, I define a macroeconomic factor that captures time-varying heterogeneity in the exposure to unemployment risk in the U.S. labour market. I build this factor in a standard "high-minus-low" fashion by taking the difference in the unemployment component of U-adjusted labour income growth experienced by the two extreme types of workers in the economy. At each time *t+1*, the employment factor $F_{t+1}^{Emp}$ is given by:

$$F_{t+1}^{Emp} = (\theta_{t+1}^{PC} - \theta_t^{PC}) - (\theta_{t+1}^{D} - \theta_t^{D}), \tag{7}$$

where $\theta_{t+1}^{e} - \theta_t^{e}$ denotes the growth rate in the employment probability of $e$-type worker and PC and D denote a representative post-college and diploma holder, respectively. $F_{t+1}^{Emp}$ captures relative changes in the exposure to "downside risk" heterogeneous agents experience in the U.S. labour market. An increase in $F_{t+1}^{Emp}$

can be interpreted as the low educated worker being more exposed to downside risk than the low educated worker over the quarter, due to a higher probability of being unemployed. Equivalently, the highly-educated worker being better off relative to the low educated worker over the quarter, having experienced a lower probability of being unemployed.

The time-series of $F_{t+1}^{Emp}$ (Figure A.1) is well behaved and tends to be positive at the beginning of each NBER recession and negative afterwards consistent with the evidence discussed in the previous section. A simple variance decomposition attributes 27% of its variation to the first term (highly educated worker) and 70% to the second term (low educated worker).

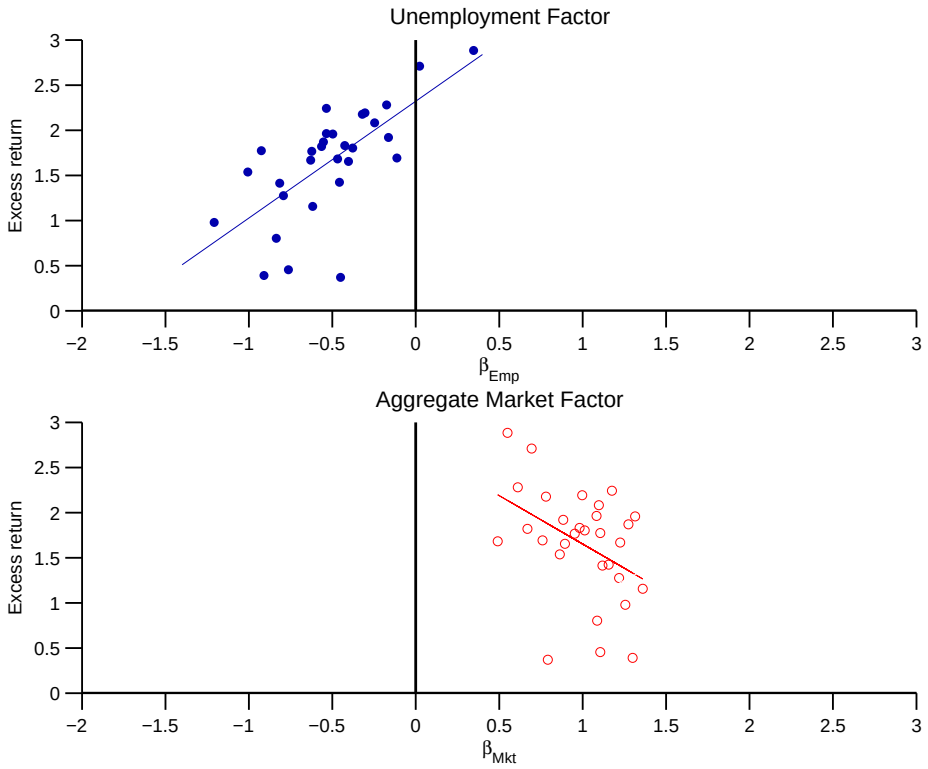
**Industry-level risk-premia.** As a second step in relating heterogeneous exposure to unemployment risk to asset prices, I use $F_{t+1}^{Emp}$ in standard asset pricing exercises. Test assets are 30 Fama-French U.S. industry portfolios, a challenging set of test assets, but also a natural place to start when macroeconomic uncertainty is the source of systematic risk to price.

As a preliminary exercise, I estimate a single factor model on each test asset using either the excess return on the aggregate U.S. stock market or $F_{t+1}^{Emp}$ as risk factor. Figure 8 reports the average excess return on each test asset (y-axis, in percentage) against its unconditional employment-beta, $\hat{\beta}^{Emp}$, in the top panel and the CAPM-beta, $\hat{\beta}^{Mkt}$, in the bottom panel. The relationship between employment-betas and average excess returns is positive and the spread in unconditional employment-betas is large ranging from about -1.3 to 0.5. The spread in unconditional market betas, instead, is bounded between 0.5 and less than 1.4. Consistent with standard asset pricing theory, industries that have low exposure to $F_{t+1}^{Emp}$ earn low risk premia on average and industries that have high exposure to $F_{t+1}^{Emp}$ earn high risk premia on average. Taken together, these two panels indicate that $F_{t+1}^{Emp}$ captures a source of risk other than pure aggregate market risk.

As a more formal exercise, I sort test assets into portfolios based on their time-varying $\hat{\beta}_t^{Emp}$-exposure, controlling for the market factor. The goal of this exercise is twofold. On the one hand, this is a test for unemployment risk being a source

Figure 8
Unemployment Risk vs Aggregate Market Risk

This figure plots industry-level average excess returns (in percentage) against their unconditional market betas in the top graph and their unconditional employment betas in the bottom graph. Risk-exposure is measured by estimating a one-factor model, that is by regressing the excess return on industry-$i$ on a constant and either the Fama-French U.S. market factor or the employment factor $F_t^{Emp}$. Test assets are 30 U.S. industry portfolios from Kenneth French's website. $F_t^{Emp}$ is growth in quarterly employment rates of post college holders minus growth in quarterly employment rates of diploma holders. (i.e., $F_{t+1}^{Emp} = (\theta_{t+1}^{PC} - \theta_t^{PC}) - (\theta_{t+1}^{D} - \theta_t^{D})$). The sample period is 1979Q1-2014Q4.




of systematic risk. Under the null, the sorted portfolios should reveal a significant risk premium. On the other hand, this is a way to investigate whether there exists a combination of industry-level traded assets that can replicate differential exposure to unemployment risk. This boils down to build a factor mimicking portfolio for the macroeconomic factor $F_{t+1}^{Emp}$.

At each time $t$ and for each test asset $i$, I estimate the following regression over a rolling window of 20 quarters (about 5 years):

$$r_t^i = \alpha_t^i + \beta_t^{Mkt,i} Mkt_t + \beta_t^{Emp,i} F_t^{Emp} + \epsilon_t^i, \tag{8}$$

where $r_t^i$ denotes the excess return on industry-$i$ between time *t-1* and time $t$, $\alpha$ is a

constant, $Mkt_t$ is the Fama-French U.S. market factor and $F_t^{Emp}$ is the employment factor. All variables are quarterly and defined over the same investment horizon.

In each window, I allocate test assets into 4 portfolios according to their estimated exposure to $F_t^{Emp}$ (henceforth $\hat{\beta}_t^{Emp}$). Low-$\hat{\beta}_t^{Emp}$ industries fall in Portfolio 1, high-$\hat{\beta}_t^{Emp}$ industries fall in Portfolio 4. In assigning industries to portfolios, I use the quartiles of the distribution of $\hat{\beta}_t^{Emp}$.

Figure 9 - Panel A plots average excess returns on these four sorted portfolios (annualized and in percentage terms) against their average time-varying employment betas. Returns do monotonically line up with their risk exposure. From Portfolio 1 to Portfolio 4, average betas vary from -1.05 to 1.22 and annualized average excess returns monotonically increase from 4.83% to 8.89%. High-$\hat{\beta}^{Emp}$ industries − industries that covary positively with $F_t^{Emp}$ − earn high risk-premia on average, and low-$\hat{\beta}^{Emp}$ industries − industries that covary negatively with $F_t^{Emp}$ − earn low risk-premia on average. Differential exposure to unemployment risk is priced at industry level. A standard long-short strategy, long in high-$\hat{\beta}^{Emp}$ industries and short in low-$\hat{\beta}^{Emp}$ industries, yields on average a statistically significant premium of about 4% over the last 35 years.

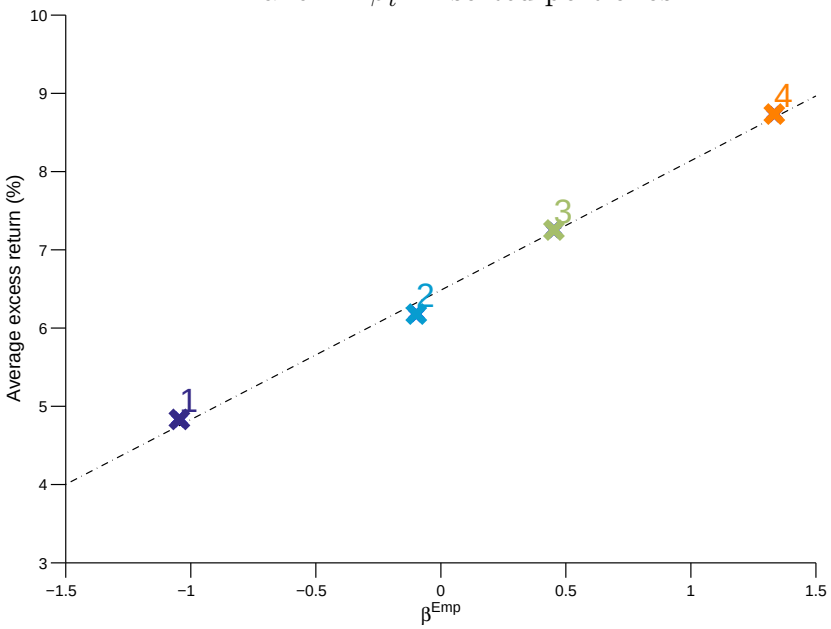
Figure 9 - Panel B shows that aggregate market risk is not priced at industry-level. When sorting test assets by their time-varying market betas $\hat{\beta}_t^{Mkt}$ estimated via Equation (8), the spread in average betas is minimal - bounded between 0.54 and 1.44 - and the resulting cross-section of average excess returns is almost flat and does not command any risk-premium. These findings corroborate the evidence in Figure 8 and confirm $F_t^{Emp}$ captures a source of systematic risk beyond aggregate market risk.

Table 5 reports risk-premia and average time-varying employment betas conditioning on subsamples. For completeness, Panel A shows descriptive statistics for the full sample period. Panel B and Panel C jointly reveal that risk-premia are driven by bad labour market states. Returns do monotonically line up with their risk exposure in both states, but the average return-spread is only 2% and is not statistically significant in the good state, while it widens up to 6.7% in the bad state. In bad times, low-$\hat{\beta}^{Emp}$ industries earn on average about 3%, while high-$\hat{\beta}^{Emp}$

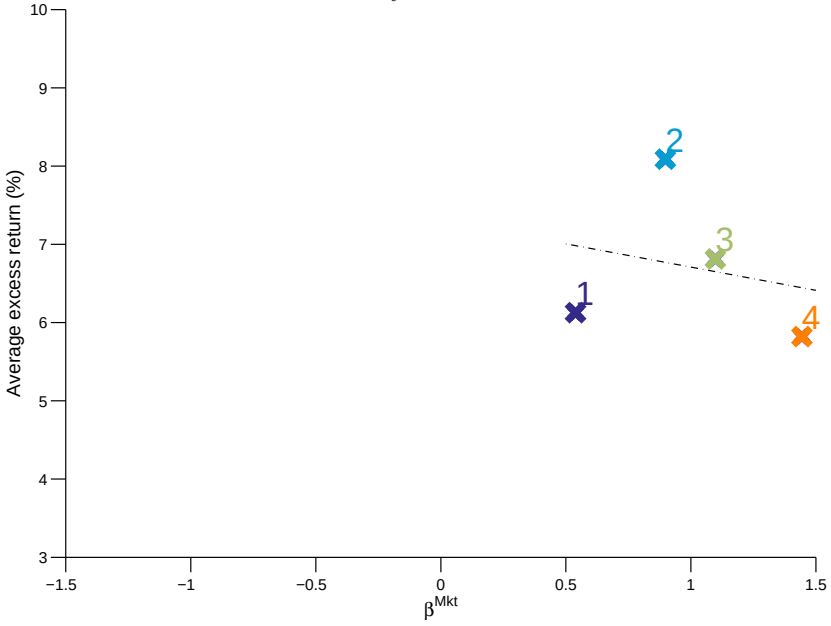
Figure 9
Industry-Level Sorted Portfolios

This figure plots the cross-section of average excess returns on industry-level sorted portfolio against their average betas. Time-varying loadings are estimated by regressing industry-$i$ excess returns on a constant, the Fama-French U.S. market factor and $F_t^{Emp}$ over rolling windows of 20 quarters. In each window, test assets are sorted into four portfolios by their time-varying exposure to either the employment factor $F_t^{Emp}$ (Panel A) or the Fama-French U.S. market factor (Panel B). Test assets are 30 U.S. industry portfolios from Kenneth French's website. $F_t^{Emp}$ is growth in quarterly employment rates of post college holders minus growth in quarterly employment rates of diploma holders. (i.e., $F_{t+1}^{Emp} = (\theta_{t+1}^{PC} - \theta_t^{PC}) - (\theta_{t+1}^{D} - \theta_t^{D})$). The sample period is 1979Q1-2014Q4.

Panel A: $\hat{\beta}_t^{Emp}$-sorted portfolios




Panel B: $\hat{\beta}_t^{Mkt}$-sorted portfolios



industries earn on average about 9.6%. As documented in Table 2, in these quarters all types of agents face a higher probability of being unemployed. Unemployment risk is thus systematic risk and this risk exposure is priced in the U.S. equity market at industry level. Panel D shows that NBER recessions, a subsample of the bad labour market state, are not the key driver. In these quarters, average excess returns on all sorted-portfolios are negative. Finally, Panel E shows that the return-spread between the top and the bottom portfolio increases up to 11% in the post-2007 period, but average returns on the four sorted portfolios do not monotonically line up. This indicates that the risk-premia are not merely compensation for the risk exposure borne by workers over the recent Great Recession. Interestingly, a simple comparison of the distribution of average $\hat{\beta}^{Emp}$ loadings across panels indicates that risk-exposure is time-varying.

**Low-$\hat{\beta}^{Emp}$ vs. high-$\hat{\beta}^{Emp}$ industries.** Is there any clear mapping between industries and sorted portfolios? Figure 10 addresses this question by reporting details about industry turnover rates, that is the percentage of quarters (out of 123) any industry falls in a given portfolio. Industries are ranked from top to bottom according to their likelihood of being low-$\hat{\beta}^{Emp}$ industries. Turnover rates are high, as any industry spans all four sorted portfolios over the sample period. This points again to pronounced time-variation in $\hat{\beta}^{Emp}$-exposure over time. Patterns do emerge, though. Coal, Telecommunication, Oil and Utilities are typical low-$\hat{\beta}^{Emp}$ industries falling about 50% of the quarters in Portfolio 1, while Smoke, Clothes and Beer are typical high-$\hat{\beta}^{Emp}$ industries falling about 50% of the quarters in Portfolio 4. Consumer Goods, Meals, Food and Construction resemble more high-$\hat{\beta}^{Emp}$ industries than low-$\hat{\beta}^{Emp}$ industries, falling more than 60% of the quarters in the two top portfolios, while the opposite conclusion holds for Financials, Wholesale and Other industries.

One potential explanation for these patterns is that time-variation in the employment-betas is driven by labour market states. To test whether industries are more likely to fall in a given portfolio in bad than good times, I run paired t-tests conditioning on any given sorted-portfolio. For each industry $i$, I first compute the

Table 5
Industry-Level Risk-Premia By States

This table reports means ($\mu$), t-statistics (in brackets) and average exposure ($\bar{\beta}_t^{Emp}$) of four portfolios of industry-level stocks sorted by their time-varying exposure to the employment factor $F_{t+1}^{Emp}$. Descriptive statistics in Panel A refer to the full sample period, results in Panels B to E are obtained conditioning on: The bad and good labour market state, NBER recessions and the post-2007 sample. The number of quarters T in each subsample is reported at the top of each panel. The last column refers to an investor that goes long Portfolio 4 (P4) and short Portfolio 1 (P1). For each test asset, time-varying loadings are estimated by regressing industry-$i$ excess returns on a constant, the Fama-French U.S. market factor and $F_{t+1}^{Emp}$ over rolling windows of 20 quarters. Test assets are 30 U.S. industry portfolios from Kenneth French's website. $F_{t+1}^{Emp}$ is growth in quarterly employment rates of post college holders minus growth in quarterly employment rates of diploma holders. (i.e., $F_{t+1}^{Emp} = (\theta_{t+1}^{PC} - \theta_t^{PC}) - (\theta_{t+1}^{D} - \theta_t^{D})$). The sample period is 1979Q1-2014Q4.

| | Portfolios | | | | |
|---|---|---|---|---|---|
| | P1 | P2 | P3 | P4 | P4-P1 |
| | A: Full sample (T=123) | | | | |
| $\mu$ | 4.83 | 6.18 | 7.25 | 8.73 | 3.90 |
| | [1.33] | [1.82] | [2.25] | [2.85] | [1.98] |
| $\bar{\beta}_t^{Emp}$ | -1.05 | -0.10 | 0.45 | 1.33 | |
| | B: Bad state (T=50) | | | | |
| $\mu$ | 2.91 | 5.79 | 6.76 | 9.57 | 6.66 |
| | [0.42] | [0.88] | [1.18] | [1.68] | [2.03] |
| $\bar{\beta}_t^{Emp}$ | -1.15 | -0.13 | 0.49 | 1.48 | |
| | C: Good state (T=73) | | | | |
| $\mu$ | 6.15 | 6.45 | 7.59 | 8.16 | 2.01 |
| | [1.58] | [1.81] | [2.01] | [2.40] | [0.83] |
| $\bar{\beta}_t^{Emp}$ | -0.97 | -0.08 | 0.42 | 1.23 | |
| | D: NBER recessions (T=14) | | | | |
| $\mu$ | -24.49 | -16.57 | -16.40 | -13.24 | 11.26 |
| | [-1.35] | [-1.02] | [-1.09] | [-0.96] | [1.31] |
| $\bar{\beta}_t^{Emp}$ | -1.52 | -0.37 | 0.23 | 1.25 | |
| | E: Post-2007 (T=32) | | | | |
| $\mu$ | -0.39 | 8.60 | 5.75 | 10.72 | 11.11 |
| | [-0.04] | [1.10] | [0.75] | [1.65] | [2.27] |
| $\bar{\beta}_t^{Emp}$ | -1.03 | -0.25 | 0.19 | 0.89 | |

probability that $i$ falls in portfolio $j$ conditioning on the economy being in either good ($p_i^{Good}$) or bad ($p_i^{Bad}$) labour market states and then test the average difference between these two probabilities across industries. Under the null, turnover rates are not driven by labour market cycles, that is $H_0 : \bar{p}_i^{BAD} - \bar{p}_i^{GOOD} = 0$.

Results are reported in Table 6. Overall, there is weak support in favour of the null hypothesis. The difference in state probabilities is not statistically significant in Portfolio 1, and it is significant, but small in magnitude in all other sorted portfolios. On average, industries have about a 60% (40%) chance to fall in portfolios other than 1 in good (bad) times, thus implying a minimal state-spread of about 20%.

**Interpreting industry-level risk-premia.** On average, industries that covary positively with the employment factor earn high risk-premia, while industries that covary negatively with $F_t^{Emp}$ earn low risk-premia. Given the nature of the risk factor underlying this novel cross-section of sorted portfolios, hiring/firing decisions taken by firms should play a major role in explaining these return differentials. The explanation I bring to the table takes the perspective of the supply side of the labour market (workers), rather than the demand side (firms). More precisely, I start from the premise that agents are income-constrained financial investors: Workers take hiring/firing decisions as given and invest in financial markets, accordingly. Under this assumption, differential exposure to unemployment risk/changes in labour market conditions should translate into heterogeneous financial investment behaviour. "Highly"-income constrained financial investors (low-educated workers) should invest in financial markets mainly for hedging purposes, while "Weakly"-income constrained financial investors (highly-educated workers) should be willing to take exposure to labour income risk and demand a premium for bearing it.

To investigate whether industry-level risk-premia reflect this behaviour, I combine two relationships, namely the employment factor $F_t^{Emp}$ and the average exposure of $\hat{\beta}_t^{Emp}$-sorted portfolios $j$, that is $\bar{\hat{\beta}}_t^{Emp,j} = Cov(r_t^j, F_t^{Emp})$. Table 7 reveals the trade-off faced by two heterogeneous agents, H (highly-educated worker) and L (low-educated worker), who own both human and financial wealth, when investing in either High- or Low-$\hat{\beta}^{Emp}$ industries. For each agent, the table discriminates

Figure 10

Low-$\hat{\beta}^{Emp}$ vs. High-$\hat{\beta}^{Emp}$ industries: Turnover Rates

This figure shows industry turnover rates for $\hat{\beta}^{Emp}$-sorted portfolios, that is the percentage of quarters (out of 123) each industry falls in any of the four $\hat{\beta}^{Emp}$-sorted portfolios. Industries are ranked from top to bottom according to their likelihood of being low-$\hat{\beta}^{Emp}$ industries (Portfolio 1). Test assets are 30 U.S. industry portfolios from Kenneth French's website. $\hat{\beta}^{Emp}$ measures exposure to the employment factor $F_t^{Emp}$, that is growth in quarterly employment rates of post college holders minus growth in quarterly employment rates of diploma holders. (i.e., $F_{t+1}^{Emp} = (\theta_{t+1}^{PC} - \theta_t^{PC}) - (\theta_{t+1}^{D} - \theta_t^{D})$). The sample period is 1979Q1-2014Q4.

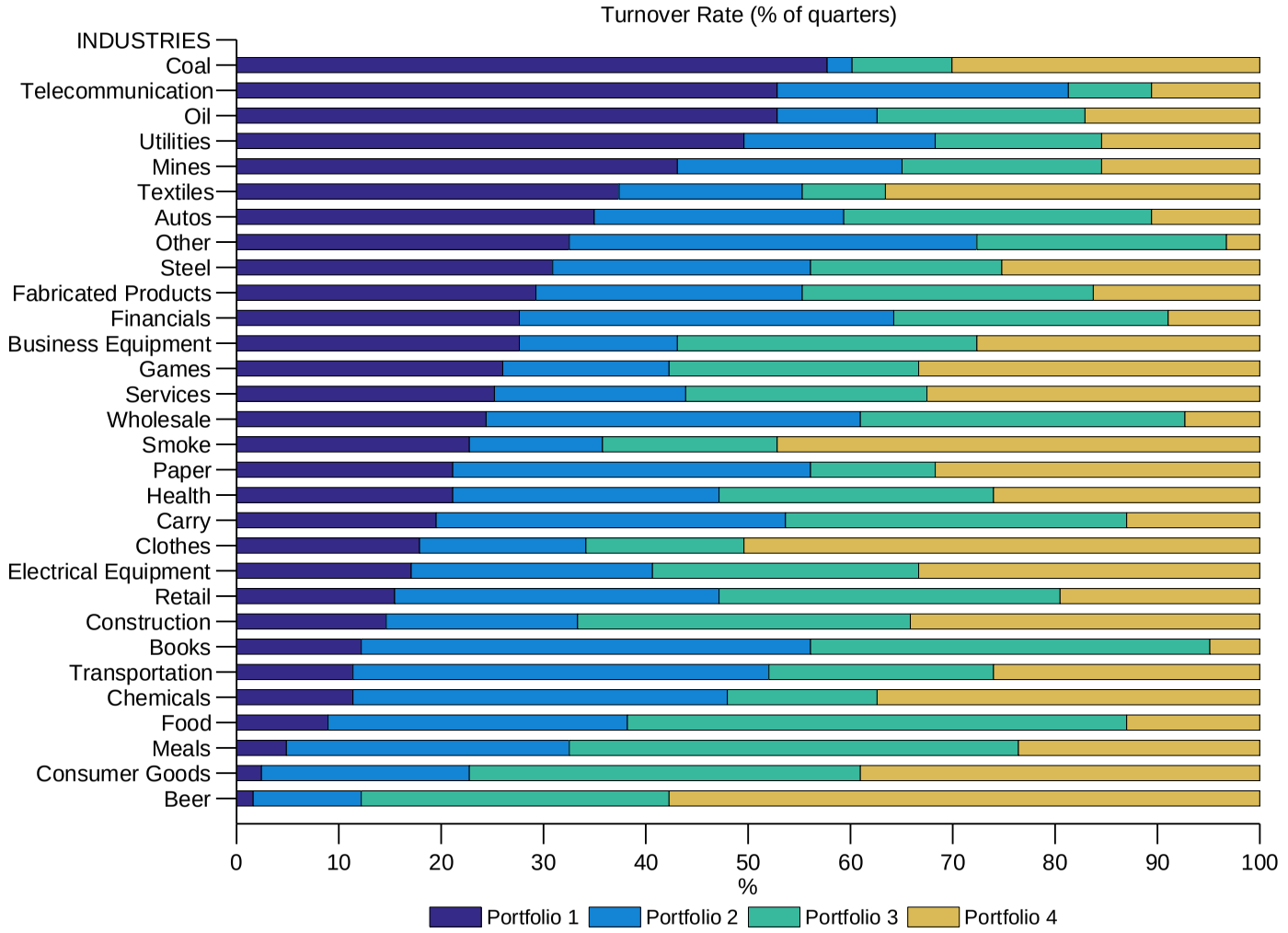


Table 6
Paired $t$-tests

This table reports paired $t$-tests under the null hypothesis that turnover rates are not driven by labour market cycles. For each industry $i$, $p_i^{GOOD}$ ($p_i^{BAD}$) denotes the probability of $i$ falling in any given $\hat{\beta}^{Emp}$-sorted portfolio conditioning on the economy being in the good (bad) labour market state. Under the null hypothesis, the average difference between these two probabilities is not statistically significant across industries, that is $H_0 : \bar{p}_i^{BAD} - \bar{p}_i^{GOOD} = 0$.

| | Port 1 | Port 2 | Port 3 | Port 4 |
|---|---|---|---|---|
| $\bar{p}_i^{BAD}$ | 0.43 | 0.40 | 0.42 | 0.40 |
| $\bar{p}_i^{GOOD}$ | 0.57 | 0.60 | 0.58 | 0.60 |
| $\bar{p}_i^{BAD} - \bar{p}_i^{GOOD}$ | -0.14 | -0.20 | -0.17 | -0.21 |
| Reject $H_0$? | No | Yes | Yes | Yes |

assets providing an hedge to labour income risk from those providing full labour income risk exposure (on average). At each point in time, hedging industries are those performing well when the agent is more likely to lose his job. Markers "+" and "-" denote good and poor asset performance, respectively. The table is filled under the assumption of an increase in $F_t^{Emp}$, that is good performance for human wealth of H (less likely to lose his job) and poor performance for human wealth of L (more likely to lose his job).

Table 7
Financial Versus Human Wealth

This table identifies $\hat{\beta}^{Emp}$-sorted portfolio of industries providing either an hedge to labour income risk or full labour income risk exposure to two representative heterogeneous workers. Worker H is highly educated (post-college degree), worker L is low educated (diploma holder). $F_{t+1}^{Emp}$ is growth in quarterly employment rates of H minus growth in quarterly employment rates of L. An increase in $F_t^{Emp}$ indicates good performance for human wealth of H (less likely to lose his job) and poor performance for human wealth of L (more likely to lose his job). $j$ denotes a $\hat{\beta}^{Emp}$-sorted portfolio of industries.

| Worker | | $j$: High-$\hat{\beta}^{Emp}$ $Cov(r_{t+1}^j, F_{t+1}^{Emp}) > 0$ | $j$: Low-$\hat{\beta}^{Emp}$ $Cov(r_{t+1}^j, F_{t+1}^{Emp}) < 0$ |
|---|---|---|---|
| H | Stocks: | + | - |
| | HC: | + | + |
| | | $j$ is risky | $j$ provides insurance |
| L | Stocks: | + | - |
| | HC: | - | - |
| | | $j$ provides insurance | $j$ is risky |

Table 7 reveals that High-$\hat{\beta}^{Emp}$ industries are hedging instruments for L, as on average they perform well ($r_t^j \uparrow$) when he is more likely to lose his job ($F_t^{Emp} \uparrow$). Low-$\hat{\beta}^{Emp}$ industries, instead, are hedging instruments for H due to a negative covariance between returns and $F_t^{Emp}$. Both investors should not require a premium to invest in these industries, being both financial instruments providing insurance to them. Standard asset pricing theory requires risky assets to command a premium (high risk premia on average) to compensate the investor for bearing the risk of a poor performance of the financial asset in bad times. Nobody should be willing to invest in risky assets, otherwise. Contrary to this reasoning, Table 7 reveals that worker L earns a premium on average when investing for hedging purposes − high-$\hat{\beta}^{Emp}$ industries outperform on average − while worker H does not.

I argue that differences in the exposure to unemployment risk can provide a rational explanation for this evidence. Agent H is typically wealthy (Mankiw and Zeldes, 1991) and historically faces a mild exposure to unemployment risk. This agent is willing to bear the risk of a poor performance of financial assets in periods in which he is more likely to lose his job, but expects to be compensated for investing in risky assets. Despite the unemployment state is historically a low probability event for this agent, the income loss he experiences in the unemployment state is severe. For this reason, H demands a premium for bearing exposure to labour income risk. Agent L, instead, is historically highly exposed to unemployment risk. This agent is a "highly"-income constrained financial investor and invests in financial markets primarily for hedging purposes. Taken together, high-$\hat{\beta}^{Emp}$ industries earn high risk premia on average driven by the demand of risky assets of agent H, while low-$\hat{\beta}^{Emp}$ industries earn low risk premia on average, being simply insurance provider for H. This reasoning is consistent with the well documented fact that stock market participation is typically higher for H-type than L-type workers (Campbell, 2006). These findings can also potentially match the argument made in Berk and Walden (2013), that "the primary reason for investing in capital markets is the attractive risk-return tradeoff offered, rather than the desire to hedge human capital risk".

**$F_t^{Emp}$ vs. HML factor.** The value premium, value stocks outperforming growth stocks on average, is a well documented fact in financial markets. Though the HML factor − the traded factor delivering this premium − is commonly used as a control variable in asset pricing tests since its discovery (Fama and French, 1992), the economic mechanism underlying the success of this trading strategy is still controversial. Since labour income hedging is a recurrent theme, I run a simple exercise to investigate the extent to which my novel $\hat{\beta}^{Emp}$-based trading strategy and a standard HML strategy relate to each other.

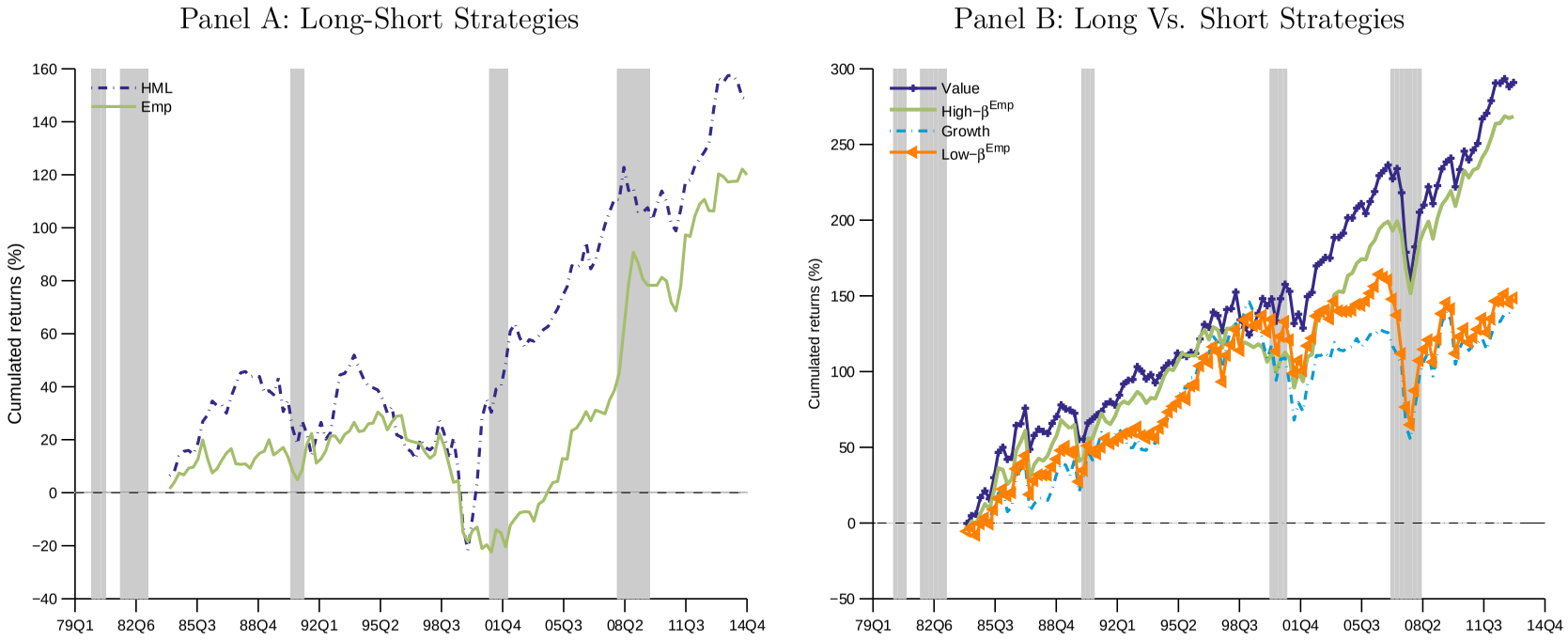
I build an industry-level HML factor by re-estimating Equation 8. At each time $t$, I sort industry-level portfolios by their exposure to the Fama-French HML factor rather than the $F_t^{Emp}$ factor. Figure 11 - Panel A compares the cumulated returns earned by an investor going long (short) industries with high (low) average exposure to the HML factor (dash-dot line) with the cumulated performance of going long (short) industries with high (low) average exposure to the employment factor $F_t^{Emp}$ (solid line). Panel B shows the cumulated performance of each portfolio of industries involved in these two trading strategies.

Panel A reveals that the value-premium historically outperforms the differential-employment premium. The two premia visibly diverge in three instances: Around 1985, 1992 and 2000. In all these cases, the HML-strategy shows an upward trend, earning positive returns, while the $F_t^{Emp}$-strategy is on a downward trend, yielding a loss to the investor. Panel B sheds lights on these patterns. Historically, value portfolios tend to outperform high-$\hat{\beta}^{Emp}$ portfolios, but low-$\hat{\beta}^{Emp}$ portfolios tend to outperform growth portfolios thus implying the value premium to be more pronounced than the differential-employment premium. The two trading strategies visibly share similar patterns as well. Except for the period 2000-2008, low-$\hat{\beta}^{Emp}$ portfolios and growth portfolios map each other pretty well, while high-$\hat{\beta}^{Emp}$ portfolios and value portfolios share similar trends but the former always outperforms the latter. Taken together, these two graphs indicate that the HML factor and the $F_t^{Emp}$ factor capture similar, but not identical, sources of systematic risk that are priced in equity markets.

Figure 11

Industry-Level Trading Strategies: $\hat{\beta}_t^{Emp}$ Vs. $\hat{\beta}_t^{HML}$ Exposure

The left graph of this figure compares the cumulated performance of going long (short) industries with high (low) average exposure to the HML factor with the cumulated performance of going long (short) industries with high (low) average exposure to the employment factor $F_t^{Emp}$. For each test asset, exposure is estimated over rolling windows of 20 quarters controlling for aggregate market risk. In each window, test assets are sorted into four portfolios, form low to high exposure to $F_t^{Emp}$. The right graph shows the cumulated performance of each portfolio of industries involved in these two trading strategies. Test assets are 30 U.S. industry portfolios from Kenneth French's website. $F_{t+1}^{Emp}$ is growth in quarterly employment rates of post college holders minus growth in quarterly employment rates of diploma holders. The sample period is 1979Q1-2014Q4.



# 5 Conclusion

This paper pins down the determinants of human wealth in the labour market starting from the simple observation that agents are human capital holder who invest their human capital in the labour market. I bring to the table a set of empirical facts suggesting that unemployment may be the main source of the human capital risk premium and may have an impact on marginal utility that goes beyond its lost-wage impact. I show that the human capital of low educated workers should be riskier than the human capital of the highly educated ones, as they experience a more severe drop in unemployment-adjusted labour income in bad times. Accounting for unemployment is key, as wages are sticky, while unemployment captures a deterioration of labour market conditions in real time. To the best of my knowledge, the role of unemployment beyond a pure cash-flow effect (wage loss) has not been explored in existent empirical asset pricing work.

My findings point to unemployment being a key driver of human capital discount rates and suggest that human capital discount rates are heterogeneous and decreasing in the endowment of human capital. These findings require theoretical foundations that I plan to provide in future work.

The evidence I provide also points to heterogeneous financial investment behaviour. Agents who are highly exposed to changes in labour marker conditions should invest in financial markets mainly for labour income hedging purposes, while agents who face mild exposure should be willing to bear the risk of a poor performance of financial assets in bad labour market states, but demand a premium for bearing it.

I document that differential exposure to unemployment risk in the U.S. labour market is priced in U.S. equity markets at industry-level. To this extent, I build a macroeconomic factor and uncover an average annualized premium of about 4% earned over the last 35 years going short industries with low exposure to my novel employment factor and long industries with high exposure to it. These risk- premia are consistent with heterogeneous financial investment behaviour and are driven by

bad labour market states, that is quarters in which unemployment risk is systematic risk, as all types of agents are more likely to lose their jobs. I provide preliminary evidence that the differential employment premium is related to the well known value-premium. I plan to investigate these patterns in detail in future work.

Overall, this set of empirical findings contributes to our understanding of human wealth and opens the door to further research on this topic.

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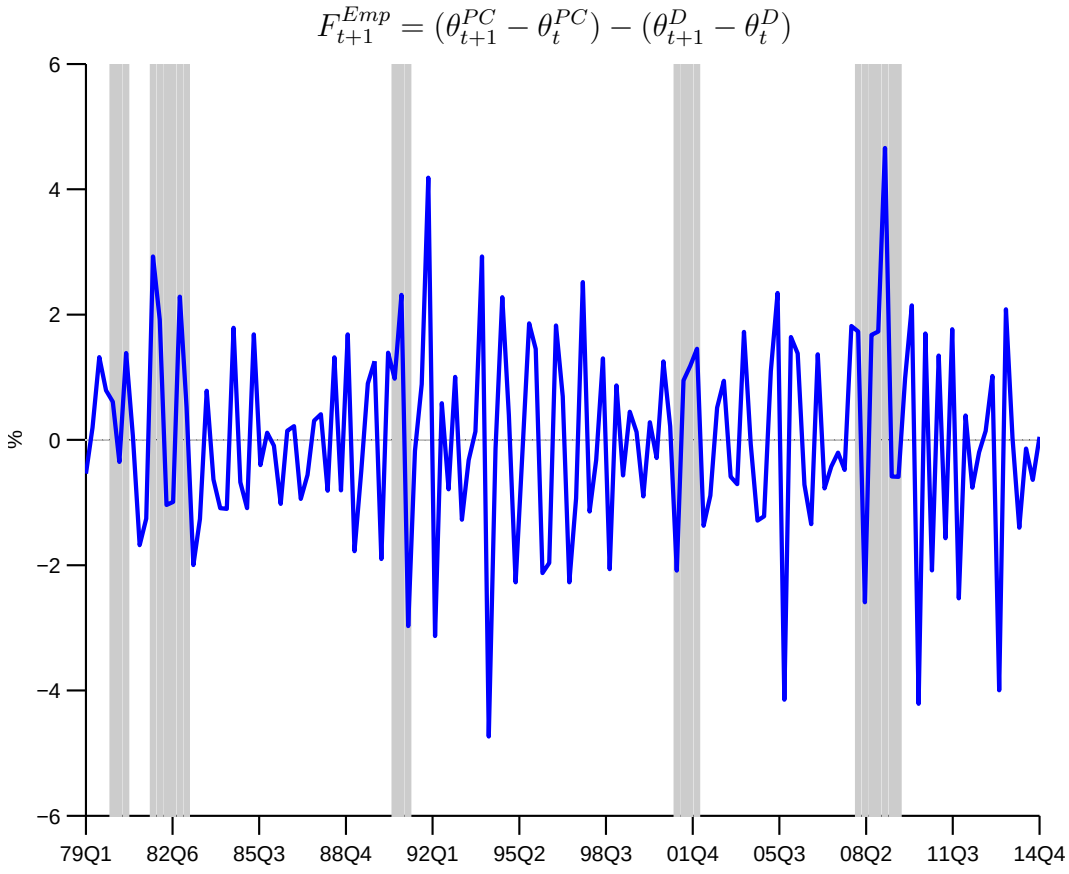
# Appendix

Figure A.1
Employment Factor

This figure shows the time-series of the employment factor $F_{t+1}^{Emp}$. $F_{t+1}^{Emp}$ is growth in quarterly employment rates of post college holders minus growth in quarterly employment rates of diploma holders. Values are in percentage. Shaded areas represent NBER recessions. The sample period is 1979:Q1-2014:Q4.



# 6 | Conclusion

In this thesis I propose three empirical asset pricing analyses on the systematic risk exposure faced by asset holders. The common idea of these studies is that agents face risks when holding human capital and financial assets, and this ultimately affects their consumption choices.

Standard asset pricing theory rationalises large differences between the average returns on risky assets and the short-term interest rate as a *premium* for bearing non-diversifiable aggregate risk. A wide empirical literature has investigated this phenomenon, aiming to identify its key risk drivers. More broadly, these studies seek to document large return differentials arising in equity as well as other asset classes (e.g., bond, foreign exchange, commodities), both within and across countries, and test the validity of a risk-based explanation in the data.

This thesis contributes to this literature by investigating three distinct sources of macroeconomic risk. The first paper, *The International CAPM Redux*, documents that investors are compensated for bearing exposure to currency risk when investing in foreign equity, either directly or via delegated portfolios of international assets. The second paper, *One Central Bank To Rule Them All*, shows that while global investors demand a premium to bear risks associated with Federal Reserve decisions, there is no comparable result for other major central banks. The third paper, *Human Capital, Unemployment Risk and Asset Prices*, argues that unemployment is a relevant state variable for the human capital asset, and shows that differential exposure to unemployment risk in the U.S. labour market is rewarded in the U.S. equity market, at least at industry level.

In the following, I discuss the general implications of each paper separately, highlighting the novelty of the analysis, its limitations and directions for future work.

The first paper - joint work with Tarun Ramadorai and Adrien Verdelhan - relies on recent advances in asset pricing research on currency markets to revisit the old-standing question of whether currency risk is priced in global equity markets. This is ultimately an empirical question, as theoretical asset pricing models already accounted for currency risk back in the 1970s.

We show that the novel empirical model we propose, the *International CAPM Redux*, successfully addresses this question capturing foreign equity risks via a set of three factors. These factors are a global equity factor denominated in local currencies, and two currency factors, dollar and carry, which are constructed from the space of foreign exchange excess returns. This specification outperforms the World and International CAPM when tested in a comprehensive set of equities from 46 developed and emerging countries spanning value, growth, small, big and aggregate market index returns from 1976 to the present. It also yields a performance comparable to the global versions of the popular Fama-French three and four factor models. Furthermore, we find evidence of substantial exposure to currency risk in a universe of global mutual funds and hedge funds, suggesting that whether exposure to international markets is direct or indirect, currency risk is a significant factor affecting the returns of international investors.

I interpret these empirical findings as a major progress relative to the existing empirical literature. As reviewed in Chapter 2, studies on global equity markets, even the most successful ones, have hardly provided robust support for the pricing of currency risk. They are implemented on a small sample of assets from developed countries, and over relatively short sample periods. With respect to delegated portfolios, our model helps shed light on the risks that foreign equity managers actually take, and potentially not disclose to customers. For example, mutual fund brochures are usually vague about currency risk, and hedge funds notoriously charge high fees for implementing strategies which could be replicated at a much lower cost.

I attribute the empirical success of our new model to two features. The first feature is the use of the currency factors, which effectively summarise a large share of common variation in currency portfolios. This helps overcome the issue of restricting

the set of currency excess returns to a small set of factors faced by previous work. Our strategy potentially accounts for all currencies in the world and also minimises the set of regressors. Moreover, being the currency factors risk factors, they capture systematic risk exposure, thereby preventing purely idiosyncratic risk – not priced in currency markets – to confound the estimation.

The second feature is the ability to disentangle the equity from the currency component in foreign equity returns. As our simple reduced form model suggests, the world aggregate equity return expressed in a common currency does contain all relevant information necessary for pricing international assets, but a one-factor model is not sufficient to capture all dynamics in the data. Our empirical specification allows to disentangle the risk drivers, which benefits the empirical tests.

We complement our empirical analysis with a reduced form model in complete markets in the tradition of Backus, Foresi, and Telmer (2001). This model is well able to replicate our key empirical findings and delivers new insights into the optimal currency hedging problem. In our framework, the optimal investment in currency assets varies over time and across investor locations, contrary to the famous result of Black (1989) who operates in a simpler mean-variance preference setting.

The aim of this model is to define a simple framework to rationalise our empirical results. This toy-model is an extension of the set up developed in previous research on the currency factors, which sounds the most natural setting to start to serve our purposes. Mapping these two currency factors in a comprehensive theoretical model, starting from the preferences of the investors, is a more ambitious goal. In my view, this is a major challenge in the field of international asset pricing nowadays, and an interesting avenue for further research.

Overall, the primary contribution of this paper is to provide an handy empirical model to price international assets. Relative to its major competitors – the global versions of the popular Fama-French three and four factor models – this model captures a well-defined source of macro-economic risk, namely currency risk, rather than pure anomalies (i.e., size, value and momentum). This is a first step to

rationalise the risk exposure of global investors, and opens the door to more theoretical work in the future.

The second paper – joint work with Pavel Savor and Mungo Wilson – contributes to the growing literature that documents high average excess returns earned around scheduled FOMC interest rate announcements. We show that this phenomenon arises in almost all stock markets across the globe. As in the United States, the FOMC effect is substantially stronger (roughly doubling) in recent years and, probably combined with an analogous currency effect, appears to be even stronger for countries that have experienced financial crises in the past two decades.

A natural question to ask is whether these international effects reflect the direct impact of FOMC announcements on foreign equity markets, or are merely a spill-over effect. A simple two-equation data generating process reveals that our estimates capture both channels in reduced form. Since it is well-known that stock-markets react simultaneously to news, the second channel should not only be at work, but it may also be the predominant one. Disentangling the two effects in reliable econometric tests is difficult, due to simultaneity bias. Using instrumental variables addresses the issue, but finding a valid instrument is difficult, especially in the context of financial markets. We leave this issue as an avenue for further research. More sophisticated econometric techniques designed to capture spill-over effects may be a suitable direction.

A primary contribution of this paper is to extend the analysis beyond the monetary policy announcements made by the Federal Reserve to other major non-U.S. central banks. We find that hardly any stock market index displays a similar effect when the Bank of England, the Bank of Japan, and the European Central Bank make interest rate announcements. The announcement premium – the average excess return difference between announcement- and non-announcement-day returns – does not even arise in the domestic stock markets of these announcing banks.

This is a puzzling finding, especially because we show that these markets do exhibit domestic inflation or unemployment announcement effects, and non-U.S. announcements of inflation and employment appear to be associated with

(marginally) higher returns even in the United States. In light of these results, we reject the possibility that economy size explains the lack of announcement premia for non-U.S. central banks. We also exclude two other candidate explanations. The first one is the possibility that non-U.S. central banks potentially matter to investors, but they choose not to pursue an active monetary policy. The second one is the possibility that these institutions actually pursue an active policy, but their decisions are widely anticipated in advance.

At this stage, we attribute our findings to the uniqueness of the Federal Reserve for global investors. Establishing why, and in what way the Fed is special is an important direction for future work. Speculative arguments are Fed capabilities, and its role in setting the global price of money. More broadly, the answer may lie in systematic differences among central banks with respect to the communication of the monetary policy, or the effectiveness in anchoring expectations.

As a final note, we interpret our findings in a risk-based framework, arguing that the announcement premia are compensation for macroeconomic risk. In related studies, Mueller, Tahbaz-Saleni and Vedolin (2015) interpret the FOMC announcement premia in currency markets as compensation for monetary policy uncertainty, and Cieslak, Morse and Vissing-Jorgensen (2015) interpret the FOMC premia in even weeks in FOMC cycle time, as risk premia for news coming from the Federal Reserve. On the contrary, Lucca and Moench (2015) interpret the pre-announcement FOMC effect as a puzzle, arguing that standard risk-based theories do not match it.

The third paper – *Human Capital, Unemployment Risk and Asset Prices* – pins down the determinants of human wealth – the present discounted value of future labour income – in the labour market. The premise is the simple observation that human capital is a risky asset, owned by workers and invested in the labour market. Human capital wealth is traditionally measured as wage income. This paper contributes to the asset pricing literature by uncovering an important role for unemployment beyond wages as a relevant state variable for human wealth.

Standard asset pricing measures allow unemployment to affect human cash-flows, but not human discount rates. The channel is the use of national labour income data, which include unemployment benefits by definition. Measures derived from life-cycle models allow unemployment to affect the human discount rate (e.g. Brown, Fang and Gomes (2012)). However, since these models concern the choices of individual households over the life-cycle, a substantial share of the risks they capture is idiosyncratic rather than systematic. In turn, these measures are not useful in asset pricing work. To the best of my knowledge, the role of unemployment beyond a pure cash-flow effect (wage loss) has not been explored in existing empirical asset pricing work.

In this paper I do not fully derive a human capital return measure. Instead, I bring to the table a set of empirical facts that suggest that unemployment may be the main source of the human capital risk premium and it may have an impact on marginal utility going beyond its lost-wage impact.

I show that, at aggregate level, the human capital of low educated workers is riskier than the human capital of the highly educated ones, as they experience a more severe drop in unemployment-adjusted labour income in bad times. Accounting for unemployment is crucial, because it captures a deterioration of labour market conditions in real time, while wages are sticky.

My findings point to unemployment being a key driver of human capital discount rates. They also suggest that human capital discount rates are heterogeneous and decreasing in the endowment of human capital. The simple unemployment-adjusted measure of labour income I propose is a first step in this direction. I am aware that providing theoretical foundations for these findings is an important venue for future work.

The evidence I provide also points to heterogeneous financial investment behaviour. Intuitively, agents who are highly exposed to changes in labour marker conditions should invest in financial markets mainly to hedge their labour income risk. On the contrary, agents who face mild exposure should be willing to bear

the risk of a poor performance of financial assets in bad labour market states, but demand a premium for bearing it.

I document that differential exposure to unemployment risk in the U.S. labour market is priced in the U.S. equity market at industry-level. To this end, I build a macroeconomic factor and uncover an average annualized premium of about 4% earned over the last 35 years going short industries with low exposure to my novel employment factor and long industries with high exposure to it. These risk-premia are driven by the demand of highly educated workers, and are more pronounced in bad labour market states. This is consistent with limited stock market participation driven by more wealthy and educated agents. These findings can potentially match the argument made in Berk and Walden (2014), that "the primary reason for investing in capital markets is the attractive risk-return tradeoff offered, rather than the desire to hedge human capital risk".

Finally, I provide preliminary evidence that the differential employment premium is related to the value-premium, and plan to investigate these patterns in detail in future work.

Taken together, these three empirical papers contribute to our understanding of the sources of systematic risk exposure faced by global and domestic asset holders. Hopefully, they will open the door to further research on these topics and in the broader asset pricing field.