An asset pricing view on international financial integration

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Preface

This thesis documents my research while I enjoyed the hospitality of the Department of Economics at the University of Dortmund and the Institute for Empirical Research in Economics at the University of Zürich as research and teaching assistant.

I owe a lot of intellectual stimulus to my supervisor Professor Mathias Hoffmann. I would like to thank him for his encouragement, support and countless helpful comments and discussions right from the beginning of my studies as doctoral student at the University of Dortmund and later as part of his team at the University of Zürich. Furthermore, I am also indebted to Professor Heinz Holländer for supervising my thesis.

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Comments by discussants and participants in various conferences, workshops and summer schools considerably helped to shape the ideas presented in this thesis. In particular, I benefited from discussions with my colleagues of the former Applied Economics (University of Dortmund) and now International Trade and Finance Team (University of Zürich), namely Torsten Ehlers, Victoria Galsband and Iryna Shcherbakova. Dr. Falko Jüssen helped me with the formal style of this thesis.

Last but not least, I would like to thank my family and friends for their support and patience during the last three years.

Introduction

The last two decades have seen a tremendous increase in research documenting the tight link between financial markets and the real economy (see e.g. the survey in Cochrane, 2006). On the one hand, macroeconomic variables help to reconcile phenomena on financial markets that have long been viewed as "anomalies". On the other hand, growing evidence supports the view that modern finance can be applied to explain puzzles in empirical macroeconomics. Both of these observations are at the heart of this thesis which focuses on the relation between financial markets and the macroeconomy in an international context to assess in how far capital markets can be described as integrated. We have witnessed tremendous increases in cross-border capital flows among industrialized economies since the mid-1980s. This finding conveys the notion of increasingly integrated capital markets in the course of what is labelled the era of financial globalisation. This notion of capital market integration is the underlying theme of this dissertation.

The framework of a cointegrated vector autoregressive representation, employed in chapter one, leaves the impression that the time series behaviour of national stock markets is heavily influenced by a global, temporary component. This common component seems to be of utmost importance in order to reconcile the variation of and comovement between developed economies' stock markets over time. This finding is suggestive of relatively well integrated stock markets of the major industrialized countries. The main results of chapter two convey the same notion. The theoretical framework that underlies chapter two is best summarized as follows: From the stance of a domestic investor, one (stochastic) discount factor should be applied to price any (foreign or domestic) asset. This statement requires that international stock markets are sufficiently integrated, otherwise a national discount factor could only be applied to national asset returns. I show in chapter two, that the use of a simple vector autoregressive representation allows to distinguish between "good" and "bad" news that drive the variation of the stochastic discount factor over time. The sensitivity of international stock returns to a national discount factor's good and bad news explains their cross-sectional dispersion from a national investor's point of view.

The third chapter shows that cross-sectional differences in exchange rate changes are well described by a conditional, consumption-based asset pricing framework, judged by either OLS cross-sectional regression results or by the comparison of pricing errors in a generalized method of moments environment. At first glance, the evidence in chapter three, again derived from a national investor's point of view, seems to underscore the impression of market integration. From the point of view of a domestic investor, cross-sectional differences in exchange rate changes are explained by a consumption-based asset pricing model conditional on instrumental variables. However, the variable that contains the conditioning information about exchange rate changes is theoretically motivated by an incomplete markets model. Market incompleteness thus seems to be key in explaining the cross-section of exchange rate changes. Chapter four uses this idea to derive a heterogeneous agent asset pricing model which allows to show that the cross-section of exchange rate changes is not driven by global but country-specific (idiosyncratic) consumption risk.

To summarize, chapter one and two provide evidence that can be interpreted in favour of relatively well integrated stock markets among the major industrialized economies, whereas chapter three and four rather convey the notion that foreign exchange markets are incomplete since idiosyncratic risks seem to be key in capturing the cross-section of exchange rate movements.

As each of the four chapters represents a self-contained paper, I use the remainder of this introduction to highlight the broader economic context in which each of the chapters should be viewed and briefly highlight the main results apart from the common topic of capital market integration.

Chapter one focuses on two of the most salient phenomena observed on international stock markets: i) the predictability of stock market returns and ii) the comovement of national stock markets. This chapter establishes a link between these two observations which has not been considered so far.

Starting with Campbell and Shiller (1988a), Fama and French (1988) and Poterba and Summers (1988), a vast amount of empirical studies highlight that returns on stock markets are predictable in the long-run. Alas, traditional financial market theory implies that expected stock market returns should be constant and hence unpredictable if stock markets process information efficiently. Stock prices thus follow a random walk. Market efficiency (in an informational sense) and stock return predictability seemed to contradict each other. This is true as long as investors' risk aversion does not vary over time. However, Constantinides (1990), Constantinides and Duffie (1996) and Cochrane and Campbell (1999) show that time-variation in stock market returns reflects time-varying risk premia on stocks as the risk aversion of stock holders varies over the business cycle. Constantinides and Duffie (1996) argue that uninsurable background risks drive the risk aversion of stock market participants. Constantinides and Duffie take into account that investors have a job and that labour income is an important source of investors' wealth (Heaton and Lucas (2000a,b) and Hoffmann (2006a) emphasize the important role of proprietary income in this context). It is more likely to loose one's job in a recession than during a boom. Furthermore, it is virtually impossible to insure against this risk. Since stocks are assets that tend to pay off well in a boom but perform poorly in recessions, the investor's attitude towards holding these risky assets heavily depends on the underlying risk of potential job loss and hence on the state of the business cycle. Constantinides (1990) and Campbell and Cochrane (1999) motivate time-varying risk aversion over the business cycle with the formation of consumption habits. When current consumption approaches the habit in a recession, then investors appreciate assets that pay off well in this state of the economy and dislike assets that pay off poorly. That is why the risk aversion of investors/consumers and hence risk premia on stocks should vary over the business cycle. Hence, predictability of stock returns at the business cycle frequency does not exclude market efficiency.

So, macroeconomic variables should explain stock market movements. But financial variables, such as the dividend-price or the price-earnings ratio, have been the only successful predictors of stock market returns until the seminal findings in Lettau and Ludvigson (2001a, 2004). Based on the idea that transitory changes in wealth leave consumption unaffected, Lettau and Ludvigson (2001a, 2004) show that temporary fluctuations of the U.S. consumption-wealth ratio predict U.S. stock market returns at the business cycle frequency. Chapter one discusses a refinement of their framework in order to show formally that short-run fluctuations of the consumption-wealth ratio should not only predict risk premia on the U.S. but also on foreign stock markets. Indeed, the U.S. consumption-wealth ratio explains time-variation in foreign stock market returns. This finding leaves the impression of a common, temporary component in national stock markets. A common component should explain the comovement between national stock markets as stressed by Campbell and Hamao (1992). The proportions of covariation among the G7 stock markets explained by the U.S. consumption-wealth ratio at three to five year horizon are substantial. Furthermore, the level of comovement explained by the common stationary stock market component rises to 70 percent in the sample from 1990 to 2005 which coincides with the observation of a tremendous increase in cross-border equity flows since the mid-1980s (Tesar and Werner, 1995; Hau and Rey, 2004, 2006). Taken together these findings suggest increased stock market integration among the major industrialized countries as the common, transitory component has gained considerable importance for the time series variation of and covariation between national stock markets in the past two decades.

Chapter two serves two purposes. First, it deals with the stylised fact that stocks with high ratios of book value relative to market value promise higher average returns than predicted by the by now standard capital asset pricing model (CAPM) proposed by Sharpe (1964) and Lintner (1965). Conversely, low book-to-market ratio stocks offer lower returns than suggested by the CAPM as Fama and French (1992, 1993) reveal for the U.S. and Chan (1991), Capaul et al. (1995) and Fama and French (1998) in international data.

The Sharpe-Lintner CAPM assumes that all investors hold a fraction of a so called market portfolio, which comprises all risky assets, and are invested in a risk-free asset according to their level of risk aversion. Average returns on stocks should be determined by their sensitivity (covariance) to the return on the market portfolio, the market beta. Value stocks typically have equal or even lower market betas than growth stocks but substantially higher average returns which conveys the notion of an additional premium on those stocks.

The CAPM does not capture the cross-section of value and growth stock returns because the spread in market betas cannot explain differences in average returns. However, Campbell and Vuolteenaho (2004) argue that the market beta hides more than it reveals. The market return is driven by news about the market's future cashflows, i.e. fundamentals, and news about expected discount rates. A change in fundamentals is permanent by nature, while high expected discount rates lower the present value of cashflows. However, high expected returns today signal better future investment opportunities. Therefore, receptiveness to the market return's cashflow news component should be compensated with a higher risk premium than sensitivity to the discount rate news component.

Campbell and Vuolteenaho (2004) focus on the U.S. and decompose the CAPM market beta into a cashflow and a discount rate variety. They show that the market beta of value stocks is dominated by the cashflow component whereas the market betas of growth stocks are predominantly discount rate betas. The spread in cashflow betas explains the cross-section of value and growth stock returns and hence the empirical failure of the CAPM in U.S. data.

I exploit these findings to show that from the point of view of an arbitrary national investor the basic reasoning by Campbell and Vuolteenao (2004) pertains to European value and growth stocks as well. European value stocks promise relatively high returns which coincides with the fact that their betas with respect to national market returns are dominated by the cashflow news variety.

Secondly, chapter two assesses in how far one implication of basic asset pricing theory can be reconciled with the data. If markets are sufficiently integrated, then one national asset pricing model should capture returns on national and international assets. However, a wealth of studies documents that even capturing the cross-section of national stock market returns with national asset pricing models constitutes a challenge. Empirical evidence presented in this thesis suggests that two-beta variants of national CAPMs explain the cross-sectional dispersion in European stock returns thus underscoring the impression of integrated stock markets among European countries.

C**hapters three** and **four** build upon the solution to the uncovered interest rate parity (UIP) puzzle recently put forward by Lustig and Verdelhan (2006). According to UIP, differences in the interest rates of foreign and domestic risk-free assets can only be caused by expected exchange rate changes. UIP predicts a foreign currency depreciation if the foreign interest rate is higher than the domestic one. However, we often observe the opposite pattern in the data, except for high inflation countries (Bansal and Dahlquist, 2000). This evidence suggests the presence of a premium for investing into foreign securities. Hence, asset pricing theory should explain deviations from UIP.

Lustig and Verdelhan (2006) are the first to show empirically that the lack of empirical support for UIP is the immediate implication of basic asset pricing theory. Lustig and Verdelhan provide evidence that currencies from countries with high interest rates relative to the U.S. are perceived as risky investment and hence have to promise high (positive) returns to compensate the investor for this riskiness. A high return on a currency corresponds to a foreign currency appreciation, which hence explains the empirical failure of UIP.

Key to success in Lustig and Verdelhan (2006) is the formation of currency portfolios with respect to interest rate differentials vis-à-vis the U.S. The building of portfolios, which are rebalanced annually, allows to incorporate the largest possible cross-section of countries and guarantees a large spread in interest rate differentials in each period. This spread in interest rate differentials creates substantial differences in average excess returns on the foreign currency portfolios. Excess returns are defined as exchange rate change less the respective interest rate differential vis-à-vis the U.S., i.e. the departure from UIP. Lustig and Verdelhan show that consumption-based asset pricing models are successful in explaining the cross-section of average returns on these portfolios and the failure of UIP as mentioned above.

In **chapter three**, I combine these findings with recent studies that emphasize a tight link between exchange rate changes and stock market return differentials among developed economies to explain cross-sectional differences in U.S. dollar exchange rate changes. Chapter three shows that quarterly returns on currencies (exchange rate changes) of developed countries are priced by the consumption-based capital asset pricing model (CCAPM) if lagged stock market return differentials are used as conditioning instrument. Motivated by Lustig and Verdelhan (2006), I start with forming currency portfolios with respect to interest rate and stock return differentials. The CCAPM explains about 50 percent of the cross-sectional variation in these currency portfolio returns. The estimated price of consumption growth risk is around two percentage points p.a. This chapter shows that the use of lagged stock return differentials derives immediately from the consumption-based asset pricing framework. Currencies from countries with past low stock market returns relative to the U.S. tend to appreciate against the U.S. dollar because they are perceived as relatively risky asset. Hence those currencies have to offer high returns.

Additionally, the CCAPM explains returns on foreign currencies on a country-by-country basis when they are scaled with the respective stock return differential. The explanatory power of the CCAPM for scaled currency returns is about as strong as for returns on currency portfolios.

Chapter four, which is joint work with Professor Mathias Hoffmann, deals with the mechanism that explains the success of Lustig and Verdelhan (2006). At the heart of the mechanism is the insight that investing into foreign currency amounts to a bet on the foreign intertemporal marginal rate of substitution: countries with relatively high consumption growth will tend to have depreciating currencies, those with low consumption growth appreciating ones.

According to the CCAPM, currencies -- like any asset -- should be priced for the exposure to world aggregate consumption growth risk they deliver. In terms of international consumption comovements, this risk has two determinants: First, given the relative volatility of consumption growth rates, the currencies of countries with high consumption correlations offer a better hedge since these currencies will depreciate less (or may even appreciate) when average world consumption growth is low. Secondly, given consumption correlations, the relative volatility of consumption determines the size of the risk premium; if consumption growth rates are positively correlated, countries with more volatile consumption growth rates offer a good hedge against world consumption growth risk since they tend to appreciate when world consumption growth is low. These currencies therefore generate low excess returns. Conversely, currencies of countries whose consumption growth is negatively related to world consumption but more volatile will offer a particularly bad hedge and will therefore show positive expected excess returns.

The logic of this mechanism implies that the international dispersion of consumption growth rates, as a measure of idiosyncratic consumption risk, should be a key determinant of the success in Lustig and Verdelhan (2006). However, they employ a representative agent model enhanced with the information about exchange rate changes inherited in interest rate differentials. We argue, and provide evidence for, that conditioning on interest rate differentials is very similar to exploiting the cross-sectional dispersion in consumption growth rates. Thus it seems natural to use a version of the CCAPM that incorporates international consumption heterogeneity to explain cross-sectional differences in currency returns. Our CCAPM accounts for consumption heterogeneity by including the cross-sectional dispersion in consumption growth rates as pricing factor in the CCAPM. This variant of the CCAPM explains more than 60 percent of the cross-sectional variation in currency returns. Empirical evidence suggests that it is the relative volatility mechanism proposed by Lustig and Verdelhan which explains the success of our international CCAPM with consumption heterogeneity.

We use this framework to dissect the consumption-exchange rate anomaly. In a world with complete financial and goods markets and homothetic preferences, consumption growth rates should equalize. Backus and Smith (1993) show that deviations from purchasing power parity drive a wedge between consumption growth rates. At the same time, real exchange rates should be determined by relative consumption growth rates. However, empirical support for this condition is virtually absent. The correlations of relative consumption growth rates with real exchange rates are extremely low. We show that a somewhat weaker version of the Backus-Smith relation only holds when UIP holds. UIP is grossly violated in the data, but departures from UIP are captured by our model. According to Backus and Smith (1993) cross-sectional differences in price level changes (inflation) should explain cross-sectional differences in (real) exchange rate changes. We therefore decompose the cross-sectional variation in consumption into one component that is due to cross-sectional differences in inflation expectations and a second component that is due to international variation in nominal interest rates (the key to success in the Lustig and Verdelhan (2006) framework). Both factors account to virtually equal parts for the pricing power of the cross-sectional variance of consumption. We thus provide (indirect) evidence for the Backus-Smith condition. Our results also suggest that the failure of uncovered interest parity and the consumption real exchange rate anomaly are more closely linked than may have commonly been believed.

Chapter 1

International evidence for return predictability and the implications for long-term covariation of the G7 stock markets

1.1 Introduction

If stock markets are perfectly integrated, then they should be driven by the same factors. Harvey (1991), Campbell and Hamao (1992) and Ferson and Harvey (1993) document the importance of global risk factors for the predictability of national stock market returns as well as explaining their cross-sectional differences. In particular, Campbell and Hamao (1992) show that U.S. financial variables, such as the dividend-price ratio (Campbell and Shiller, 1988a; Fama and French, 1988) or the relative treasury-bill rate (Campbell, 1991; Hodrick, 1992), do not only predict stock returns on the U.S. but also on the Japanese stock market. Building upon the seminal contributions of Lettau and Ludvigson (2001a,2004), Guo (2006) shows that short-run fluctuations of the U.S. consumption-wealth ratio predict quarterly excess returns on U.S. as well as risk premia on foreign stock markets. These findings leave the impression of a common, temporary component in national stock markets.

Based on these findings, the aim of this chapter is twofold. First, I show formally that the U.S. consumption-wealth ratio has to forecast returns on foreign stock markets if the basic logic applies that motivates its use as a predictive variable for U.S. stock returns, thus complementing the recent evidence by Guo (2006) who stops short of providing a formal rationale for his findings. In addition, this chapter shows that the predictive power of the U.S. consumption-wealth ratio is most pronounced at the business cycle frequency which is in line with theoretical macroeconomic models that rationalize predictability of stock returns with time-varying risk aversion over the business cycle.[1] The finding of international stock return predictability by short-run fluctuations of the U.S. consumption-wealth ratio, henceforth abbreviated with *cay*, stresses the impression of a common, temporary component in national stock markets. However, the implications of a global, stationary stock market component for the comovement of stock markets have not been considered so far. It is the second main contribution of this chapter to fill this gap.

The existing literature on international stock market comovement has taken two extreme points of view as of yet. On the one hand, a wealth of studies focuses on short-term correlations documenting a considerable increase in the short-term (monthly) correlation of global stock market returns (e.g. Longin and Solnik, 1995; Brooks and Del Negro, 2004, 2006; Goetzmann et al., 2005; Berben and Jansen, 2005). On the other hand, the idea that covariation is the reflection of a common permanent component in national stock markets has attracted a lot of attention since the dispute between Kasa (1992) and Richards (1995). Kasa (1992) provides evidence for one common stochastic trend driving the five major stock markets in the long-run using the Johansen cointegration test. Richards (1995) questions the statistical basis of Kasa´s analysis and tests empirically some of the restrictions of cointegration on the time-series behaviour of national stock markets. Cointegration among national stock markets requires the predictability of relative stock returns, since a deviation from the common trend by one stock market should be exactly offset by another. For the same reason, cointegrated stock markets imply virtually no long-run gains from international diversification. Richards (1995) provides some evidence for relative stock return predictability. Hence, we cannot rule out the existence of a common, permanent stock market

[1] Time variation in risk aversion can be induced by uninsurable background risks (Constantinides and Duffie, 1996; Heaton and Lucas, 2000a,b), limited stock market participation (Polkovnichenko, 2004; Vissing-Joergensen, 2002) or the formation of consumption habits (Constantinides, 1990; Campbell and Cochrane, 1999).

component. In addition, Goetzmann et al. (2005) highlight that -- judged by short-term correlations -- the scope of international diversification among the stock markets of the major economies is very limited.

This chapter takes an intermediate position in terms of the frequency (monthly versus several years) at which the comovement of stock market returns is regarded compared to the existing literature on this issue. In this chapter, I argue that the predictive power of U.S. *cay* for foreign stock market returns reflects the substantial importance of a common, temporary component in explaining the evolution of national stock markets at the business cycle frequency. The forecast ability of *cay* peaks at three-year horizon. I assess how much of the covariation of the G7 stock markets at three-year horizon can be reconciled with the exposure to *cay*, the mirror image of the common, transitory stock market component.

Sensitivity to *cay* explains between 15 to 60 percent of the comovement of three-year returns on the G7 markets over the sample period from the fourth quarter of 1969 to the first quarter of 2005. A visual comparison of the realised three-year excess returns with the corresponding fitted values from long-horizon regressions reveals that the fit is best in the sample period from 1990 to 2005. For that time period comovement between the G7 stock markets implied by their exposure to the common stationary component is even more pronounced. This finding coincides with the observation of a tremendous rise in cross-border capital flows in the past two decades (e.g. Tesar and Werner, 1995; Hau and Rey, 2004, 2006; Lane and Milesi-Ferretti, 2001) and with firm level evidence of the increasing importance of global shocks for short-term comovement of stock returns in the course of the last twenty years (Brooks and Del Negro, 2006). Thus a higher degree of integration among international stock markets seems to have raised the importance of the common, stationary component for the time-series variation in and comovement between the G7 stock markets.

In the next section, I sketch my slight refinement of the framework of Lettau and Ludvigson (2001a, 2004) to show formally why the U.S. consumption-wealth ratio should be informative about the future path of foreign stock markets. Section 1.3 assesses the cointegration properties of the resulting four variables comprising decomposition of the U.S. consumption-wealth ratio in detail. The predictive power of U.S. *cay* for excess returns on the MSCI indexes of the G7 is assessed in section 1.4. Section 1.5 discusses the role of *cay* in explaining stock market comovement. Finally, I conclude in section 1.6. The appendix of this thesis contains details about the construction and sources of data used in this chapter.

1.2 The relation between U.S. consumption, wealth and expected returns on foreign stock markets

The main results of Lettau and Ludvigson (2001a, 2004) convey the notion that time variation in the U.S. consumption-wealth ratio, i.e. time variation in *cay*, finds its main source in cyclical changes in the market value of U.S. households´ stock market wealth. These market value changes are induced by expected returns on stock market wealth mirrored in broad stock indexes. Hence, *cay* predicts stock returns. I use the remainder of this section to show formally that this reasoning pertains to U.S. households' foreign stock holdings as well, such that U.S. *cay* has to forecast returns on foreign stock markets. Therefore, I tautologically rewrite the trivariate consumption-wealth ratio approximation of Lettau and Ludvigson (2001a, 2004) to explicitly take account of U.S. households' foreign stock holdings.

Following Campbell and Mankiw (1989), Lettau and Ludvigson (2001a) regard a representative agent economy in which all wealth is traded.

The representative household faces an intertemporal budget constraint of the form:

$$W_{t+1} = (1 + R_{w,t+1})(W_t - C_t) \tag{1.1}$$

where W_t denotes aggregate wealth (human wealth plus asset wealth) in period t. C_t denotes consumption and $R_{w,t+1}$ the net return on aggregate wealth.

Rearranging the budget constraint for the ratio of consumption to wealth and taking a loglinear approximation around the mean consumption-wealth ratio under the assumption that this mean is covariance stationary leads to the following law of motion for the log consumption-wealth ratio.

$$c_t - w_t = E_t \sum_{i=1}^{\infty} \rho_w^i (r_{t+i}^w - \Delta c_{t+i}) \tag{1.2}$$

Lower-case letters denote natural logarithms throughout the chapter, Δ denotes the difference operator.

In order to make equation (1.2) empirically tractable, Lettau and Ludvigson (2001a, 2004) decompose aggregate wealth into its components asset and human wealth and loglinearise around the long-run mean of the ratio of human and asset wealth which leads to

$$w_t \approx va_t + (1-v)h_t \tag{1.3}$$

with v interpretable as average share of asset wealth in aggregate wealth, a_t, log asset wealth and h_t, log human wealth such that the log consumption-wealth obeys

$$c_t - va_t - (1-v)h_t = E_t \sum_{i=1}^{\infty} \rho_w^i (vr_{t+i}^a + (1-v)r_{t+i}^h - \Delta c_{t+i}) \tag{1.4}$$

decomposing the return on aggregate wealth, r_t^w accordingly (Campbell, 1996).

I further decompose asset wealth into foreign stocks held by U.S. households, abbreviated with fs_t, and the rest of asset wealth (U.S. stocks, real estate, etc.) which I will refer to as domestic asset wealth, daw_t. A loglinear approximation of asset wealth around the foreign equity to domestic asset wealth ratio yields

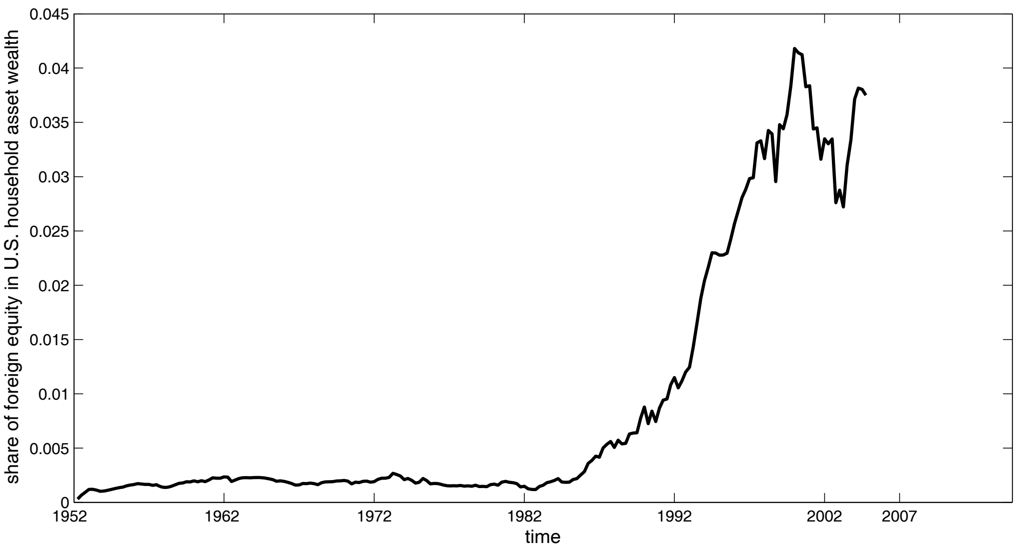
$$a_t \approx \lambda_t daw_t + (1-\lambda_t) fs_t \tag{1.5}$$

with λ_t the (time-varying) share of domestic asset wealth in U.S. households´ asset wealth and $1-\lambda_t$ the share of foreign equity.[2] As the shares of domestic asset wealth and foreign

[2] I thank Mathias Hoffmann for suggesting this course of analysis.

equity are observable, I do not require to assume constant foreign equity and domestic asset wealth shares which would be more than inappropriate as figure 1.1 shows. Figure 1.1 presents the share of foreign equity in U.S. households' asset wealth (vertical axis) over time from 1952 to 2005 (horizontal axis). Apparently, the share of foreign equity in U.S. asset wealth is small in absolute terms but experienced a substantial increase since the mid-1980s. Since λ_t could also be interpreted as a measure of home bias in U.S. asset wealth, this figure visualizes the decline in home bias as documented in Lane and Milesi-Ferretti (2001) for a large set of countries.

Figure 1.1: Share of foreign equity in U.S. households' asset wealth



Combining (1.4) with (1.5) gives

$$w_t \approx v(\lambda_t daw_t + (1-\lambda_t) fs_t) + (1-v)h_t \tag{1.6}$$

A loglinear approximation of the gross return on asset wealth, r_t^a, with respect to foreign equity and domestic asset wealth follows the same pattern as the decomposition of the asset wealth components (Campbell, 1996), such that we obtain $r_t^a = \lambda_t r_t^{daw} + (1-\lambda_t) r_t^{fs}$ and hence

$$c_t - v(\lambda_t daw_t + (1-\lambda_t) fs_t) + (1-v)h_t$$
$$= E_t\{\sum_{i=1}^{\infty} \rho_w^i [(v(\lambda_t r_{t+i}^{daw} + (1-\lambda_t) r_{t+i}^{fs}) + (1-v) r_{t+i}^{h}) - \Delta c_{t+i}]\} \tag{1.7}$$

by plugging (1.6) into (1.2).

Multiplying out gives

$$c_t - vfs_t - v(\lambda_t (daw_t - fs_t)) + (1-v)h_t$$
$$= E_t\{\sum_{i=1}^{\infty} \rho_w^i [(vr_{t+i}^{fs} - v(\lambda_t (r_{t+i}^{daw} - r_{t+i}^{fs})) + (1-v) r_{h,t+i}) - \Delta c_{t+i}]\} \tag{1.8}$$

However, (1.8) cannot be employed for empirical purposes because one part of aggregate wealth, human wealth, is unobservable. I follow Lettau and Ludvigson (2001a) and assume labour income to represent the dividend paid from human wealth and thus its non-stationary component to overcome this obstacle. Then log human wealth, h_t, obeys

$$h_t = \kappa + y_t + z_t \tag{1.9}$$

with y_t, log labour income, κ, a constant term and a covariance stationary term z_t. Plugging (1.9) into (1.8) and assuming that the return on labour income equals the return on human wealth gives

$$c_t - vfs_t - v(\lambda_t (daw_t - fs_t)) + (1-v)y_t$$
$$= E_t\{\sum_{i=1}^{\infty} \rho_w^i [(vr_{t+i}^{fs} - v(\lambda_t (r_{t+i}^{daw} - r_{t+i}^{fs})) + (1-v)\Delta y_{t+i}) - \Delta c_{t+i}] + (1-v) z_{t+i}\} \tag{1.10}$$

According to (1.10), c_t, log consumption, fs_t, log foreign equity, λ_t (daw_t - fs_t), the interaction term of foreign equity with domestic asset wealth and the (time-varying) share of domestic asset wealth in U.S. households´ asset wealth and y_t, log labour income, cointegrate provided they are integrated of order one, I(1). Hence, time variation of the consumption-wealth ratio, i.e. a temporary deviation from the common trends, mirrors either returns on (changes of) foreign equity holdings, returns on the interaction term, changes of labour income or consumption growth or some combination of those. Admittedly, the interaction term is hard to interpret economically and so are returns on this variable. However, the only purpose of this decomposition of the consumption-wealth ratio is to examine the error-correction properties of foreign stock holdings of U.S. households in the cointegrated relation between consumption and aggregate wealth to provide a formal rationale for the predictive power of U.S. *cay* for foreign stock market returns.

1.3 Empirical Evidence: Cointegration and Error Correction

If foreign stock holdings adjust temporary deviations from the common trend among consumption and aggregate wealth, then U.S. *cay* has to mirror future returns on foreign stock markets. This is the implication from the present value representation (1.10) of the U.S. consumption-wealth ratio.

This section is thus devoted to assess the cointegration properties of my four variables comprising loglinear proxy of the U.S. consumption-wealth ratio. All variables are quarterly, per capita, real in billions of chain-weighted 2000 U.S. dollars and transformed to natural logarithms for the sample period from the second quarter of 1952 to the first quarter of 2005.

Labour income and consumption are approximated as proposed by Lettau and Ludvigson (2001a, 2004). As pointed out by Lettau and Ludvigson (2001a) as well as Rudd and Whelan (2002), the budget constraint (1.1) refers to total personal consumption flows. Since we do not observe consumption flows we rely on expenditures as best proxy. Lettau and Ludvigson (2001a) follow Blinder and Deaton (1985) and assume that the log consumption expenditure on non-durables and services consumption is a constant multiple of the logarithm of total consumption expenditure. Rudd and Whelan (2002) provide evidence that the relation between log non-durable and services consumption and log total personal consumption expenditure is not constant over time. However, Lettau and Ludvigson (2004) argue that durable consumption expenditure represents rather replacements or additions to an existing stock than a service flow from the stock of durable goods and hence is better described as wealth which is the view I follow in this chapter. The choice of consumption approximation will have an impact on the estimates of the cointegration coefficients as e.g. discussed in Hoffmann (2006a) but it turns out that the qualitative results in terms of error correction properties and forecasting power for stock returns remain unaltered, even though the by now common assumption of a constant relation between total consumption and expenditures on non-durables and services is debatable in this context.

Unit root tests show that the four variables under consideration are I(1), which conveys the notion that my four-variable approximation of the log consumption-wealth ratio should cointegrate.[3]

Table 1.1 displays results of the Johansen cointegration test, critical values for Trace and L-max test as well as the test statistics for both of the tests. Akaike (AIC) and Schwartz (SIC) information criteria suggest an appropriate lag length of one quarter for the vector autoregressive representation (VAR) of the four variables. According to the test statistics, the null of no cointegration for the relation between non-durables and services consumption expenditure excluding clothing and shoes, foreign equity holdings, domestic asset wealth interacted with foreign equity and labour income is rejected at 90 percent confidence level. Hence, the test results can be interpreted in favour of one cointegrating relation among the four variables under consideration.

The cointegrating relation (1.10) imposes the restriction that the cointegration coefficients of foreign equity holdings and the interaction term of foreign equity with domestic asset wealth should equalize. I assess these restrictions by estimating the cointegration vector via a dynamic least squares regression (DOLS) proposed by Stock and Watson (1993). OLS estimates of cointegrated variables converge to their true value with the sample size rather than with the square root of the sample size (Stock, 1987). Thus, these estimates are "superconsistent" and simple OLS provides consistent point estimates.

[3] Unit root tests are not reported to conserve space

Table 1.1: Johansen Cointegration Test

	Critical Values Trace			Test Statistic Trace
	10%	5%	1%	
r=0	44.4929	47.8545	54.6815	**45.5429**
r=1	27.0669	29.7961	35.4628	14.0055
r=2	13.4294	15.4943	19.9349	2.5366
r=3	2.7055	3.8415	6.6349	0.7659

	Critical Values L-Max			Test Statistic L-Max
	10%	5%	1%	
r=0	25.1236	27.5858	32.7172	**26.5361**
r=1	18.8928	21.1314	25.8650	11.4689
r=2	12.2971	14.2639	18.5200	1.7707
r=3	2.7055	3.8415	6.6349	0.7659

	AIC	SIC
l=1	-22.0350	-21.7791
l=2	-21.9342	-21.4224

Notes: The variables under consideration are non-durables and services consumption expenditure excluding expenditures on clothing and shoes, foreign stock holdings, log asset wealth ex foreign stocks minus log foreign stocks and labour income. All variables are measured at the quarterly frequency. The sample starts second quarter 1952 and ends first quarter 2005. All variables are in natural logarithms, real, p.c. in 2000 chain weighted U.S. dollars. The Johansen test is performed under the assumption of an unrestricted constant but no time trend in the data. The Trace test tests the null hypothesis of r cointegrating relations against the alternative of p, the number of variables in the tested system, cointegrating relations. The L-Max test tests the null of r cointegrating relations against the alternative of r+1. AIC is the Akaike information criterion, SIC the Schwartz information criterion.

But the error terms of the individual time-series variables could be correlated with each other. Hence, the OLS estimates are consistent but potentially biased. The DOLS estimate equation takes the following form:

$$\begin{aligned} c_t &= \alpha + \beta_{fe} fs_t + \beta_{daw}(\lambda_{t-i}(daw_{t-i} - fs_{t-i})) + \beta_y y_t \\ &+ \sum_{i=-k}^{k} b_{fs,i}\Delta fs_{t-i} + \sum_{i=-k}^{k} b_{daw,i}\Delta(\lambda_{t-i}(daw_{t-i} - fs_{t-i})) + \sum_{i=-k}^{k} b_{y,i}\Delta y_{t-i} + \varepsilon_t \end{aligned} \tag{1.11}$$

Estimation of the cointegration coefficients, β_i with i = fs, $\lambda(daw\text{-}fs)$, y gives the cointegration vector $\hat{\beta}$ if the coefficient on consumption (here: non-durables and services) is normalised to unity.[4] Newey-West corrected t-statistics appear in parenthesis (Newey and West, 1987). The coefficients of differences in lead or lag are omitted.

$$\hat{\beta} = [1 \quad \underset{(8.77)}{-0.22 fs_t} \quad \underset{(8.07)}{-0.21(\lambda_t(daw_t - fs_t))} \quad \underset{(24.19)}{-0.63 y_t}]'$$

Note that the cointegration coefficient estimates of foreign equity and log domestic asset wealth minus log foreign equity are approximately the same as required by the decomposition of the consumption-wealth ratio.

Engle and Granger (1987) point out that for every cointegrating relation an error-correction representation exists. Here, the error correction properties of the cointegrated system should reveal if foreign equity adjusts a temporary deviation from the common trend among consumption and aggregate wealth which would rationalize the predictive power of *cay* for foreign stock market returns.

The vector error correction representation (VECM) of $\boldsymbol{x}_t = [c_t, fs_t, \lambda_t(daw_t - fs_t), y_t]'$ is

$$\boldsymbol{\Gamma(L)\Delta x}_t = \boldsymbol{\alpha\beta' x}_{t-1} + \boldsymbol{\varepsilon}_t \tag{1.12}$$

in which $\boldsymbol{\Delta x}_t = [\Delta c_t, \Delta fs_t, \Delta(\lambda_t(daw_t - fs_t))\ \Delta y_t]'$ is the vector of first differences and $\boldsymbol{x}_{t-1}$ the vector of lagged levels, $\boldsymbol{\alpha} \equiv [\alpha_c, \alpha_{fs}, \alpha_{daw\text{-}fs}, \alpha_y]'$ is the vector of error correction coefficients. $\boldsymbol{\Gamma(L)}$ denotes a (4-by-4) matrix in the lag operator and $\hat{\boldsymbol{\beta}} = [1, -\beta_{fs}, -\beta_{daw\text{-}fs}, -\beta_y]'$ represents

[4] The estimates do not vary much from one to eight leads and lags. Here eight leads and lags are employed. Johansen´s maximum likelihood procedure provides very similar estimates. Furthermore, the cointegration coefficients sum to roughly unity if total consumption is employed in the estimation (not reported).

the vector of the above estimated cointegration coefficients when non-durables and services consumption expenditures are used as consumption proxy. The (4-by-1) vector of shocks in the cointegration relation is represented by $\boldsymbol{\varepsilon}_t$ with covariance matrix $\boldsymbol{\Omega}$. Lower-case letters in bold face denote vectors, bold upper-case letters represent matrices.

The term $\hat{\boldsymbol{\beta}}'\boldsymbol{x}_{t-1}$ gives the cointegration residual, $\boldsymbol{\alpha}$ is the adjustment vector that displays what variables adjust a deviation from the common trend. If $\boldsymbol{x}_t$ is cointegrated, at least one of the adjustment coefficients α_c, α_{fs}, α_{daw-fs} or α_y has to take values different from zero in the error-correction representation.

Table 1.2 reports VECM coefficient estimates. The lag length of one has been chosen according to Akaike, Schwartz and Hannan-Quinn information criteria. T-statistics of the coefficient estimates are in parenthesis. Bold faces indicate significance at the 95 percent confidence level.

Table 1.2: VECM estimates

	Δc_t	Δfs_t	$\Delta(\lambda_t\,(daw_t - fs_t))$	Δy_t
Δc_{t-1}	**0.2533** **(3.3216)**	1.7526 (1.1040)	−1.0757 (−0.7524)	**0.5227** **(3.3527)**
Δfs_{t-1}	0.0073 (0.3994)	−0.1228 (−0.3217	0.2228 (0.6479)	−0.0455 (−1.2131)
$\Delta(\lambda_{t-1}\,(daw_{t-1} - fs_{t-1}))$	0.0081 (0.4073)	−0.2843 (−0.6882)	0.4046 (1.0877)	−0.0557 (−1.3731)
Δy_{t-1}	**0.0942** **(2.3584)**	0.5390 (0.6483)	−0.1436 (−0.1918)	−0.0672 (−0.8229)
$\hat{\boldsymbol{\beta}}'\boldsymbol{x}_{t-1}$	−0.0057 (−0.1955)	**2.8203** **(4.6829)**	**−2.3607** **(-4.3522)**	0.0775 (1.3113)

Notes: This table reports VECM estimates for the cointegrated VAR consisting of non-durable consumption and services consumption expenditure excluding clothing and shoes, *c*, foreign stock holdings, *fs*, domestic asset wealth interacted with foreign equity holdings and the share of domestic asset wealth in asset wealth, $\lambda(daw\text{-}fs)$, and labour income, *y*, for the sample period from the second quarter of 1952 to the first quarter of 2005. $\hat{\boldsymbol{\beta}}'\boldsymbol{x}_{t-1}$ is the cointegration residual obtained with the cointegration vector: $\hat{\beta} = [1 \quad \underset{(8.77)}{-0.22 fs_t} \quad \underset{(8.07)}{-0.21\lambda_t (daw_t - fs_t)} \quad \underset{(24.19)}{-0.63 y_t}]'$

I focus on the adjustment coefficients in the last row of table 1.2. The adjustment coefficient estimates of both of the asset wealth components, foreign equity and foreign equity interacted with domestic asset wealth are statistically different from zero which mirrors the responsibility of the two components for the error correction in the cointegrated system. From (1.10) error correction through foreign equity implies that the cointegration residual, *cay*, mirrors expected returns on foreign equity and hence on foreign stock markets which thus explains the findings of Guo (2006).

1.3.1 Identification of permanent and transitory shocks and variance decomposition

To strengthen the points made in the previous subsection, I follow Hoffmann (2001a,b) in identifying permanent and transitory shocks in the cointegrating system to quantify their contribution to the forecast error variance of the level of consumption, foreign stock holdings, domestic asset wealth interacted with foreign equity and the share of domestic asset wealth in asset wealth as well as labour income.

As I regard a cointegrated system with four variables and one single cointegrating relation, there are three permanent shocks representing the innovations to the three common trends and one single transitory shock (Stock and Watson, 1988). Identification is achieved by inverting the vector error correction representation of $\boldsymbol{x}_t = [c_t, fs_t, \lambda_t (daw_t - fs_t), y_t]'$ into a multivariate Beveridge-Nelson moving average representation in terms of the reduced form disturbances (Beveridge and Nelson, 1981), which is given by

$$\boldsymbol{x}_t = \boldsymbol{C(1)} \sum_{i=0}^{t} \varepsilon_i + \boldsymbol{C}^*\boldsymbol{(L)}\varepsilon_t \tag{1.13}$$

The first term on the right-hand side of (1.13) represents the random walk and the second term the stationary component of $\boldsymbol{x_t}$.

Johansen (1995) shows that $\boldsymbol{C(1)}$ can be identified with the parameters of the VECM, such that

$$\boldsymbol{C(1)} = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Gamma(1)}\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}'_{\perp} \tag{1.14}$$

where $\boldsymbol{\beta}_{\perp}$, $\boldsymbol{\alpha}_{\perp}$ are the orthogonal complements of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$. The Granger representation theorem implies that $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ satisfy $\boldsymbol{\beta}'\boldsymbol{C(1)} = 0$ and $\boldsymbol{C(1)}\boldsymbol{\alpha} = 0$.

The common trends, π_t, thus are

$$\pi_t = \boldsymbol{\alpha}'_{\perp}\sum_{i=0}^{t}\varepsilon_i = \sum \eta_t \tag{1.15}$$

Let $\eta_t^P = \boldsymbol{\alpha}'_{\perp}\varepsilon_t$ denote the permanent shocks to the cointegrated system and $\eta_t^T = \boldsymbol{\alpha}'\boldsymbol{\Omega}^{-1}$ the transitory shock if it is orthogonal to the permanent shocks. Hence, the structural permanent shocks and the structural transitory shock are identified via

$$\boldsymbol{\eta}_t = \boldsymbol{S}\varepsilon_t \tag{1.16}$$

with $\boldsymbol{\eta}_t = \begin{pmatrix} \eta_t^P \\ \eta_t^T \end{pmatrix}$ and $\boldsymbol{S} = \begin{pmatrix} (\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Omega}\boldsymbol{\alpha}_{\perp})^{1/2}\boldsymbol{\alpha}'_{\perp} \\ (\boldsymbol{\alpha}'\boldsymbol{\Omega}\boldsymbol{\alpha})^{1/2}\boldsymbol{\alpha}'\boldsymbol{\Omega}^{-1} \end{pmatrix}$ to make sure that η_t^P and η_t^T have unit variance.

With this identification it is straightforward to quantify the contribution of the three permanent shocks and the single transitory shock to the forecast error variance of the four cointegrated variables.

Table 1.3 presents the decomposition of the forecast error variance of the levels of *c*, *fs*, *λ*(*daw-fs*) and *y* into the components that can be attributed to the three permanent shocks combined and to the transitory shock. I identify the transitory shock as orthogonal to the permanent shocks. The top panel reports variance decompositions if statistically insignificant adjustment coefficient estimates are set to zero as recommended by Gonzalo and Ng (2001). The bottom panel displays variance decompositions if all adjustment coefficients are set to their estimated values.

The transitory shock should have the strongest effect on the forecast error variance of the asset wealth components because their adjustment coefficient estimates are statistically significant. This implies that both of the variables participate in the correction of a temporary deviation from the common trends among c, fs, $\lambda(daw\text{-}fs)$ and y and hence should be primarily driven by the transitory shock. The variance decompositions mirror exactly this reasoning. Note also that the impact of the transitory shock on the variance of foreign equity is stronger than on domestic asset wealth which is in line with the magnitude of the error correction coefficient estimates. The foreign equity adjustment coefficient is larger than that of domestic asset wealth, i.e. the transitory shock has to have a stronger impact on foreign equity than on domestic asset wealth.

Consumption and labour income do not participate in the error correction mechanism. Their adjustment coefficients are statistically indistinguishable from zero, which means that both of the variables should be predominantly driven by the permanent shocks. Variance decompositions for consumption and labour income support this reasoning. Almost all of the variation in consumption and labour income can be attributed to the three permanent shocks at any time horizon.

Table 1.3: Forecast error variance decompositions of the levels of the four cointegrated variables

	$\alpha_c = \alpha_y = 0$							
	c_{t+h}-$E_t(c_{t+h})$		fs_{t+h} - $E_t(fs_{t+h})$		λ_{t+h} (daw_{t+h} - fs_{t+h}) -E_t (λ_{t+h} (daw_{t+h}- fs_{t+h}))		y_{t+h} - $E_t(y_{t+h})$	
h	P	T	P	T	P	T	P	T
1	1.0000	0.0000	0.1567	0.8433	0.2716	0.7284	1.0000	0.0000
4	0.9999	0.0001	0.3645	0.6355	0.4624	0.5376	0.9994	0.0006
8	0.9999	0.0001	0.5675	0.4325	0.6452	0.3548	0.9994	0.0006
16	0.9999	0.0001	0.7959	0.2041	0.8332	0.1668	0.9994	0.0006
24	0.9999	0.0001	0.8805	0.1195	0.9009	0.0991	0.9994	0.0006
	α_c and α_y estimated							
	c_{t+h}-$E_t(c_{t+h})$		fs_{t+h} - $E_t(fs_{t+h})$		λ_{t+h} (daw_{t+h} - fs_{t+h}) -E_t (λ_{t+h} (daw_{t+h}- fs_{t+h}))		y_{t+h} - $E_t(y_{t+h})$	
h	P	T	P	T	P	T	P	T
1	0.9986	0.0014	0.1957	0.8043	0.3053	0.6947	0.9369	0.0631
4	0.9984	0.0016	0.4603	0.5397	0.5509	0.4491	0.9620	0.0380
8	0.9983	0.0017	0.6575	0.3425	0.7224	0.2766	0.9800	0.0200
16	0.9884	0.0016	0.8264	0.1736	0.8616	0.1384	0.9891	0.0109
24	0.9884	0.0016	0.8861	0.1139	0.9088	0.0912	0.9926	0.0074

Notes: This table reports the forecast error variance share of the level of the cointegrating variables, consumption, *c*, foreign equity, *fs*, domestic asset wealth interacted with foreign stock holdings and the share of domestic asset wealth in asset wealth, λ(*daw-fs*) and labour income, *y*, that can be attributed to the combined three permanent shocks (columns "P") and the single transitory shock (columns "T"). The forecast horizon h is in quarters.

1.4 U.S. cay and excess returns on foreign stock markets

The forecast ability of *cay* for stock market returns has raised continuous debates. Brennan and Xia (2005) are concerned that the predictive power of *cay* is due to a look-ahead bias because the cointegration parameters used to estimate *cay* are estimated over the full sample period which explains why *cay* does not predict stock returns out-of-sample but only in-sample. Additionally, Hahn and Lee (2003) argue that the forecasting power of *cay* is largely driven by a time trend.

In their reply to this critique, Lettau and Ludvigson (2005 a, b) show that the success of *cay* in explaining excess returns does not necessarily involve estimation of the cointegration coefficients. Hence, look-ahead bias by estimating the cointegration coefficients is not likely a reason for the predictive power of *cay*. Moreover, they argue that if one estimates the cointegration coefficients, it would be inappropriate not to use information from the full sample period since cointegration is a long-run phenomenon. Hoffmann (2006b) shows that the Lettau and Ludvigson (2001a, 2004) framework implies two cointegrating relationships that are not discovered by standard cointegration tests because of deterministic trends and a structural break. However, he also shows that the predictive power of *cay* is still a salient feature of the data if one properly takes account of the before mentioned issues.

I thus focus on the in-sample predictability of long-horizon excess returns on the Morgan Stanley Capital International (MSCI) indices of the G7 stock markets. Since *cay* is only observed at quarterly frequency, I calculate end-of-quarter returns from monthly data on the MSCI G7 indexes denominated in U.S. dollar. Following Campbell and Hamao (1992), the U.S. three-month t-bill rate is used to obtain excess returns. Table 1.4 reports estimates of regressions of the form

$$r_{t+h}^{i,e} = \alpha_h + \beta_h cay_t + \varepsilon_{t+h} \tag{1.17}$$

with $r_{t+h}^{i,e}$ the log excess return on the MSCI stock index of country *i* at time horizon *t*+*h*. Newey-West corrected t-statistics appear below the regressor estimates (Newey and West , 1987). R^2 reports the adjusted R^2. Bold faces highlight significant estimates.

Table 1.4: Long-horizon regressions of foreign stock market excess returns on U.S. cay

	h=1	h=4	h=8	h=12	h=16	h=20	h=24
CND	1.03 (1.45)	2.97 (1.17)	5.54 (1.71)	**9.02** **(2.94)**	**9.52** **(2.73)**	**7.77** **(2.00)**	**6.69** **(2.32)**
R^2	0.01	0.02	0.06	0.16	0.20	0.09	0.08
FRA	**2.26** **(2.28)**	**8.41** **(2.84)**	**12.13** **(3.39)**	**19.01** **(5.00)**	**19.27** **(4.68)**	**14.09** **(3.37)**	**9.16** **(2.59)**
R^2	0.04	0.14	0.15	0.29	0.24	0.11	0.05
GER	**1.77** **(2.08)**	**6.23** **(2.55)**	**8.36** **(2.04)**	**12.22** **(2.05)**	10.62 (1.59)	4.61 (0.73)	0.62 (0.17)
R^2	0.03	0.09	0.09	0.15	0.09	0.01	0.01
ITA	1.42 (1.36)	**7.54** **(2.02)**	**14.42** **(2.33)**	**24.43** **(3.41)**	**26.54** **(3.25)**	**23.40** **(2.55)**	20.28 (1.85)
R^2	0.01	0.08	0.14	0.28	0.27	0.18	0.12
JPN	0.52 (0.55)	0.65 (0.15)	1.94 (0.27)	4.20 (0.45)	1.11 (0.0.12)	−9.45 (−1.35)	−19.05 (−1.96)
R^2	0.00	0.00	0.00	0.00	0.00	0.02	0.10
UK	**2.44** **(2.96)**	**9.56** **(3.31)**	**13.96** **(4.26)**	**19.31** **(4.94)**	**16.55** **(4.59)**	**12.87** **(3.18)**	**6.53** **(2.00)**
R^2	0.06	0.24	0.30	0.37	0.23	0.13	0.03
US	**2.08** **(3.55)**	**6.49** **(3.72)**	**11.85** **(5.13)**	**16.38** **(4.36)**	**18.47** **(4.35)**	**17.91** **(3.89)**	**17.49** **(4.41)**
R^2	0.07	0.20	0.34	0.44	0.45	0.34	0.31

Notes: This table displays OLS estimates from regressions of the form

$$r_{t+h}^{i,e} = \alpha_h + \beta_h cay_t + \varepsilon_{t+h}$$

with $r_{t+h}^{i,e}$ the log excess return on the MSCI stock index of country *i* at time horizon t+h. Newey-West (Newey and West, 1987) corrected t-statistics appear below the regressor estimates. R^2 reports the adjusted R^2. Bold faces highlight significant estimates. The sample period runs from 1969Q4 to 2005Q1.

The only exception is Japan for which *cay* does not reveal predictive power at any time horizon. But apart from Japan all stock market excess returns of the remaining G7 economies are predicted by *cay* which corroborates Guo (2006). The peak of the forecast ability of *cay* is reached at time horizons between 12 and 16 quarters consistent with theoretical macroeconomic models that motivate time-varying risk premia with time variation in risk aversion (Constantinides, 1990; Constantinides and Duffie, 1996; Campbell and Cochrane, 1999; Heaton and Lucas, 2000a,b; Vissing-Joergensen, 2002; Polkovnichenko, 2004).

The R^2 statistics display that substantial proportions in the variation of foreign stock market returns are explained by *cay*. The highest R^2 statistics for foreign stock markets is 0.37 for the United Kingdom.

Evidence for predictability of stock returns from long-horizon regressions in relatively small samples should be regarded with healthy scepticism since the standard errors of the regressor estimates rely on asymptotic distribution theory (see e.g. Valkanov (2003) and the literature surveyed therein). However, Hodrick (1992) shows that a large amount of long-horizon predictability is consistent with a small portion of short-run predictability. Here, for four of seven countries, U.S. *cay* predicts the one-quarter excess returns statistically significantly. Guo (2006) examines a wider cross-section of countries at the one-quarter horizon and finds most of developed economies' stock market returns to be explained by *cay*. Hence, it seems justified to conclude that U.S. *cay* predicts foreign stock market returns. The (conventional) long-horizon regressions conducted in this section serve to illustrate at what time horizon the predictive power of *cay* for foreign stock market returns reaches its peak. As emphasized above, the peak of predictability at three and four year horizon is consistent with theoretical macroeconomic models.

When stock market excess returns denominated in local currency are considered, the qualitative results remain the same. Moreover, the results do not change if I employ the respective countries´ short-term interest rate to obtain excess returns (not reported).

The forecast ability of U.S. *cay* for risk premia on foreign stock markets is interesting in the light of recent studies that examine the predictive power of national consumption-wealth ratios for national stock markets. Evidence by Fernandez-Corugedo et al. (2003), Fisher and Voss (2003), Tan and Voss (2004) and Ioannidis et al. (2005) suggests that short-run fluctuations in national consumption-wealth ratios predict national stock market returns. However, these studies focus on Anglo-Saxon countries while Hamburg et al. (2007) present evidence that German *cay* does not predict German stock market returns but macroeconomic variables as the unemployment rate. The predictive power of a Japanese consumption-wealth ratio for national stock market returns is virtually zero as well (Nagayasu, 2006). Obviously, U.S. *cay* explains risk premia on both Anglo-Saxon and European countries' stock markets.

The explanatory power of *cay* for excess returns on the G7 stock markets seems to be substantial. Recent empirical studies (Hau and Rey, 2004, 2006; Lane and Milesi-Ferretti, 2001) emphasize that cross-border equity flows have increased tremendously in the last two decades. Hence the explanatory power of *cay* should be especially pronounced since the mid-1980s. In order to visualize this point, figures 1.2 to 1.8 present realised three-year excess returns on the G7 stock markets together with the corresponding fitted values from regression (1.17).

Figures 1.2. – 1.8: Predicted versus realized three-year returns

Figure 1.2: Canada

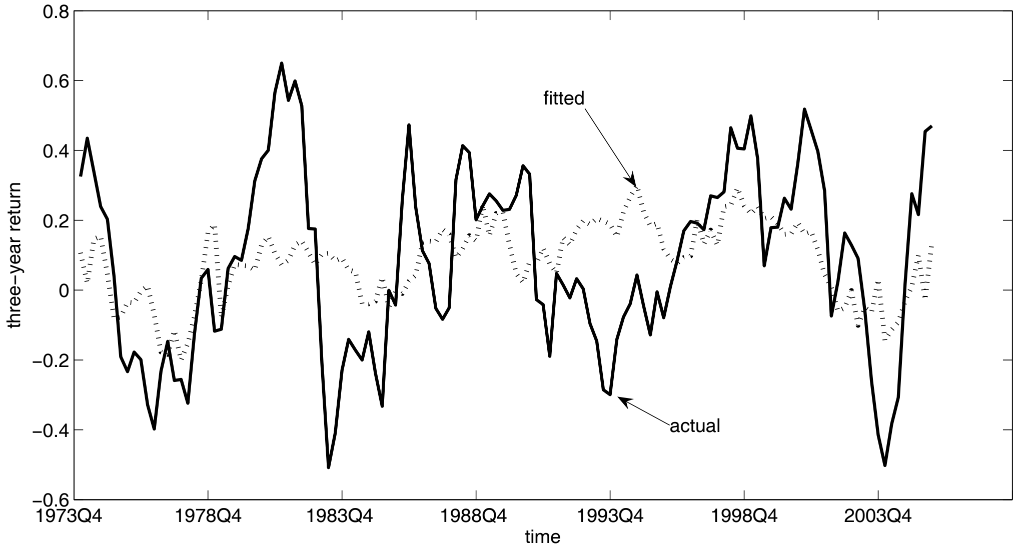


Figure 1.3: France

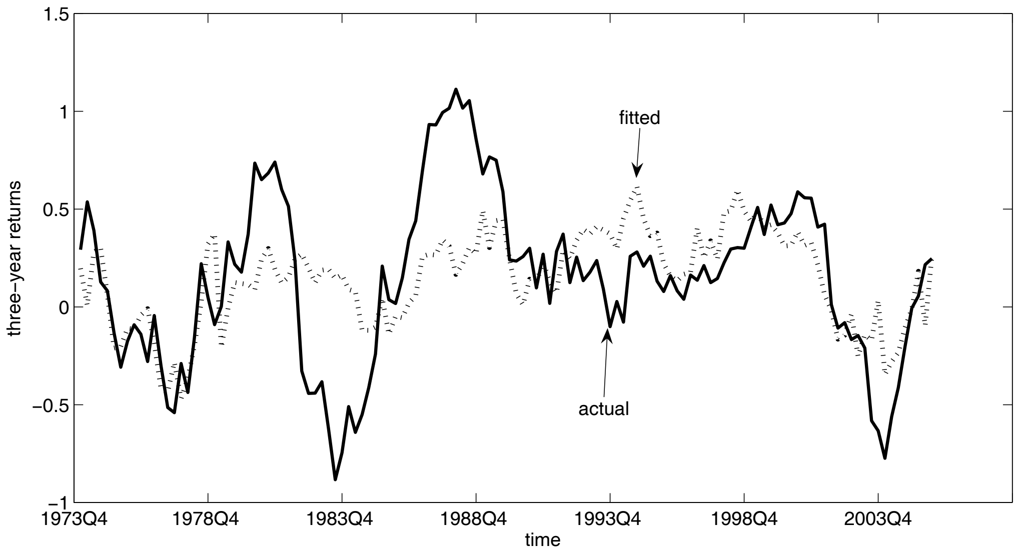


Figure 1.4: Germany

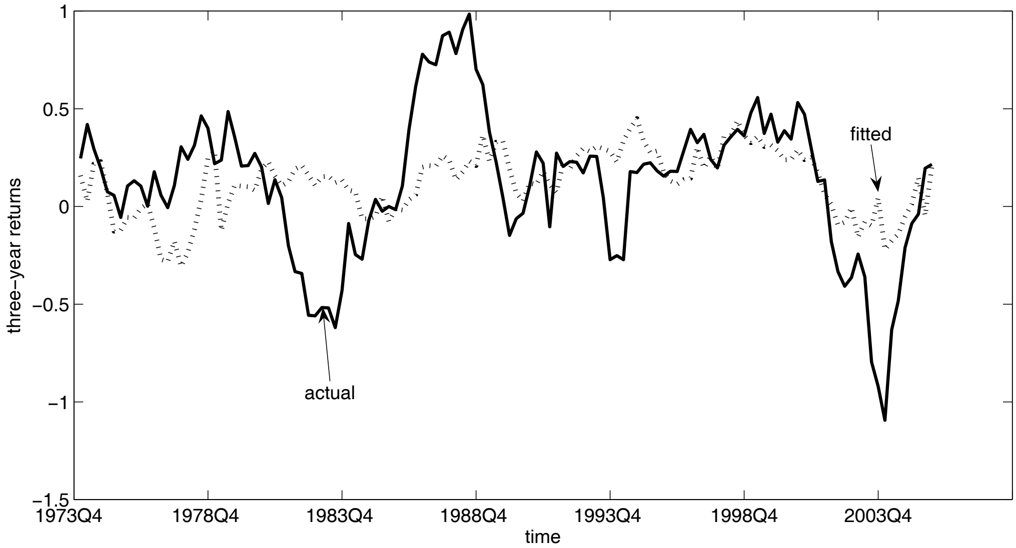


Figure 1.5: Italy

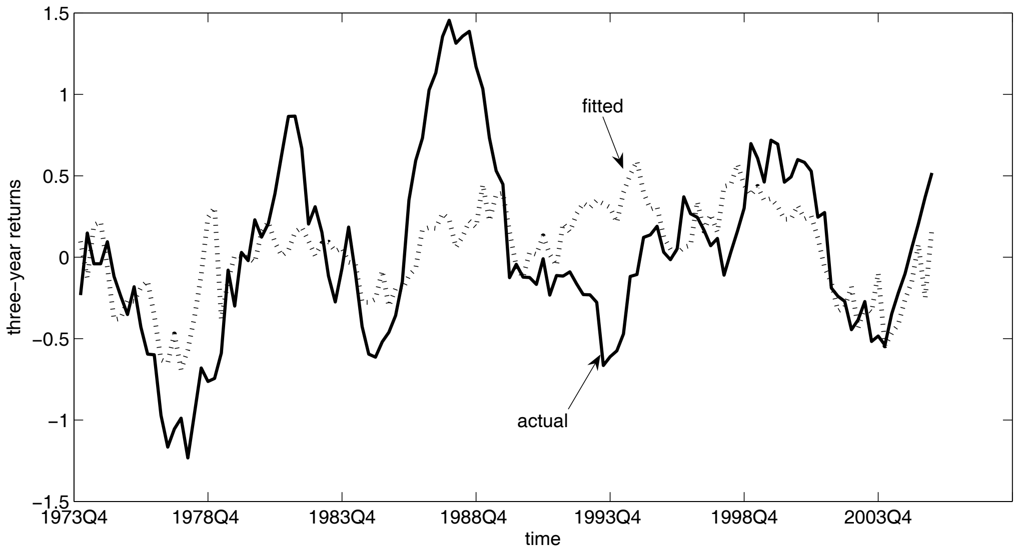


Figure 1.6: Japan

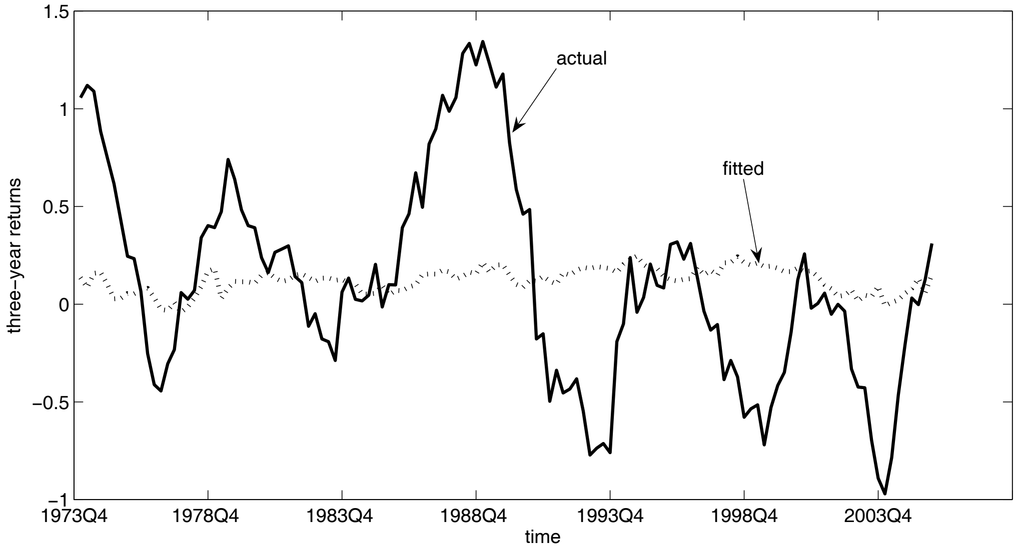


Figure 1.7: United Kingdom

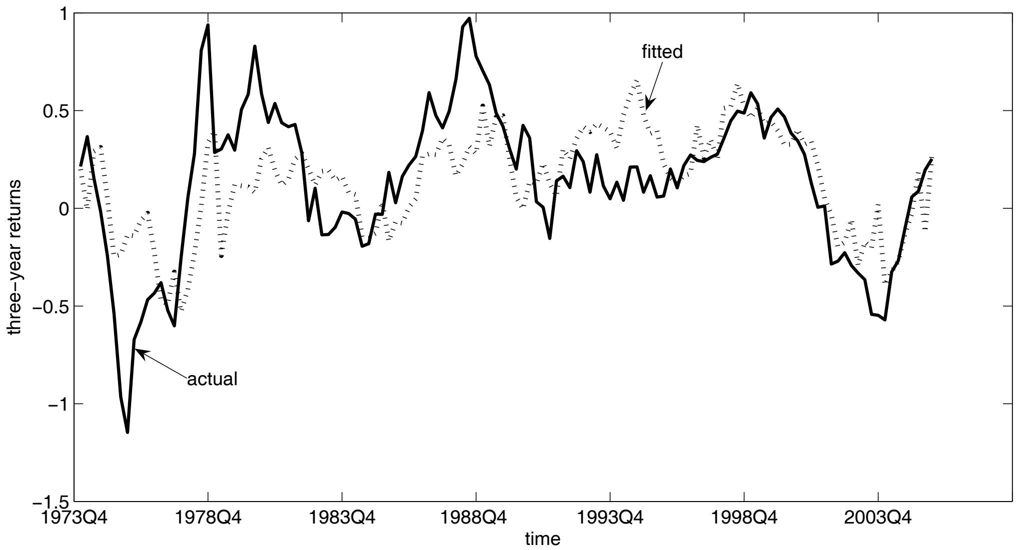
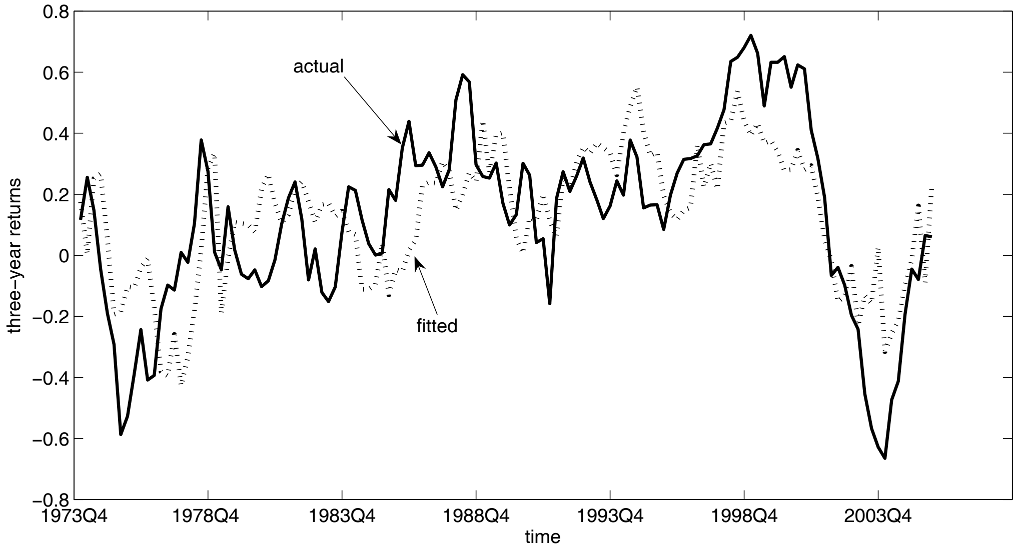


Figure 1.8: United States



The overall picture that emerges is that *cay* captures the three-year variation in long-horizon excess returns reasonably well, especially in the sample period from 1990 to 2005 thus corroborating the conjecture made above.

As suggested by the cointegration framework, the U.S. consumption-wealth ratio is informative about the future path of foreign stock markets. But the question remains if the forecast ability of *cay* survives once a country-specific predictor of stock returns, e.g. the respective national dividend-price ratio, is considered as an additional forecast variable in the long-horizon regressions.

I construct national dividend-price ratios from MSCI total return and price indexes as described in Kasa (1992) to run regressions of the form

$$r_{t+h}^{i,e} = \alpha_h + \beta_{h,cay} cay_t + \beta_{h,dp}(d_t^i - p_t^i) + \varepsilon_{t+h} \tag{1.18}$$

with $d_t^i - p_t^i$ the log dividend-price ratio of country *i*. Table 1.5 presents the results.

Table 1.5: Long-horizon regressions of foreign stock market excess returns on U.S. cay and national dividend-price ratio (sample period 1969Q4 – 2005Q1)

	h=1	h=4	h=8	h=12	h=16	h=20	h=24
CND							
cay	1.08	2.50	4.79	**8.61**	**9.45**	7.90	**7.47**
	(1.54)	(0.96)	(1.51)	**(2.66)**	**(2.52)**	(1.76)	**(2.31)**
d-p	0.00	0.08	0.12	0.11	0.10	0.24	0.27
	(0.06)	(0.89)	(1.19)	(1.04)	(1.00)	(1.44)	(1.58)
R^2	0.00	0.03	0.08	0.19	0.23	0.17	0.16
FRA							
cay	**2.31**	**8.62**	**12.53**	**20.03**	**21.17**	**17.11**	**12.51**
	(2.31)	**(2.92)**	**(3.56)**	**(5.76)**	**(5.03)**	**(3.59)**	**(2.56)**
d-p	0.00	0.03	0.06	0.11	0.16	0.22	0.17
	(0.21)	(0.68)	(0.78)	(1.29)	(1.63)	(1.37)	(1.00)
R^2	0.03	0.14	0.16	0.32	0.31	0.18	0.09
GER							
cay	**1.91**	**6.96**	**9.98**	**14.81**	**14.42**	10.72	6.65
	(2.19)	**(2.84)**	**(2.51)**	**(2.73)**	**(2.50)**	(1.59)	(1.30)
d-p	0.02	0.09	0.17	**0.25**	**0.37**	**0.44**	**0.35**
	(0.87)	(1.47)	(1.69)	**(1.96)**	**(2.60)**	**(2.60)**	**(2.12)**
R^2	0.02	0.11	0.14	0.24	0.26	0.20	0.11
ITA							
cay	1.48	**7.39**	**13.69**	**23.97**	**26.84**	**23.68**	20.46
	(1.36)	**(1.96)**	**(2.16)**	**(3.16)**	**(3.05)**	**(2.34)**	(1.76)
d-p	0.01	0.11	**0.26**	0.18	0.05	0.02	– 0.00
	(045)	(1.15)	**(2.00)**	(1.95)	(0.48)	(0.08)	(–0.01)
R^2	0.00	0.10	0.19	0.30	0.28	0.19	0.13
JPN							
cay	1.17	3.60	7.52	12.87	12.34	5.40	– 0.80
	(1.29)	(0.95)	(1.32)	(1.84)	(1.58)	(0.79)	(–0.10)
d-p	0.04	**0.18**	**0.34**	**0.46**	**0.58**	**0.62**	**0.66**
	(1.95)	**(2.13)**	**(2.93)**	**(4.30)**	**(5.27)**	**(4.20)**	**(4.66)**
R^2	0.02	0.10	0.17	0.25	0.31	0.31	0.37
UK							
cay	**2.51**	**9.64**	**14.36**	**20.61**	**18.57**	**15.97**	**12.00**
	(2.87)	**(3.37)**	**(4.86)**	**(5.83)**	**(4.89)**	**(5.16)**	**(6.16)**
d-p	0.05	**0.18**	**0.29**	**0.50**	**0.64**	**0.77**	**0.82**
	(1.25)	**(2.71)**	**(3.48)**	**(4.55)**	**(6.21)**	**(7.95)**	**(8.37)**
R^2	0.08	0.31	0.40	0.59	0.60	0.58	0.59
US							
cay	**2.11**	**6.34**	**11.47**	**15.89**	**18.06**	**17.57**	**18.08**
	(3.63)	**(3.58)**	**(4.80)**	**(4.01)**	**(4.17)**	**(3.91)**	**(4.92)**
d-p	– 0.00	0.02	0.06	0.12	0.15	**0.25**	**0.29**
	(–0.02)	(0.41)	(0.61)	(1.12)	(1.44)	**(3.32)**	**(2.80)**
R^2	0.08	0.20	0.36	0.48	0.54	0.44	0.40

Notes: This table displays regressor estimates from regressions of the form

$$r_{t+h}^{i,e} = \alpha_h + \beta_{h,cay} cay_t + \beta_{h,dp}(d_t^i - p_t^i) + \varepsilon_{t+h}$$

with $r_{t+h}^{i,e}$ the log excess return on the MSCI stock index of country *i* at time horizon t+h and $(d_t^i - p_t^i)$ the respective national log dividend-price ratio. Newey-West (Newey and West, 1987) corrected t-statistics appear below the regressor estimates. R^2 reports the adjusted R^2. Bold faces highlight significant estimates.

It is apparent that the forecast ability of *cay* is not materially affected by the presence of the national dividend-price ratio as additional regressor. Noteworthy is the finding that the Japanese, British and German dividend-price ratio predict excess returns on the respective national stock market. While the British and German dividend-price ratio as well as *cay* jointly predict long-horizons returns on the respective MSCI index, the Japanese dividend-price ratio is the only variable with forecast ability of excess returns on the Japanese stock market.

These results corroborate that a common temporary component in national stock markets exists. Its importance is mirrored in the predictive power of *cay* for foreign stock returns. The impact of the common component on international stock markets is highest at business cycle frequency. In addition, *cay* forecasts risk premia on foreign stock markets even in the presence of country-specific predictors of stock returns.

1.5 Covariation of returns

As Campbell and Hamao (1992) point out, a single variable that predicts stock returns on several national stock markets can be interpreted as the reflection of a common stock market component. Furthermore, the sensitivity of national stock markets to this common component should explain covariation among these markets.

Following this logic, the predictive power of U.S. *cay* seems to reflect a temporary component common to national stock markets as it explains a considerable fraction of the variation in risk premia on the G7 stock markets. Hence we should also obtain information about the degree of covariation between the G7 stock markets by their sensitivity to *cay*. Information about stock market comovement should thus be inferred from

$$\boldsymbol{r} = \boldsymbol{\gamma} cay + \boldsymbol{\varepsilon} \tag{1.19}$$

where $\boldsymbol{r}$ is the N-by-1 vector of excess returns on the G7 markets at a particular time horizon with N the number of returns, *cay* reflects the common temporary component, $\boldsymbol{\gamma}$ denotes the N-by-1 vector of coefficients from the long-horizon regressions (1.17) at the respective time horizon, and $\boldsymbol{\varepsilon}$ denotes the N-by-1 vector of the corresponding error terms.

Hence, the covariance of the excess returns obeys

$$\boldsymbol{var(r)} = \boldsymbol{\gamma\gamma}'\,\mathrm{var}(cay) + \boldsymbol{var(\varepsilon)} \tag{1.20}$$

so that we can measure covariation implied by the exposure to *cay* with a multivariate "R^2" measure of the form

$$\mathbf{R}^2 = \mathbf{I} - \frac{\boldsymbol{var(\varepsilon_{i,j})}}{\boldsymbol{var(r_{i,j})}} \tag{1.21}$$

I focus on the covariation between the 12-quarter excess returns on the MSCI stock indexes of the G7 economies as the predictive power of *cay* and hence the influence of the common component peaks at that horizon. Table 1.6 presents my measure of covariation for the three-year excess returns of the G7 economies for the sample period from 1973 to 2005. The values on the diagonal reflect the R^2 statistics from the long-horizon regressions while the off-diagonal elements give the pairwise amount of covariation among the three-year returns on the G7 stock markets implied by their exposure to *cay*. As an example, consider the row GER and its intersection with the column FRA. The number of 0.27 in this cell says that 27 percent of the common movement in the French and German three-year stock returns can be rationalized by their sensitivity to *cay*, the mirror image of the common temporary component.

Table 1.6: Covariation of 12-quarter returns (1973Q4 – 2005Q1)

	CND	FRA	GER	ITA	JPN	UK	US
CND	0.16						
FRA	0.34	0.29					
GER	0.38	0.27	0.15				
ITA	0.33	0.38	0.39	0.27			
JPN	0.08	0.08	0.06	0.09	0.01		
UK	0.37	0.43	0.37	0.51	0.11	0.35	
US	0.42	0.56	0.37	0.61	0.23	0.49	0.40

Notes: This table presents the amount of covariation in three-year returns on the G7 stock markets explained by their exposure to the common, temporary stock market component reflected in *cay*.

Not surprisingly, little of the Japanese stock market´s covariation with the other G7 economies is captured by the common component. It varies between one and 23 percent with regard to the U.S. But *cay* explains between 15 to 60 percent of the covariation between 12-quarter excess returns on the remaining G7 stock markets. Interestingly, the comovement among the three core European countries is less pronounced than their common movement with the U.S.

There is still some covariation between the G7 markets that is unexplained which could be caused by a common permanent component in international stock markets or due to country-specific effects as suggested by Richards (1995). Nonetheless, table 1.6 highlights the importance of the international stock market component in explaining the covariation of national stock markets at the business cycle frequency.

Furthermore, this finding sheds light on the degree of integration of the G7 stock markets. Comovement caused by a common component can be interpreted as reflection of integration of capital markets (Campbell and Hamao, 1992). The amount of covariation explained by the common stock market component is not negligible which is suggestive of relatively well integrated stock markets among the G7.

In addition, figures 1.2 to 1.8 convey the notion that the importance of the common component for the covariation of the G7 returns is particularly pronounced in the sample period from 1990 to 2005. Table 1.7 reports the amount of covariation explained by *cay* for this sample period.

Table 1.7: Covariation of 12-quarter returns (1990Q1 – 2005Q1)

	CND	FRA	GER	ITA	JPN	UK	US
CND	0.22						
FRA	0.41	0.50					
GER	0.43	0.58	0.44				
ITA	0.31	0.39	0.57	0.10			
JPN	0.27	0.36	0.31	0.43	0.05		
UK	0.51	0.65	0.65	0.49	0.68	0.63	
US	0.49	0.63	0.55	0.65	0.55	0.67	0.58

Notes: see table 1.6

Consistent with the visual impression from figures 1.2 to 1.8 the explanatory power of the international stock market component for comovement of the G7 markets is considerably stronger than in the full sample. This finding does even pertain to Japan vis-à-vis the other G7 economies. The common temporary component explains up to 70 percent of the covariation among the G7 stock markets since the 1990s and leaves the impression of increased market integration among the major industrialized countries.

1.6 Conclusions

This chapter corroborates and extends earlier evidence for the existence of a common temporary component in international stock markets that is reflected in the predictive power of short-run variations in the U.S. consumption-wealth ratio, *cay*, for excess returns on foreign stock markets at the business cycle frequency. This common component is responsible for 15 to 60 percent of the covariation between 3-year excess returns on the G7 stock markets in the period from 1969Q4 to 2005Q1. This finding is most pronounced for the sample period from 1990 to 2005 where we date the onset of financial globalisation. Taken together these findings are suggestive of an increase in the importance of the common stock market component for the time series variation in and covariation between major stock markets. The latter result ties in with empirical studies that highlight the growing importance of global factors for the comovement of national stock markets at relatively high (monthly) frequency (Brook and DelNegro, 2006).

In addition, this chapter provides a formal rationale for the predictive power of U.S. *cay* for excess returns on foreign stock markets. The market value of U.S. households' foreign equity holdings changes in response to expected returns on foreign stocks to adjust deviations from the common trend among consumption and aggregate wealth in the U.S.

Chapter 2

Cashflow news, the value premium and an asset pricing view on European stock market integration

2.1 Introduction

The main results of the preceding chapter convey the notion that a common, temporary component explains substantial proportions of the comovement between stock markets of the G7 countries. A strong covariation among the stock markets of the major industrialized economies seems to be a salient feature of the data irrespective of the frequency at which it is regarded. This evidence leaves the impression of relatively well integrated stock markets which implies that differences in the sensitivity to common, global factors should explain differences in average returns on international stock markets (Harvey, 1991; Ferson and Harvey, 1993).

However, if capital markets are sufficiently integrated, then cross-sectional dispersion in international asset returns should be explained by national risk factors as well. It is this latter line of thought that this chapter pursues. This argument follows immediately from the so called basic pricing equation for asset returns

$$1 = E_t(M_{t+1}R^i_{t+1}) \tag{2.1}$$

with M_{t+1} the stochastic discount factor and R^i_{t+1} the gross return on asset or portfolio i. In words, an expected asset return should be constant once discounted with the stochastic discount factor (SDF) that is the same for all assets. This equation should hold for any asset from a national investor’s point of view. The question is now what is the form of the national SDF, i.e. what model explains asset returns, and what are the asset returns to examine.

Value stocks, defined as stocks with high book value relative to market value (B/M), high earnings-to-price ratio (E/P), high cashflow-to-price ratio (C/P) and high dividend-to-price ratio (D/P) receive a lot of attention by practitioners as well as academics since they offer higher average returns than expected from their market betas in a Sharpe (1964) and Lintner (1965) capital asset pricing model (CAPM). Conversely, growth stocks (stocks with e.g. low book-to-market value ratio) promise lower returns than predicted by the CAPM. This finding is not a unique observation on U.S. stock markets but by now well documented in international data (e.g. Capaul et al., 1995; Chan et al.,1991; Fama and French, 1998).

The Sharpe-Lintner CAPM assumes the existence of a so called market portfolio comprising all risky assets. The excess return on this market portfolio is a measure of all systematic sources of risk. Differences in the sensitivity to the market return (“betas”) should thus explain differences in average asset returns. In empirical work the market return is typically proxied by broad stock market indexes. While this practice can be criticized on various grounds (e.g. Roll, 1977; Campbell, 1996; Jagannathan and Wang, 1996; Lettau and Ludvigson, 2001a), Davis et al. (2000) show that the CAPM works well when confronted with U.S. value and growth stock data from the sample period from 1929 to 1963 but works poorly in the modern time period from 1963 to the present.

Campbell and Vuolteenaho (2004) explain the difference in the performance of the CAPM in the two sample periods by decomposing CAPM market betas into a cashflow ("bad") and discount rate ("good") variety. Intuitively, bad news about the market´s future cashflows reflect a decrease of wealth and hence lead to a fall in the value of the market but leave future investment opportunities unaffected. The value of the market portfolio could also decline because investors increase the discount rate applied to cashflows, which at the same time mirrors better future investment opportunities. Furthermore, the intertemporal CAPM of

Merton (1973) suggests that the receptiveness to innovations in dividends (cashflows) should be rewarded with a higher price of risk than sensitivity to discount rate news. Campbell and Vuolteenaho (2004) thus show that value stocks´ market betas in U.S. post-war data contain a substantially higher cashflow component than growth stocks´ market betas which explains seemingly abnormally high average returns on value portfolios.[5]

Based on these findings for the U.S., the contribution of this chapter is twofold. On the one hand, I ask if the explanation for differences in average returns on value and growth portfolios proposed by Campbell and Vuolteenaho (2004), Cohen et al. (2003) and Campbell et al. (2005) applies to international data from a national investor's perspective. On the other hand, I assess in how far national two-beta versions of the CAPM capture the international cross-sectional dispersion in average returns on seven European countries[6]. I use the European value and growth portfolios for the latter exercise as well. But note that it is not the aim of this chapter to explain the cross-sectional dispersion of returns in the value versus growth domain. Rather the question is: Can we explain why e.g. the average return on the Belgium value portfolio is different from the German value portfolio from a Swiss investor's point of view? But the sorting of stocks into portfolios according to book-to-market ratios guarantees large spreads in average returns and thus facilitates the cross-sectional analysis.

The countries taken into question in this chapter (with the exception of Switzerland) experienced the establishment of the European monetary system with a period of monetary and fiscal policy convergence, interrupted by several currency crises in the 1990s and finally the launch of the European Monetary Union (EMU) in 1999. The empirical exercises conducted in this chapter can thus also be viewed as an assessment of the integration of the core European stock markets in the sample period from the first quarter of 1975 to the fourth quarter of 2005.

[5] Lettau and Wachter (2004) underscore this intuition focusing on value and growth returns in the U.S.

[6] These countries are Belgium, France, Germany, Italy, the Netherlands, Spain and Switzerland

From an asset pricing perspective, the choice of European countries is guided by two additional considerations. First, I would like to minimize the impact of exchange rate risk and focus on purely stock market based explanations of cross-sectional differences in stock returns. Various versions of international asset pricing models show that exchange rate risks are an important factor in explaining the cross-sectional dispersion in international stock market returns (e.g. Dumas and Solnik, 1995; Gerard and De Santis, 1997; Harvey, 1991; Solnik, 1974). Since this paper concentrates on the core EMU countries plus Switzerland, the impact of foreign exchange risks on cross-sectional stock returns is likely to be relatively small. Secondly, Lane and Milesi-Feretti (2005) show that European investors predominantly invest into euro-area equity. Thus the European value and growth portfolios should represent the investment opportunities of national investors in European countries.

The main results are easily summarized. European value stocks offer higher excess returns than their growth portfolio counterparts. In line with the findings of Campbell and Vuolteenaho (2004), Cohn et al. (2003) and Campbell et al. (2005), high average returns are associated with disproportionately high cashflow components in national market betas. Furthermore, from the perspective of a national investor, average returns on European value and growth portfolios can be reconciled with two-beta variants of national CAPMs. Hence, the implications of capital market integration on asset pricing theory seem to be reflected in the sample of European countries under consideration.

Interestingly, even though high average returns go hand in hand with high cashflow betas relative to discount rate betas, cross-sectional differences among returns on the European value and growth stocks seem to be explained by differences in national discount rate betas for five of the seven countries. The lower the sensitivity to better than expected discount rate news, i.e. “good” news, the higher is the average return. Differences in national cashflow betas explain cross-sectional dispersion in European stock returns from the perspective of a Belgium and Dutch investor.

The remainder of this chapter is organized as follows. In section 2.2, I sketch the framework of Campbell (1991) and Campbell and Vuolteenaho (2004) used to identify cashflow and discount rate betas. Thereafter, I briefly discuss the choice of state variables in section 2.3 and provide details of the data employed in this paper in 2.4. Section 2.5 discusses the empirical evidence. Finally, section 2.6 concludes.

2.2 Theoretical framework

The identification of cash flow and discount rate news driven components in simple and excess stock returns is based on the relationship between prices, dividends and returns as formulated in the dividend ratio model of Campbell and Shiller (1988a).

A log-linear approximation of the stock return, R_{t+1}, gives

$$r_{t+1} \approx k + \rho p_{t+1} + (1-\rho)d_{t+1} - p_t \tag{2.2}$$

where r_{t+1} is the log stock return, p_t the log stock price at time t, d_{t+1} log dividends, k summarizes constant terms and $\rho = \dfrac{1}{1+\exp(d-p)}$,with d - p the long-run mean of the log dividend-price ratio, is a weight obtained in the log-linearization. Rearranging (2.2) for the stock price, expanding to the infinite horizon and taking expectations on both sides of the equation yields

$$p_t = \frac{k}{1-\rho} + E_t\left[\sum_{j=0}^{\infty} \rho^j ((1-\rho)d_{t+1+j} - r_{t+1+j})\right] \tag{2.3}$$

Substituting (2.3) into (2.2), Campbell (1991) shows that unexpected changes in stock returns obey

$$r_{t+1} - E_t r_{t+1} = (E_{t+1} - E_t)\sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - (E_{t+1} - E_t)\sum_{j=1}^{\infty} \rho^j \Delta r_{t+1+j} \tag{2.4}$$

where Δ denotes the difference operator and E_t rational expectations at time t. Revisions of expected future dividend growth are written as $(E_{t+1} - E_t)\sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j}$, changes of future discount rates as $(E_{t+1} - E_t)\sum_{j=1}^{\infty} \rho^j \Delta r_{t+1+j}$.

Equation (2.4) states that unexpected changes of stock returns have to be associated with revisions of expectations of future cashflows or discount rates or both. Following Campbell (1991), equation (2.4) can be written in more compact notation as

$$v_{r,t+1} = \eta_{CF,t+1} - \eta_{DR,t+1} \tag{2.5}$$

with $v_{r,t+1} \equiv r_{t+1} - E_t r_{t+1}$ the unexpected component of the stock return, $\eta_{CF,t+1} \equiv (E_{t+1} - E_t)\sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j}$ representing news about dividend changes, i.e. cash flows and $\eta_{DR,t+1} \equiv (E_{t+1} - E_t)\sum_{j=1}^{\infty} \rho^j \Delta r_{t+1+j}$ which denotes news about returns, i.e. discount rates.

In order to identify cash flow and discount rate components in stock returns, Campbell (1991) suggests to use a first-order VAR of the form

$$\boldsymbol{z}_{t+1} = \boldsymbol{\mu} + \boldsymbol{\Gamma z}_t + \boldsymbol{u}_{t+1} \tag{2.6}$$

where $\boldsymbol{z}_{t+1}$ is a k-by-1 state vector with the stock return, r_{t+1}, as first element and variables which predict stock returns, $\boldsymbol{\mu}$ is a k-by-1 vector of constants and $\boldsymbol{\Gamma}$ a k-by-k matrix of VAR parameters. Shocks are i.i.d. and represented by the k-by-1 vector $\boldsymbol{u}_{t+1}$. The assumption of a first-order VAR is not restrictive because a higher-order VAR can be written in first-order companion form (Campbell and Shiller, 1988a).

Since the state vector, $\boldsymbol{z}_{t+1}$, includes variables that predict stock returns, the discount rate news component is directly estimated in the VAR whereas the cash flow news component is a residual. It is that part of the return which is not explained by the state variables.

Under the assumption that the data is generated by (2.6), forecasts of future returns obey

$$E_t r_{t+1+j} = \mathbf{e1}'\boldsymbol{\Gamma}^{j+1}\boldsymbol{z}_t \tag{2.7}$$

with $\boldsymbol{e1}$ a k-by-1 vector whose first element is one and all other elements zero. The discounted sum of changes in the expectation of future returns, i.e. the discount rate component of the return, can thus be written as

$$\begin{aligned}\boldsymbol{\eta}_{DR,t+1} &\equiv (E_{t+1} - E_t)\sum_{j=1}^{\infty}\rho^j \Delta r_{t+1+j} \\ &= \mathbf{e1}'\sum_{j=1}^{\infty}\rho^j \boldsymbol{\Gamma}^j \boldsymbol{u}_{t+1} \\ &= \mathbf{e1}'\rho\boldsymbol{\Gamma}(\boldsymbol{I} - \rho\boldsymbol{\Gamma})^{-1}\boldsymbol{u}_{t+1} = \boldsymbol{\lambda}'\boldsymbol{u}_{t+1}\end{aligned} \tag{2.8}$$

with $\boldsymbol{\lambda}' = \boldsymbol{e1}'\rho\boldsymbol{\Gamma}(\boldsymbol{I} - \rho\boldsymbol{\Gamma})^{-1}$. The cash flow news component is then given by

$$\boldsymbol{\eta}_{CF,t+1} = (\mathbf{e1}' + \boldsymbol{\lambda}')\boldsymbol{u}_{t+1} \tag{2.9}$$

implied by equations (2.5) and (2.8) because $v_{r,t+1}$ can be picked out with $\boldsymbol{e1}'\boldsymbol{u}_{t+1}$.

I report the receptiveness of value and growth stocks to cashflow news and discount rate news as cashflow ("bad") beta and discount rate ("good") beta. Intertemporal asset pricing theory suggests that the former type of risk should be associated with a higher risk premium than the latter one (Merton, 1973). Intuitively, bad news about the market´s future cashflows reflect a decrease of wealth and hence lead to a fall in the value of the market but leave future investment opportunities unaffected. The value of the market portfolio could also decline because investors increase the discount rate applied to cashflows, which at the same time mirrors better future investment opportunities. Hence, receptiveness to discount rate news is less risky than sensitivity to cashflow news and therefore the terminology "bad" cashflow and "good" discount rate beta introduced by Campbell and Vuolteenaho (2004). The decomposition into cashflow and discount rate components could also be interpreted as decomposing the market return into its permanent and transitory parts according to Campbell and Vuolteenaho (2004). A stock price is the net present value of future discounted dividend growth. Hence a drop in dividend growth permanently affects the stock price and thus returns, whereas temporarily high discount rates could be offset by relatively low discount rates in the future. This interpretation underlies the assumption that dividend growth is unpredictable. However, appropriately defined macroeconomic variables do predict dividend growth in the long-run (Lettau and Ludvigson, 2005b; Hoffmann, 2006a), which leaves the impression that not all changes in dividends can be considered as permanent and hence this latter interpretation should be considered with caution.[7]

In order to obtain "bad" and "good" betas, I follow Campbell and Vuolteenaho (2004) and calculate cashflow betas from

$$\beta_{i,CF} = \frac{\operatorname{cov}(r_{i,t}, \eta_{CF,t})}{\operatorname{var}(r_{M,t} - E_t r_{M,t})} \tag{2.10}$$

Discount rate betas are obtained from

$$\beta_{i,DR} = \frac{\operatorname{cov}(r_{i,t}, -\eta_{DR,t})}{\operatorname{var}(r_{M,t} - E_t r_{M,t})} \tag{2.11}$$

where cov and var denote sample covariances and variances respectively, $r_{i,t}$ is the log excess return on stock *i* over the risk-free rate, $\eta_{CF,t}$, the estimated cashflow news term, $\eta_{DR,t}$,the estimated discount rate news component and $r_{M,t}$ - $E_t r_{M,t}$ the unexpected return on the market portfolio. The discount rate beta is here defined as the covariance of a stock return with lower than expected discount rates, i.e. "good" news. Note that these beta definitions differ from regression estimates. Betas are measured separately and conditioned on the variance of the unexpected market return not on the variance of the estimated news terms as would be the case in a regression. This definition implies that the sum of cashflow and discount rate betas equals the market beta, such that

$$\beta_{i,M} = \beta_{i,CF} + \beta_{i,DR} \tag{2.12}$$

[7] I thank Mathias Hoffmann for clarifying this point to me

2.3 State variables: Predictors of international stock returns

The success of the VAR in identifying cashflow and discount rate news components of a stock return relies on the choice of state variables, which have to explain stock market returns. Cashflow, i.e. dividend, news components are obtained as residual from the VAR. A major problem in applying this approach to international stock markets is to find a common set of predictive variables for national market returns.

In the previous chapter I have shown that the U.S. consumption-wealth ratio captures a common, transitory component in national stock markets. Short-run fluctuations of the ratio of consumption to aggregate wealth in the U.S. - *cay* – do not only predict time-varying excess returns on U.S. but also on foreign stock markets with considerable success. Hence, *cay* is a natural candidate as state variable.

Campbell and Vuolteenaho (2004) point out that the two-beta CAPM explanation for the U.S. value premium hinges on the use of the small-stock value spread, i.e. the difference in the logarithmic book-to-market value ratio on a small value portfolio and a small growth portfolio, as state variable in the market return decomposition into news components.[8] Campbell and Vuolteenaho (2004) motivate the use of the small-stock value spread by the inability of the Sharpe-Lintner CAPM to explain returns on small value and growth portfolios, which reflects that the value spread inherits information about systematic sources of risk not captured by the CAPM. However, the use of the small stock value spread as predictive variable has not remained uncontroversial (Liu and Zhang, 2006), in particular because the value spread's forecast ability seems to occur only in conjunction with other predictive variables.

[8] See the appendix to Campbell and Vuolteenho (2004) for further details on the construction of the small stock value spread

To address this concern and encouraged by the main results of the previous chapter I run univariate regressions of the G7 market excess returns on the U.S. value spread for which results are presented in table 2.1. If the basic logic of a common, transitory component in national stock markets pertains, then the U.S. value spread is also a potential explanatory variable for foreign stock market returns.

Table 2.1: Long-horizon regressions of foreign stock market excess returns on U.S. small stock value spread (sample period 1969Q4 – 2005Q4)

	h=1	h=4	h=8	H=12	h=16	h=20	h=24
CND	– 0.05 (–0.87)	**– 0.33** **(–2.18)**	– 0.33 (–1.49)	– 0.35 (–1.78)	– 0.24 (–1.31)	0.09 (0.33)	0.37 (1.16)
R^2	0.00	0.07	0.04	0.05	0.02	0.00	0.02
FRA	**– 0.13** **(–2.36)**	**– 0.45** **(–2.81)**	**– 0.61** **(–2.19)**	– 0.71 (–1.70)	– 0.44 (–0.77)	0.17 (0.24)	0.72 (1.02)
R^2	0.02	0.09	0.08	0.08	0.01	0.00	0.00
GER	**– 0.11** **(–2.01)**	– 0.36 (–1.89)	– 0.41 (–1.15)	– 0.44 (–0.80)	– 0.06 (–0.09)	0.69 (0.90)	1.19 (1.89)
R^2	0.02	0.07	0.05	0.04	0.00	0.04	0.13
ITA	– 0.10 (–1.94)	**– 0.44** **(–2.54)**	– 0.65 (–1.75)	– 0.74 (–1.71)	– 0.70 (–1.68)	– 0.21 (–0.41)	0.22 (0.31)
R^2	0.01	0.06	0.06	0.05	0.03	0.00	0.00
JPN	**– 0.16** **(–2.85)**	**– 0.46** **(–2.36)**	**– 0.74** **(–2.00)**	– 0.90 (–1.86)	– 0.99 (–1.88)	– 0.84 (–1.48)	– 0.65 (–0.82)
R^2	0.04	0.07	0.09	0.09	0.07	0.04	0.00
UK	**– 0.13** **(–3.07)**	**– 0.46** **(–4.57)**	**– 0.59** **(–3.10)**	**– 0.79** **(–3.11)**	**– 0.68** **(–2.24)**	– 0.40 (–0.92)	0.06 (0.16)
R^2	0.03	0.12	0.11	0.13	0.07	0.01	0.00
US	**– 0.08** **(–2.16)**	**– 0.32** **(–2.48)**	– 0.34 (–1.52)	– 0.43 (–1.28)	– 0.27 (–0.56)	0.17 (0.31)	0.68 (1.27)
R^2	0.02	0.10	0.06	0.06	0.01	0.00	0.05

Notes: This table displays OLS estimates from regressions of the form

$$r_{t+h}^{i,e} = \alpha_h + \beta_h vs_t + \varepsilon_{t+h}$$

with $r_{t+h}^{i,e}$ the log excess return on the MSCI stock index of country i at horizon t+h and vs_t is the U.S. small stock value spread. Newey-West (Newey and West, 1987) corrected t-statistics appear below the regressor estimates. R^2 reports the adjusted R^2. Bold faces highlight significant estimates.

The results in table 2.1. show that the value spread explains excess returns on the G7 stock markets at rather short, one and four quarter, horizon for all of the G7. It thus seems to be the ideal complement to *cay* which performs best at the business cycle frequency. The predictive power of the value spread highlights again the importance of the common, transitory component in international stock markets even at short horizons and the international evidence supports the argumentation line of Campbell and Vuolteenaho (2004).

Ideally, country-specific predictors of stock returns should complement the U.S. value spread and U.S. *cay* as forecast variables. However, evidence from the previous chapter leaves the impression that the forecast ability of a national financial variable as the dividend-price ratio varies across countries. Interest rate based predictors are not available for all of the countries under consideration because especially data on short-term interest rates is simply not available for the whole sample period from the first quarter of 1975 to the fourth quarter of 2005. More general, it is not clear if the same set of national variables exhibit predictive power for their national stock market returns at all as the example of national consumption-wealth ratios from the previous chapter has shown.

Because of these considerations, I push the idea of the common temporary component in stock markets to the extreme and use the U.S. price-earnings ratio as final stock market predictor for the European stock markets. The price-earnings ratio is defined as log of the S&P 500 stock index less a ten-year moving average of earnings on the S&P 500 as in Campbell and Shiller (1988b, 1998).[9] I also experimented with other U.S. variables that have been used as predictors of the U.S. stock market such as the relative treasury bill rate (Campbell, 1991; Hodrick, 1992), the rate of return on a 3-month treasury bill less a one-year backward moving average, as well as the term spread, the interest rate on a long-term U.S. government bond less the interest rate on a short-term note (Keim and Stambaugh,1986; Campbell, 1987; Fama and French, 1989). It turns out that the results only depend on the use

[9] Data is freely available on Robert J. Shiller's webpage http://www.econ.yale.edu/~shiller/data.htm

of the value spread and *cay*. All the other variables can be employed interchangeably. However, the price-earnings ratio predicts stock returns best at time horizons spanning several years (Campbell and Shiller, 1988b), such that this variable seems to be the ideal complement to the value spread (peak of forecast ability at one to four-quarters) and *cay* (predictive power peaks at three to four-year horizon), whereas interest rate based predictors seem to track the cyclical fluctuation in stock returns very much as *cay* does.

2.4 Data

Data on monthly and annual international value and growth returns is freely available on Kenneth French´s website.[10] Since I use *cay* as state variable, which is only observed at the quarterly frequency, I construct end-of-quarter return series from the monthly observations. Value and growth portfolios employed in this paper are book-to-market ratio sorted. The portfolios are formed at the end of December each year by sorting on their book-to-market value ratios and then value-weighted returns are calculated for the following 12 months. The value portfolios contain firms in the top 30 percent of a ratio and the growth portfolios contain firms in the bottom 30 percent. I use returns on value and growth portfolios of Belgium, France, Germany, Italy, The Netherlands, Spain and Switzerland to investigate the sensitivity of these returns with respect to cashflow and discount rate news on the respective market indexes. I employ the respective countries' Morgan Stanley Capital International (MSCI) indices as national stock market indexes which can be freely downloaded from http://www.mscibarra.com.

All indexes and hence returns are expressed in local currency, i.e. if I examine the cross-section of all of the European value and growth stock returns from a German perspective, then all returns are denominated in Deutschmark/Euro. If I take the perspective of a Swiss investor,

[10] http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/

then all returns are in Swiss Francs, etc. Excess returns are obtained with local short-term interest rates or in cases where these interest rates are not available the German three-month call money market rate. These data are from the IFS CD November 2006. U.S. *cay* is constructed as in chapter one, the U.S. price-earnings ratio is defined as log of the S&P 500 stock index less a 10-year moving average of log earnings on the S&P 500 as in Campbell and Shiller (1988b, 1998). The small stock value spread is constructed as described in the appendix to Campbell and Vuolteenaho (2004). It is the difference in the logarithmic book-to-market value ratio on a small value portfolio and a small growth portfolio measured at June of each year for which data can be downloaded from Kenneth French's website as well. Intrayear values (from July to May) are obtained by adding the cumulative log return on the small-book-to-market portfolio to, and subtracting the cumulative log return on the high-book-to-market portfolio from, the end-of-June value spread. The sample period runs from the first quarter of 1975 to the fourth quarter of 2005.

2.5 Empirical Evidence

This section is organized as follows. In the first subsection, I provide descriptive statistics that stress the presence of a premium on European value stocks. The second subsection gives details of the VAR characteristics when the Belgium, Dutch, German, French, Italian, Spanish and Swiss market returns are decomposed into cashflow and discount rate components. Then I assess the sensitivity of value and growth portfolio returns of the seven countries in question with respect to each of the estimated national market return's news components. Finally, this chapter examines the international cross-sectional dispersion in the value and growth portfolios to show that two-beta variants of national CAPMs explain average returns across European stock markets as suggested by basic asset pricing theory under the assumption of sufficiently integrated stock markets.

2.5.1 Descriptive statistics

Table 2.2 presents the annualized mean excess returns on the European value and growth portfolios from the perspective of a German investor.[11] Hence all returns are denominated in Deutschmark and Euro after 1998Q4. The German three-month call money market rate is used to obtain excess returns. When focusing on the mean excess returns, there is strong evidence for the value premium on these European stock markets. The only exception is Italy for which the growth portfolio promises a higher risk premium than the value portfolio. However, the sharpe ratios provide a less clear cut picture. In only four of the seven cases the sharpe ratio of the value portfolios is higher than that of the growth portfolio returns since the latter ones are less variable than the former ones.

Table 2.2: Descriptive statistics (1975Q1 – 2005Q4)

	mean excess returns (percentage points p.a.)	standard deviation (percentage points p.a.)	sharpe ratio
Value Belgium	13.08	50.22	0.26
Growth Belgium	7.32	41.98	0.17
Value France	11.04	59.56	0.19
Growth France	5.57	47.64	0.12
Value Germany	10.57	47.79	0.22
Growth Germany	5.15	48.66	0.11
Value Italy	3.50	66.70	0.05
Growth Italy	5.35	58.30	0.09
Value Netherlands	10.36	57.73	0.18
Growth Netherlands	8.01	38.78	0.21
Value Spain	4.38	65.97	0.07
Growth Spain	2.31	64.94	0.04
Value Switzerland	8.94	55.68	0.16
Growth Switzerland	6.35	40.63	0.16

Notes: This table presents the annualized mean excess returns on value and growth portfolios of Belgium, France, Germany, Italy, The Netherlands, Spain and Switzerland in percentage

[11] The mean returns, their standard deviations and the corresponding sharpe-ratios provide the same picture qualitatively when denominated in one of the other currencies.

points per annum as well as the respective standard deviation and sharpe ratio for the sample period 1975Q1 – 2005Q4. All returns are expressed in Deutschmark and Euro after 1998Q4. The German 3-month call money market rate is used to obtain excess returns.

2.5.2 VAR estimates

Table 2.3, panel A presents OLS coefficient estimates of a VAR consisting of the return on the German market portfolio, the small stock value spread, *vs*, short-term fluctuations in the U.S. consumption-wealth ratio, *cay*, and the U.S. logarithmic price-earnings ratio, $p - e$. A lag length of one quarter is suggested by Akaike, Schwartz and Hannan-Quinn information criteria. Each row of Panel A corresponds to one equation estimated in the VAR. T-statistics are displayed in parenthesis below the VAR estimates. R^2 denotes the adjusted R^2. All VAR estimates rely to some extent on the parameter ρ which should obey $\rho = \frac{1}{1+\exp(d-p)}$. I use sample means of the dividend yield to estimate ρ for each of the stock markets under consideration. All of the results remain qualitatively unaltered if I follow Campbell and Vuolteenaho (2004) who use an annual value of $\rho = 0.95$ and employ $\rho = 0.95^{1/4}$, since I deal with quarterly data, or if I let ρ vary around values between 0.9 and 0.99.

I discuss the results for Germany because they are very much representative for all of the other countries. Thus I take the freedom to present only the market return equations of the other VARs in table 2.4. Focusing on the return equation in the first row, the state variables predict about four percent of the variation in the one-quarter excess return on the German market portfolio. The value spread is marginally insignificant while U.S. *cay* marginally significantly explains the German market return at the 95 percent confidence level. This finding is noteworthy, as Hamburg at al. (2007) find the German consumption-wealth ratio to predict macroeconomic variables such as the unemployment rate rather than German stock market returns. Interestingly, the coefficient of the price-earnings-ratio is incorrectly signed and not statistically distinguishable from zero.

Table 2.3: VAR characteristics (German market return)

Panel A: VAR estimates					
	r_t^M	vs_t	cay_t	$(p-e)_t$	R^2
r_{t+1}^M	-0.07 (-0.72)	-0.13 (-1.83)	**2.56** **(1.97)**	0.05 (1.50)	0.04
vs_{t+1}	-0.12 (-1.53)	**0.77** **(13.13)**	0.52 (0.47)	**0.07** **(2.18)**	0.97
cay_{t+1}	**- 0.02** **(-3.67)**	-0.00 (-0.88)	**0.66** **(9.23)**	-0.00 (-1.84)	0.83
$(p-e)_{t+1}$	-0.06 (-1.02)	-0.06 (-1.44)	1.18 (1.48)	**0.96** **(17.46)**	0.97
Panel B: Variance share of news terms and correlation					
var_{NCF} :0.95	var_{NDR} :0.28	$-2\text{cov}(\eta_{DR},\eta_{CF})$: -0.23		$\rho_{NCF,NDR}$:	0.23

Notes: Panel A of this table displays the estimated VAR coefficients. Newey-West (Newey and West, 1987) corrected t-statistics appear in parenthesis below the estimates. The lag length of the VAR is one quarter. r_{t+1}^M denotes the natural logarithm of the excess return on the German market portfolio, vs_t is the U.S. small stock value spread, cay_{t+1} is the residual of the cointegrating relation between U.S. consumption and aggregate wealth, $(p-e)_t$ is the U.S. price-earnings ratio constructed as in Campbell and Shiller (1988b) . R^2 is the adjusted R^2

Panel B gives the shares of the market return variation explained by the variation in the two news series, $\text{var}_{NCF}, \text{var}_{NDR}$, and the covariance between the news components, $-2\text{cov}(\eta_{DR},\eta_{CF})$. It also presents the correlation coefficient between cashflow and discount rate components of the German market return, $\rho_{NCF,NDR}$. Bold faces highlight significant estimates

Panel B gives the share of the variance of the unexpected market return explained by cashflow and discount rate news respectively as well as the share captured by the covariance between the two news series. The cashflow news component clearly dominates variation in the German market return. This result is in stark contrast to the findings of Campbell (1991) and Campbell and Vuolteenaho (2004) that discount rate news predominantly cause variation in the U.S. market return in post-war data. However, the news terms are almost uncorrelated with each other. The correlation coefficient between the news series is about 0.23.

Table 2.4. presents the return equation from the VARs of the other six countries. The results are qualitatively similar to the VAR characteristics obtained for the German market return. Furthermore, the correlation between the news series varies between -0.12 for Italy and Spain and 0.22 for Belgium and Switzerland.

Table 2.4: Return forecasting equation from countries' VARs

	r_t^M	vs_t	cay_t	$(p-e)_t$	R^2
Belgium					
r_{t+1}^M	-0.17 (-1.84)	$\mathbf{-0.13}$ $\mathbf{(-2.08)}$	1.37 (1.35)	0.04 (1.37)	0.03
France					
r_{t+1}^M	-0.07 (-0.71)	-0.13 (-1.83)	2.22 (1.75)	0.05 (1.36)	0.02
Italy					
r_{t+1}^M	-0.02 (-0.19)	-0.08 (-1.01)	1.90 (1.32)	0.04 (0.91)	0.00
Netherlands					
r_{t+1}^M	-0.06 (-0.63)	-0.10 (-1.68)	1.37 (1.50)	0.02 (0.74)	0.02
Spain					
r_{t+1}^M	-0.08 (-0.84)	-0.09 (-1.16)	1.58 (1.11)	0.04 (1.04)	0.00
Switzerland					
r_{t+1}^M	-0.11 (-1.20)	-0.10 (-1.80)	2.20 (1.84)	0.05 (1.67)	0.03

Notes: This table presents the return equations from the VARs of Berlgium, France, Italy, the Netherlands, Spain and Switzerland to decompose the respective market excess returns into their cashflow and discount rate components. Further details are given in the notes to table 2.3

2.5.3 Bad and good betas

Cashflow and discount rate news components of the German and the other countries' market portfolio are almost uncorrelated with each other. This observation conveys the notion that different types of stocks could react differently to cashflow and discount rate news. Furthermore, intertemporal asset pricing theory suggests that receptiveness to the market portfolio´s cashflow news should be compensated with a higher risk premium than sensitivity to discount rate news (Merton, 1973). For the U.S., Campbell and Vuolteenaho (2004), Cohn et al. (2003) and Campbell et al. (2005) show that value stocks promise higher returns than growth stocks because their Sharpe-Lintner CAPM market betas are dominated by the cashflow variety whereas growth stocks' market betas are primarily driven by discount rate news.

Table 2.5 presents bad and good betas of value and growth portfolios of the seven European countries under consideration with respect to the news series obtained for the German market return. Again, I focus on Germany because the bad and good beta estimates of the other countries basically provide the same message.

The picture that emerges from the results shown in table 2.5 is that high returns on value portfolios are associated with relatively high cashflow betas from the perspective of a German investor. Notice that the definitions of the betas imply that their sum should equal the corresponding market beta. Then the market betas of most of the value portfolio returns are higher than the growth portfolio market betas anyway. But this finding is caused by their high cashflow betas compared to growth portfolios.

Table 2.5: Bad and good betas from German perspective

	β_{CF}	β_{DR}
Value BEL	0.49	0.05
Growth BEL	0.36	0.09
Value FRA	0.62	0.10
Growth FRA	0.43	0.17
Value GER	0.77	0.08
Growth GER	0.69	0.21
Value ITA	0.60	0.10
Growth ITA	0.44	0.17
Value NL	0.70	0.11
Growth NL	0.39	0.11
Value ESP	0.42	0.06
Growth ESP	0.41	0.16
Value CH	0.67	0.20
Growth CH	0.34	0.13

Notes: This table presents cashflow and discount rate beta estimates of value and growth stocks of Austria, Belgium, France, Germany, Ireland, Italy, Netherlands, Spain, Switzerland and the United Kingdom conditional of cashflow and discount rate components of the excess return on the German market portfolio.
Betas are calculated from

Cashflow beta: $$\beta_{i,CF} = \frac{\mathrm{cov}(r_{i,t}, \eta_{CF})}{\mathrm{var}(r_{M,t} - E_{t-1}(r_{M,t}))}$$

Discount rate beta: $$\beta_{i,DR} = \frac{\mathrm{cov}(r_{i,t}, -\eta_{DR})}{\mathrm{var}(r_{M,t} - E_{t-1}(r_{M,t}))}$$

where cov and var denote sample covariances and variances.
The cashflow component is abbreviated with η_{CF}, the discount rate news component with η_{DR}, $r_{i,t}$ denotes the individual value or growth stock excess return and $r_{M,t} - E_{t-1}(r_{M,t})$ represents the unexpected market return. The discount rate beta is here defined as the covariance of a stock return with lower than expected discount rates.

The only exceptions are the German value and growth portfolios which mirror a salient feature of the data: The market betas of national growth stocks are higher than their value stocks' counterparts. However, the cashflow news driven component in value stocks' market betas is substantially higher than in the market beta of growth stocks. Hence, the evidence provided in table 2.5 corroborates Campbell and Vuolteenaho (2004), Cohn et al. (2003) and Campbell et al. (2005) in European data. High bad betas relative to discount rate betas are associated with relatively high average returns. This conclusion is stressed by figures 2.1 to 2.7. These figures display mean excess returns on the European national value and growth portfolios denominated in the respective local currency on the horizontal axis compared to the ratio of their cashflow beta, obtained for each of the seven market return news components, with the respective market betas on the vertical axis.

Figures 2.1. to 2.7: Average excess returns on value and growth portfolios relative to the ratio of cashflow to respective market beta

Figure 2.1: Belgium

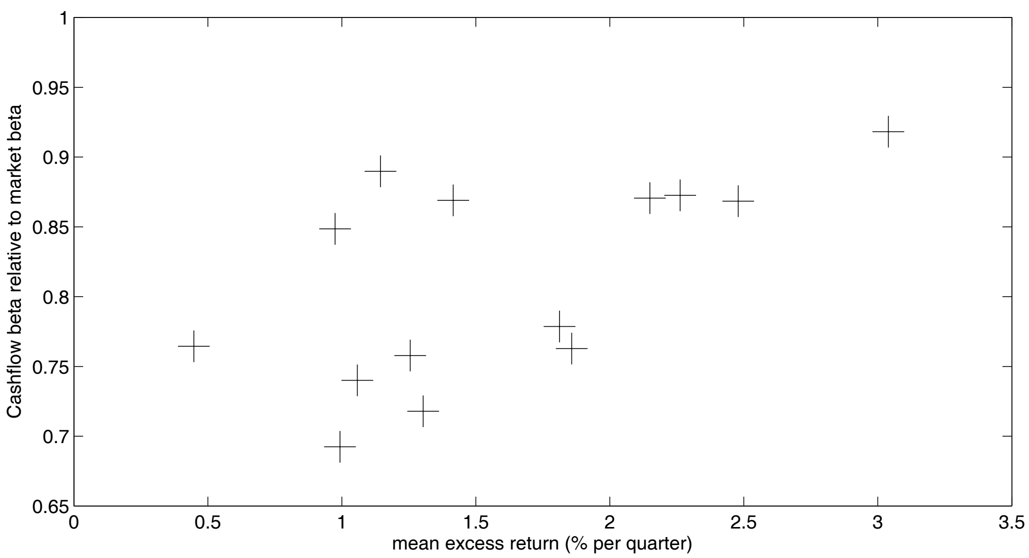


Figure 2.2: France

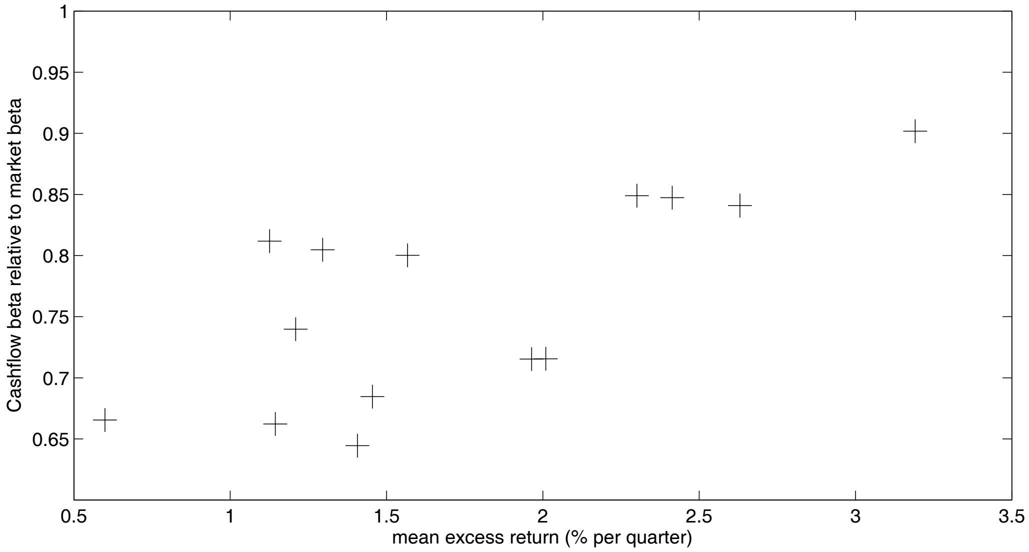


Figure 2.3: Germany

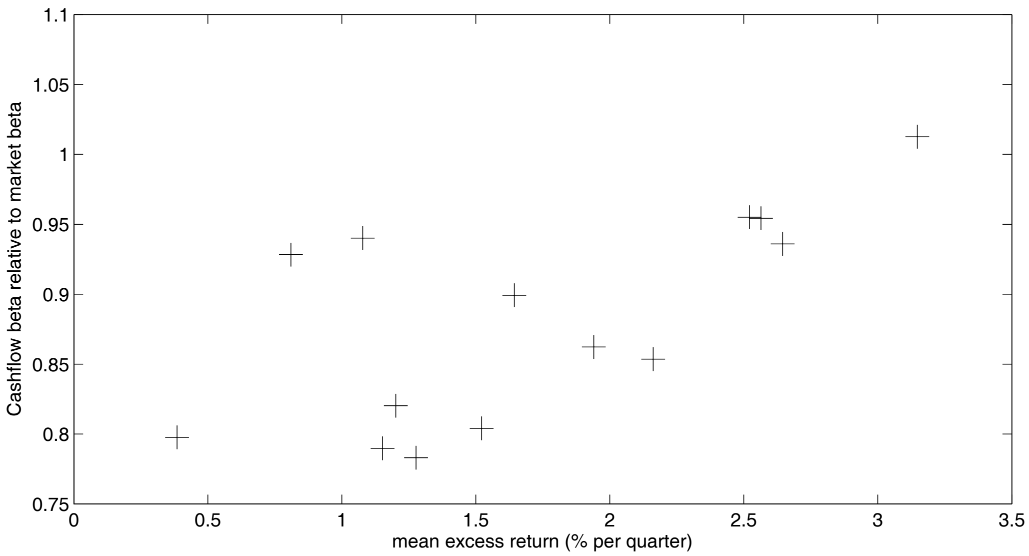


Figure 2.4: Italy

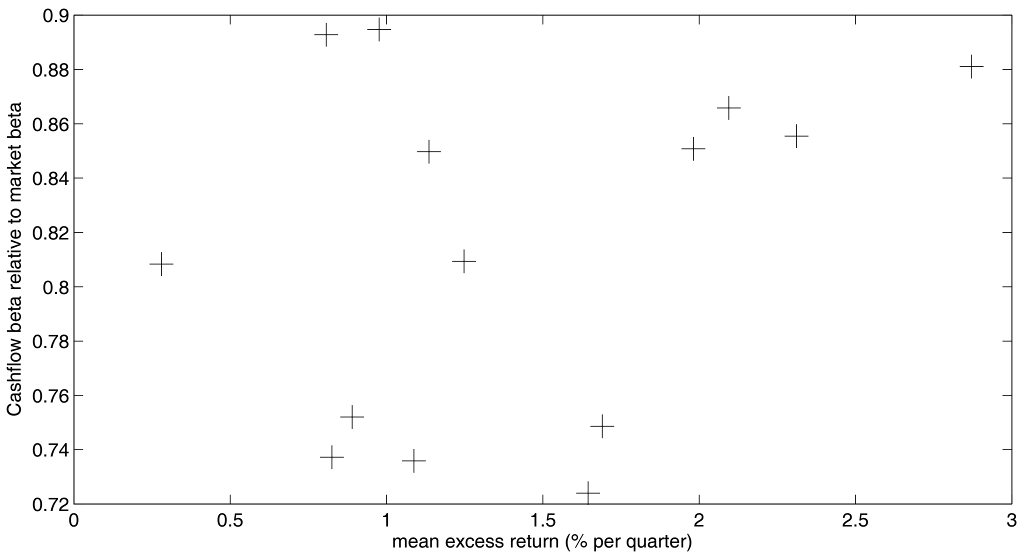


Figure 2.5: Netherlands

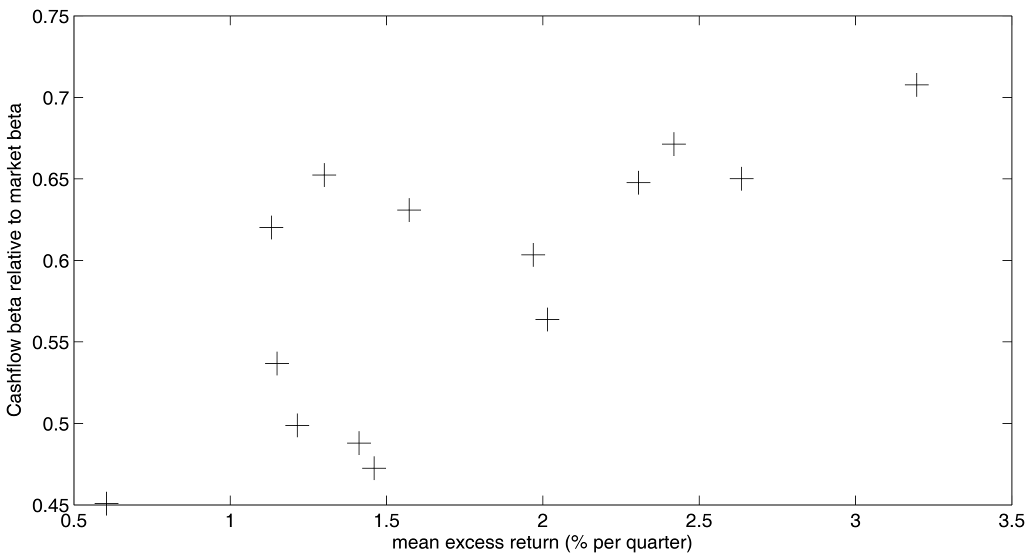


Figure 2.6: Spain

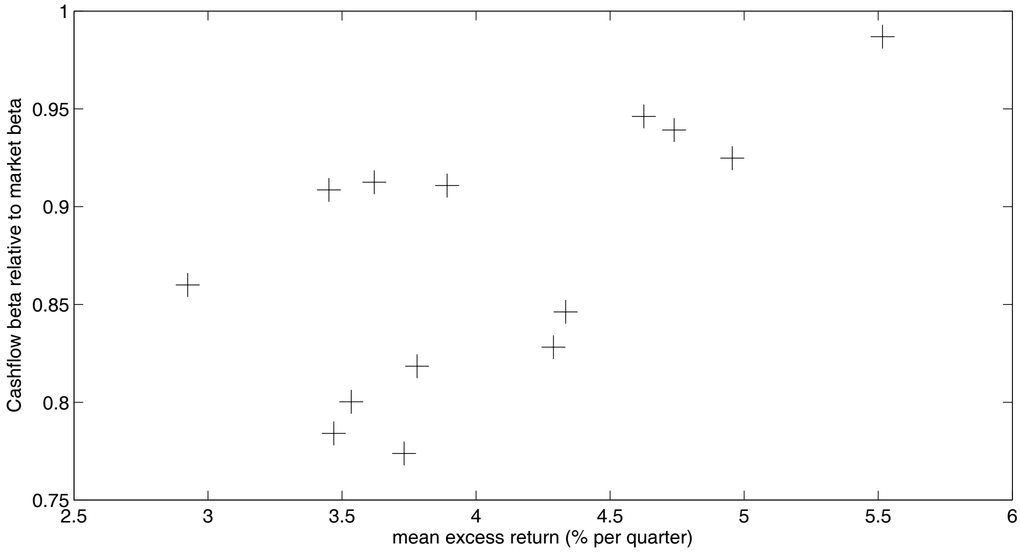
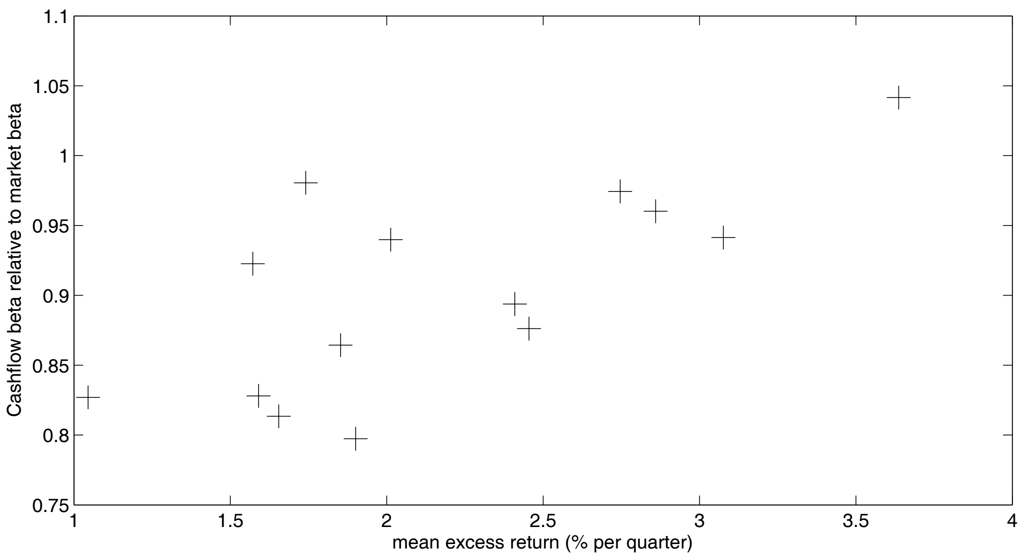


Figure 2.7: Switzerland



The relationship between average returns and the size of the cashflow component in the market beta is positive. The figures thus provide a visual impression of the conclusion drawn before.

2.5.4 The international cross-section of returns

The focus of this subsection is closely related to the implications of the basic pricing equation (2.1) and the observation in table 2.1 that returns on value and growth portfolios vary internationally. The basic pricing equation states that one (national) discount factor should be applicable to price any (international) asset. Hence, international differences in stock returns should be captured by a national asset pricing model. So, the question addressed in this section is basically: Does the national two-beta CAPM provide a rationale for differences in average stock returns across countries? The focus is thus more on the international cross-sectional dimension of stock returns than on the distinction between value and growth stocks, although the sorting of stocks with respect to their book-to-market ratios ensures large spreads in the European stock returns and is hence also important in this context.

The cross-sectional implication of the basic pricing equation can be seen by a simple rearrangement of equation (2.1), exploiting that $E(XY) = E(X)E(Y) + \text{cov}(X,Y)$ which gives

$$E_t(R^i_{t+1}) = \frac{1}{E_t(M_{t+1})} + \left(\frac{\text{cov}(R^i_{t+1}, M_{t+1})}{\text{var}(M_{t+1})}\right)\left(-\frac{\text{var}(M_{t+1})}{E_t(M_{t+1})}\right) \tag{2.13}$$

which can be summarized to

$$E_t(R^i_{t+1}) - R^f_t = \beta^i_M \lambda_M \tag{2.14}$$

with $R^f_t = \dfrac{1}{E_t(M_{t+1})}$, the risk-free rate, $\beta^i_M = \dfrac{\text{cov}(R^i_{t+1}, M_{t+1})}{\text{var}(M_{t+1})}$, the regression coefficient of excess return i on the discount factor, representing the asset-specific quantity of systematic risk, and $\lambda_M = -\dfrac{\text{var}(M_{t+1})}{E_t(M_{t+1})}$ interpreted as the price of risk.

As the excess return on the market portfolio is the only source of systematic risk in a Sharpe (1964) and Lintner (1965) CAPM, (2.14) collapses to

$$E_t(R^i_{t+1}) - R^f_t = \hat{\beta}^i_{R^M} \lambda_{R^M} \tag{2.15}$$

in which λ_{R^M} is the price of the market risk and $\hat{\beta}^i_{R^M}$ is the estimated asset-specific exposure to the market portfolio. High excess returns should thus be associated with high market betas as λ_{R^M} is the same for all assets.

In the case of the two-beta CAPM, the cross-sectional regression (2.15) looks as follows

$$E_t(R^i_{t+1}) - R^f_t = \hat{\beta}^i_{CF}\lambda_{CF} + \hat{\beta}^i_{DR}\lambda_{DR} \tag{2.16}$$

where the betas represent the sensitivity of the stock excess returns on the cashflow and discount rate news terms of the respective market returns as given in table (2.3) for Germany. The pure market return betas for the simple CAPM are obtained from time series regressions of the value and growth stock returns on the respective market excess returns.

I follow Fama and MacBeth (1973) and estimate a cross-sectional regression of the value and growth portfolio returns on the estimated betas at each point in time. The risk prices are then averages of the estimated risk price series.

The cross-sectional results for the Sharpe (1964) and Lintner (1965) CAPM are given in table 2.6 with Shanken (1992) corrected t-statistics in parenthesis since the betas are generated regressors. R^2 is the cross-sectional R^2. Mean absolute pricing errors (mape) and mean squared pricing errors (mspe) are given in percentage points per quarter.

The results displayed in table 2.6 leave the impression that national CAPMs perform poorly in explaining international cross-sectional differences in stock returns. The only exception is Belgium for which the CAPM works fairly well. However, for the remaining countries the national CAPMs capture virtually none on the cross-sectional dispersion in European value and growth portfolios returns.

Table 2.6: Cross-sectional regressions (CAPM)

	λ_{R^M}	R^2	mspe	mape
Germany				
Sharpe-Lintner CAPM	0.65 (0.43)	0.01	0.59	0.66
Belgium				
Sharpe-Lintner CAPM	3.50 (1.67)	0.36	0.30	0.46
France				
Sharpe-Lintner CAPM	0.33 (0.21)	0.01	0.46	0.58
Italy				
Sharpe-Lintner CAPM	0.91 (−0.59)	0.08	0.43	0.54
Netherlands				
Sharpe-Lintner CAPM	1.77 (0.87)	0.10	0.42	0.50
Spain				
Sharpe-Lintner CAPM	−1.57 (−0.81)	0.18	0.38	0.52
Switzerland				
Sharpe-Lintner CAPM	0.56 (0.29)	0.01	0.46	0.57

Notes: This table presents results from cross-sectional Fama-MacBeth regressions (Fama and MacBeth, 1973) for national CAPMs when confronted with value and growth stock returns of Berlgium, France, Germany, Italy, the Netherlands, Spain and Switzerland. R^2 is the cross-sectional R^2 used in Jagannathan and Wang (1996). The mean absolute (mape) and mean squared pricing errors (mspe) are reported in percentage points per quarter. The sample spans the period from 1975Q1 to 2005Q4.

This picture dramatically changes once we consider the market return's cashflow and discount rate risk separately. Table 2.7 presents the cross-sectional regression estimates from the two-beta variants of the CAPM.

Table 2.7: Cross-sectional regressions (two-beta CAPM)

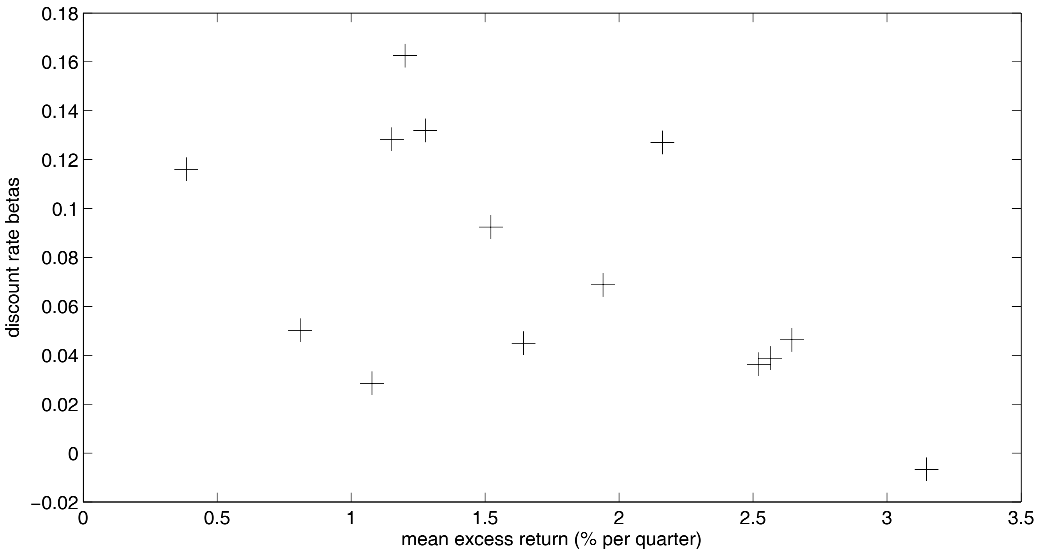
	λ_{NCF}	λ_{NDR}	R^2	mspe	mape
Germany					
Two-beta CAPM	2.35 (1.35)	**−8.78** **(−2.03)**	0.48	0.31	0.42
Belgium					
Two-beta CAPM	2.87 (1.76)	−2.80 (−0.60)	0.53	0.22	0.39
France					
Two-beta CAPM	2.34 (1.45)	**−7.59** **(−2.06)**	0.54	0.21	0.38
Italy					
Two-beta CAPM	0.10 (0.05)	−12.29 (−1.48)	0.28	0.33	0.51
Netherlands					
Two-beta CAPM	**4.21** **(2.11)**	−3.97 (−1.44)	0.62	0.18	0.35
Spain					
Two-beta CAPM	0.51 (0.25)	**−18.74** **(−2.44)**	0.66	0.20	0.36
Switzerland					
Two-beta CAPM	1.76 (1.09)	**7.13** **(−1.99)**	0.58	0.20	0.38

Notes: This table presents results from cross-sectional Fama-MacBeth regressions (Fama and MacBeth, 1973) for two-beta versions of national CAPMs when confronted with value and growth stock returns of Belgium, France, Germany, Italy, the Netherlands, Spain and Switzerland. R^2 is the cross-sectional R^2 used in Jagannathan and Wang (1996). The mean absolute (mape) and mean squared pricing errors (mspe) are reported in percentage points per quarter. The sample spans the period from 1975Q1 to 2005Q4.

The fit of this model is by far better and produces statistically significant risk prices. The R^2 statistics range from 0.28 for Italy to 0.66 for Spain. Moreover, mean squared and mean absolute pricing errors are substantially lower when compared with the ones obtained from the Sharpe-Lintner CAPMs. Thus the distinction between the market's cashflow and discount rate risks considerably improves the performance of the CAPM.

For five of the seven countries, differences in national discount rate betas account for the cross-sectional dispersion in European stock returns. The negative risk prices simply reflect the fact that discount rates are defined as better news about discount rates than expected. Hence there is a negative relation between average returns and discount rate betas for the respective countries. Figure 2.8 shows this relationship for Germany.

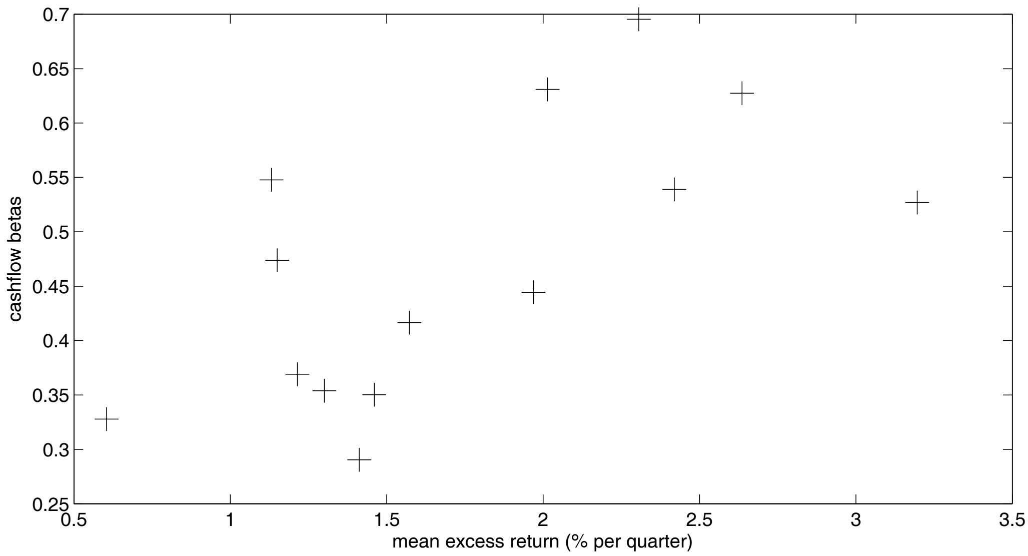
Figure 2.8: Average returns vs. German discount rate betas



The higher the mean excess return (horizontal axis), the lower the beta to good news about the German stock market return (vertical axis).

This picture is quite different when we regard the Netherlands. From the perspective of a Dutch investor differences in the sensitivity to the Dutch market's cashflow news explain international differences in stock returns. Figure 2.9 thus displays the positive relation between Dutch cashflow betas (vertical axis) and average returns on the European value and growth portfolios (horizontal axis).

Figure 2.9: Average returns vs. Dutch cashflow betas



High average returns are hence the outcome of relatively high cashflow betas from the perception of a Dutch investor.

It is interesting that two-beta versions of the CAPM work so well in capturing cross-sectional differences in European stock returns. However, there are fundamental differences in the systematic sources of risk that explain average returns from a national investor's perspective.

Remember that discount rate news are directly estimated in the VAR. Hence, all of the information about stock market returns inherited in the state variables is shifted to the discount rate components. The state variables are motivated under the assumption that their predictive power is the mirror image of a common, temporary component in national stock markets. The very fact that the spreads in discount rate betas of the German, French, Italian, Spanish and Swiss market return help to explain the cross-sectional differences between the European value and growth portfolio returns suggests that these national investors regard differences in the sensitivity to the common stock market component as the decisive feature in order to judge the riskiness of stock returns. The residual information, i.e. cashflow news, plays a negligible role for these investors.

For a Benelux investor, the opposite reasoning seems to apply. Receptiveness to national market's cashflow news capture cross-sectional differences in average returns.

2.6 Conclusions

Employing the framework of Campbell (1991) and Campbell and Vuolteenaho (2004), I show that high average returns on European value portfolio returns can be reconciled with the two-beta variant of the Sharpe-Lintner CAPM from a national investor's perspective. High returns on value stocks are associated with relatively high cashflow betas compared to the respective discount rate betas. This finding is a salient feature of the data irrespective if one takes the stance of a Belgium, French, German, Italian, Dutch, Spanish or Swiss investor. This finding is in line with studies by Campbell and Vuolteenaho (2004), Cohen et al. (2003) and Campbell at el. (2005) with regard to book-to-market ratio and size sorted portfolios in the U.S. In addition, two-beta versions of national CAPMs capture the cross-sectional dispersion in European value and growth stock returns. This chapter thus provides empirical evidence for the implication that a national asset pricing model should explain cross-sectional dispersion in any asset as suggested by basic asset pricing theory if national capital markets are sufficiently integrated.

Chapter 3

Consumption growth, uncovered equity parity and the cross-section of returns on foreign currencies

3.1 Introduction

At the end of August 2006 the U.S. dollar has appreciated against the British Pound by 1.8 percent and at the same time depreciated against the Swiss Franc by 0.12 percent compared to end of July 2006. Why are there such differences in changes of U.S. dollar exchange rates?

We know since Mussa (1979) that exchange rates essentially behave like asset prices. Hence, exchange rate changes can be interpreted as returns on holding foreign currency and asset pricing theory should therefore apply to explain cross-sectional differences in exchange rate changes. In this chapter I show that the consumption-based capital asset pricing model (CCAPM) canonized in Breeden (1979) and Breeden et al. (1989) explains up to 50 percent of the cross-sectional variation in quarterly returns on foreign currencies in the post Bretton-Woods period. This finding hinges on the use of stock market return differentials as conditioning instrument and is obtained from the perspective of a U.S. investor for U.S. dollar exchange rates of a set of 15 developed countries.

This chapter is heavily influenced by the finding that consumption growth risk is priced in risk premia on foreign currencies in Lustig and Verdelhan (2006). They show that various versions of the CCAPM do not only allow to explain the cross-section of domestic stock returns (Lettau and Ludvigson, 2001b; Parker and Julliard, 2005; Yogo, 2006) but also of currency excess returns, i.e. the excess return on investment in foreign bonds. Key to success is the formation of annually rebalanced currency portfolios for a wide cross-section of countries sorted with respect to interest rate differentials. Lustig and Verdelhan present evidence that consumption-based models capture the cross-section of average excess returns on these currency portfolios, especially the model proposed by Yogo (2006). Market return based models, such as the Sharpe (1964) and Lintner (1965) CAPM or the Fama and French three-factor model (Fama and French, 1993), perform poorly in this respect.

The main findings of Lustig and Verdelhan suggest that the empirical failure of the uncovered interest rate parity condition (UIP) reflects basic asset pricing theory. Currencies from countries with high interest rates relative to the U.S. are perceived as relatively risky asset by the U.S. investor. Hence, those currencies have to offer relatively high returns, i.e. the foreign currency has to appreciate, while UIP predicts a foreign currency depreciation.

In addition, I build on a recent strand of literature that theoretically and empirically emphasizes the tight link between cross-border capital flows, relative stock market returns and exchange rate changes among developed economies. This literature, briefly summarized in the subsequence, conveys the notion of an uncovered equity return parity condition (UEP).

Capital flows between stock markets have increased strongly in the past two decades (Tesar and Werner, 1995) and have gained considerable importance for exchange rates. Siourounis (2004) shows that flows between equity markets explain exchange rate movements among the five major currencies while bond flows do not when both kinds of flows are considered simultaneously.

Based on a model with incomplete financial markets, Hau and Rey (2006) argue that portfolio rebalancing considerations lead to capital flows between equity markets that contemporaneously affect exchange rates. Investment into foreign equity exposes the investor to both equity return and currency return risk. Since capital markets are incomplete, exchange rate risk cannot be hedged. A higher foreign stock market return in local currency compared to the home stock market biases the investor´s portfolio return to foreign stocks and exposes her to a higher currency risk. If this risk exposure is high enough, it is optimal for the investor to rebalance her stock portfolio and sell foreign stocks when they offer higher returns than the home stock market. Portfolio rebalancing thus creates flows between stock markets and hence exchange rate movements in the way observed in the data, i.e. high stock returns relative to the home country are associated with a foreign currency depreciation (Brooks et al., 2001; Hau and Rey, 2004).

Cappiello and De Santis (2005) argue that an arbitrage relation between expected returns on exchange rates and expected returns on stock markets implies UEP. Taking account of country-specific risk premia on stocks they find stock return differentials to be good predictors of the sign of expected exchange rate changes in- and out-of-sample. This latter result could be driven by the fact that the response of exchange rates to shocks in relative stock market returns persists one or two periods after the shock has occurred (Hau and Rey, 2004).

I exploit these findings and use lagged interest rate together with lagged stock return differentials as characteristics to form portfolios of currencies for the developed countries that are taken into consideration by Lustig and Verdelhan (2006). Relative stock returns and interest rate differentials are treated as instrumental variables for expected exchange rate changes. We know since Meese and Rogoff (1983) that our ability to forecast quarterly exchange rate changes is limited.[12] The best we can do is to use previous period's interest rate differentials as suggested by UIP or, as I argue in this chapter, employ lagged stock return differentials. The sorting of currencies into portfolios serves to ensure that there are large spreads in currency returns at quarterly frequencies. The portfolio formation in Lustig and Verdelhan (2006) for annual excess returns on foreign bonds has the same purpose.

The simple CCAPM explains roughly 50 percent of the cross-sectional dispersion in average, quarterly returns on currency portfolios sorted by interest rate and stock return differentials.

[12] However, Mark (1995) and Hoffmann and MacDonald (2006) show that fundamentals (relative liquidities or real interest rate differentials) capture the time series variation in nominal and, respectively, real exchange rate changes in the long-run. Starting with Evans and Lyons (2002), a microfounded literature shows that exchange rate changes are predicted by order flows at daily frequencies.

The estimated price of consumption growth risk is about two percentage points per annum and we cannot reject the hypothesis that all pricing errors are zero. Both interest rate differentials and stock return differentials seem to provide information about future exchange rate changes according to the portfolio characteristics.

Moreover, evidence presented in this chapter leaves the impression that -- judged by their sensitivity to consumption growth conditional on time variation in relative stock returns -- currencies from countries with high past stock returns seem to be less risky than their counterparts from past low stock return differential countries. This finding thus provides an additional rationale for using stock return differentials as instrumental variable in order to explain the cross-section of exchange rate changes.

Finally, I follow Cochrane (1996) and assess the conditional implications of the CCAPM by considering currency returns scaled with the respective stock return differential as test assets on a country-by-country basis. The conditional estimate of the price of consumption growth risk is statistically significantly different from zero and close to the values presented by Lustig and Verdelhan (2006). Furthermore, the fit of the CCAPM is about as well as in the case of currency portfolios. This latter result conveys the notion that prior portfolio formation does not seem to be necessary to explain the cross-section of returns on foreign currencies in a consumption-based asset pricing framework in a data set of countries for which UEP holds.

The remainder of the chapter is organized as follows. Section 3.2 describes the theoretical asset pricing framework while section 3.3 focuses on the empirical evidence for currency portfolios. In section 3.4, I assess empirically the relation between stock return differentials, individual foreign currency returns and U.S. consumption growth. Section 3.5 reports the performance of the CCAPM when confronted with managed returns on foreign currencies. Section 3.6 concludes.

3.2 The CCAPM and returns on foreign currencies

The first-order conditions of an investor´s utility maximization problem give the basic consumption-based asset pricing equation

$$P_t^i = E_t\left[\beta \frac{u'(C_{t+1})}{u'(C_t)} X_{t+1}^i\right] \tag{3.1}$$

with P_t^i the price of asset *i* and X_{t+1}^i the corresponding payoff. E_t is the expectation operator conditional on time t information, β is the subjective discount factor and $u'(C)$ denotes the first derivative of the utility function $u(C)$. C_t represents consumption at time t.

Equivalently, gross returns, $\frac{X_{t+1}^i}{P_t^i}$, obey

$$1 = E_t(M_{t+1} R_{t+1}^i) \tag{3.2}$$

with $M_{t+1} \equiv \beta \frac{u'(C_{t+1})}{u'(C_t)}$ the stochastic discount factor and R_{t+1}^i the gross return on asset or portfolio *i*. In words, any asset return should be discounted with the same stochastic discount factor (SDF) which is directly related to the intertemporal marginal rate of substitution of consumption.

We know since Mussa (1979) that exchange rates essentially behave like asset prices. I thus interpret a change in exchange rates as return on holding foreign currency. The focal point of this chapter is pure currency risk and not excess returns that accrue from the investment in e.g. foreign bonds as in Bekaert and Hodrick (1992) or Lustig and Verdelhan (2006).

Hence, equation (3.2) becomes

$$1 = E_t(M_{t+1} \frac{S_{t+1}^i}{S_t^i}) \tag{3.3}$$

where S_{t+1}^i represents a change in the exchange rate defined as foreign currency *i* over U.S. dollar. From the perspective of a U.S. investor, a foreign currency depreciation corresponds to a low or negative return on the foreign currency while a foreign currency appreciation implies a high or positive return.

I assume that the representative investor maximizes a power utility function. The first-order conditions imply

$$M_{t+1} = \beta \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma} \tag{3.4}$$

where γ represents the coefficient of relative risk aversion. A log-linear version of M_{t+1} obeys

$$m_{t+1} = \alpha - b_{\Delta c} \Delta c_{t+1} \tag{3.5}$$

with $\alpha = \log(\beta)$ and $b_{\Delta c} = \gamma$.

Hence, equation (3.3) reduces to

$$0 = E_t(m_{t+1} \Delta s_{t+1}) \tag{3.6}$$

with Δs_{t+1}^i the log exchange rate change of country *i*. Since consumption has been deflated, I define the nominal U.S. dollar exchange rate deflated with realized U.S. inflation as return on foreign currency, i.e.

$$r_{t+1}^i = \Delta s_{t+1}^i - \pi_{t+1} \tag{3.7}$$

with r_{t+1}^i the deflated return on currency *i* and π_{t+1} realized U.S. inflation, such that equation (3.3) becomes

$$0 = E_t(m_{t+1} r_{t+1}^i) \tag{3.8}$$

In the empirical part of this chapter, I regard either returns on portfolios of currencies or currency returns scaled with an instrumental variable. Hence, I basically examine

$$0 = E(m_{t+1} z_t r_{t+1}^i) \tag{3.9}$$

under the assumption that the instrument (or portfolio characteristic) *z* observed at time t contains sufficient information about the investor's information set concerning expected exchange rate changes realized in t+1. Then equation (3.9) can be interpreted as the pricing equation of a managed portfolio of foreign currency returns (Cochrane, 1996). Of course, z could also be a vector of instruments.

3.3 Currency portfolios

Even though the implications of UIP do not hold in the data, interest rate differentials mirror expected exchange rate changes (Lustig and Verdelhan, 2006). Hau and Rey (2004) and Cappiello and De Santis (2005) suggest that stock return differentials are a good statistic for expected currency returns at quarterly or higher frequency. I take both of these findings into account and form portfolios of currencies with respect to one-quarter lagged interest rate and stock return differentials to ensure large spreads in currency returns. This sorting would be inappropriate if the two characteristics were highly correlated. However, the correlations between interest rate and stock return differentials vary between -0.24 and -0.02.

The sample spans the period from 1974Q1 to 2005Q4 and takes the following countries under consideration: Australia, Austria, Belgium, Canada, Denmark, France, Germany, Greece, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland and the United Kingdom.

I obtain data on short-term interest rates from the IFS CD November 2006. Quarterly data on stock returns is calculated from end-of-period returns on the Morgan Stanley Capital International (MSCI) stock index of the respective countries in local currency.[13] I calculate quarterly currency returns from end-of-quarter MSCI stock returns in local currency at time t and the respective stock returns in U.S.-dollars. Positive foreign currency returns are thus associated with a depreciation of the U.S. dollar. Returns on foreign currencies are deflated with the change in the U.S. consumer price index (CPI) published in the IFS CD November 2006.

[13] These stock indexes are freely available on www.mscibarra.com .

3.3.1 Portfolio characteristics

First, I form three portfolios sorted by lagged interest rate differentials. The first portfolio consists of the 30% of countries with highest interest rate differentials, the second one of the 40% of countries with middle interest rate differential and portfolio three contains the remaining 30% of countries with lowest interest rate differential via-à-vis the U.S. Then I distinguish between high and low stock return differentials compared to the U.S. within the interest rate differential sorted portfolios, such that I obtain six portfolios which are rebalanced every quarter.[14]

Table 3.1 presents average excess returns and the Sharpe-ratio of all six portfolios. Note that portfolio P11 contains the countries with highest interest rate differential compared to the U.S. (the first "1") and among those the countries with high stock return differentials (the second "1"). Portfolio P12 consists of countries with highest interest rate differential relative to the U.S. and among those of the countries with low stock return differential. Accordingly, P32 is the portfolio of exchange rates from countries with lowest interest rate differential and relatively low stock returns.

Table 3.1: Portfolio characteristics

	P11	P12	P21	P22	P31	P32
Mean excess return	0.82	1.38	0.03	-1.83	-3.33	-1.87
Sharpe-ratio	0.04	0.07	0.00	-0.09	-0.15	-0.08

Notes: This table presents annualized average returns and the corresponding sharpe ratio of six currency portfolios that are rebalanced quarterly. First, currencies are sorted into portfolios according to their interest differential vis-à-vis the U.S. In a second step, I distinguish between high and low stock return differentials compared to the U.S among the interest rate sorted portfolios. Portfolio P11 contains the countries with highest interest rates and among those with highest stock returns relative to the U.S. Portfolio P32 hence consists of countries with lowest interest rate and among those with high stock return differential. The Sharpe-ratio is the ratio of mean return and standard deviation of the return. The sample period covers the

[14] The number of six portfolios is chosen to make sure that each of the portfolios contains currencies from at least two countries over the whole sample period.

time from the first quarter of 1974 to the fourth quarter of 2005 and comprises data on 20 developed countries.

The portfolio characteristics convey the notion that a) high interest rate currencies offer relatively high average returns and b) high foreign stock returns relative to the U.S. are associated with depreciating foreign currencies and hence negative or lower foreign currency returns than low stock return countries from the perspective of a U.S. investor. The clear exceptions are the middle interest rate portfolios, P21 and P22, for which high stock return differential currencies offer higher returns than countries with low stock returns compared to the U.S. Nonetheless, the portfolio characteristics support Lustig and Verdelhan (2006) as well as Cappiello and De Santis (2005) and Hau and Rey (2004). They reflect the failure of uncovered interest rate parity and at the same time the presence of the persistent impact of stock return differential shocks on nominal exchange rate changes. High interest rates relative to the U.S. signal appreciating foreign currencies and hence high foreign currency returns for a U.S. investor. Past high foreign stock returns in excess of the U.S. stock market return reflect low returns on these currencies.

3.3.2 Risk prices

As in the second chapter, I receive a representation of equation (3.2) in terms of risk prices by exploiting $E(XY) = E(X)E(Y) + \text{cov}(X,Y)$, such that

$$E_t(R_{t+1}^i) = \frac{1}{E_t(M_{t+1})} + \left(\frac{\text{cov}(R_{t+1}^i, M_{t+1})}{\text{var}(M_{t+1})}\right)\left(-\frac{\text{var}(M_{t+1})}{E_t(M_{t+1})}\right) \qquad (3.10)$$

which can be summarized to

$$E_t(R_{t+1}^i) - R_t^f = \beta_M^i \lambda_M \qquad (3.11)$$

with $R_t^f = \dfrac{1}{E_t(M_{t+1})}$, the risk-free rate, $\beta_M^i = \dfrac{\operatorname{cov}(R_{t+1}^i, M_{t+1})}{\operatorname{var}(M_{t+1})}$, the regression coefficient of excess return i on the discount factor, representing the asset-specific quantity of systematic risk, and $\lambda_M = -\dfrac{\operatorname{var}(M_{t+1})}{E_t(M_{t+1})}$ interpreted as the price of risk.

As consumption growth is the only source of systematic risk in this setting, (3.11) collapses to

$$E_t(R_{t+1}^i) - R_t^f = \beta_{\Delta c}^i \lambda_{\Delta c} \tag{3.12}$$

in which $\lambda_{\Delta c}$ is the price of consumption growth risk and $\beta_{\Delta c}^i$ the asset-specific exposure to consumption growth risk. High currency returns should thus be associated with high consumption growth betas as $\lambda_{\Delta c}$ is the same for all assets.

I use the Fama-MacBeth cross-sectional regression (Fama and MacBeth, 1973) to estimate the beta representation (3.12). The first stage of the Fama-MacBeth estimation is a time series regression of foreign currency returns on consumption growth

$$r_t^i = \mu + \beta_{\Delta c}^i \Delta c_t + \varepsilon_t^i \tag{3.13}$$

The estimated betas measure the exposure of returns on foreign currencies to consumption growth risk. In the second step of the Fama-MacBeth regression, I assess if differences in the exposure to consumption growth risk can account for differences in average returns on currencies. I thus run cross-sectional regressions at each point in time, i.e.

$$r_t^i = \lambda_0 + \lambda_{\Delta c} \hat{\beta}_{\Delta c}^i + v_t^i, \forall t \tag{3.14}$$

allowing for an intercept λ_0.

Table 3.2 presents the results. I report the point estimates in percentage points per quarter. T-statistics in parenthesis appear below the estimates and are corrected for the fact that the $\hat{\beta}_{\Delta c}^i$

are generated regressors (Shanken, 1992). The column R² gives the cross-sectional R² adjusted for the number of regressors as used in Campbell and Vuolteenaho (2004).[15]

Table 3.2: Risk price estimates (currency portfolios)

	λ_0	$\lambda_{\Delta c}$	R^2
CCAPM	-1.22	**0.50**	0.49
	(-1.73)	**(2.32)**	

Notes: This table reports the risk price of consumption growth (in percentage points per quarter) from a cross-sectional Fama-MacBeth CCAPM regression of the form:

$$r_t^i = \lambda_0 + \lambda_{\Delta c}\hat{\beta}_{\Delta c}^i + \varepsilon_t^i, \forall t$$

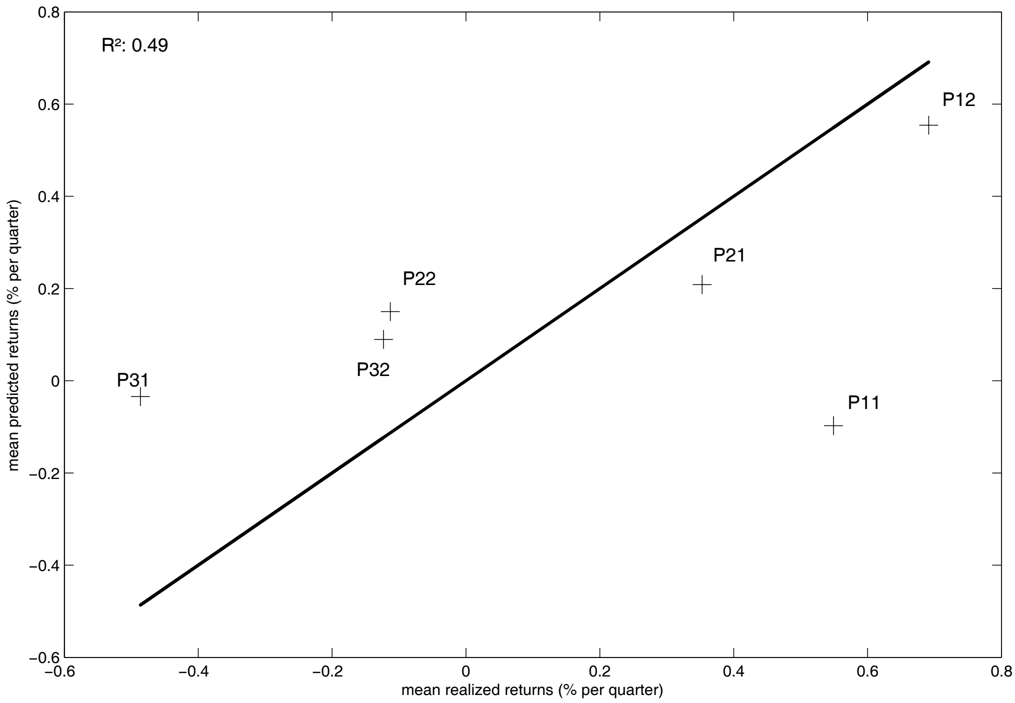
in which r_t^i denotes deflated returns on currency portfolio *i*. The betas are obtained from multiple time-series regressions of the currency portfolio returns on U.S. consumption growth. The sample period runs from the first quarter of 1974 to the second quarter of 2003. As the betas are generated regressors, t-statistics that are corrected for this errors-in-variables issue (Shanken, 1992) are presented in parenthesis. R² denotes the adjusted cross-sectional R².

The estimated price of consumption growth risk, $\lambda_{\Delta c}$, is statistically significantly different from zero. The price of 0.5 percentage points per quarter corroborates the annual price of consumption growth risk of approximately two percentage points p.a. estimated by Lustig and Verdelhan (2006). Judged by the R^2, the simple CCAPM explains about half of the cross-sectional variation in average returns on currency portfolios.

Figure 3.1 plots mean realized excess returns on currency portfolios against their values predicted by the CCAPM to provide a visual impression of the fit of model. The CCAPM has difficulties to price the high stock return portfolios in the high and low interest rate differential bin while it does reasonably well for the other portfolios.

[15] The R² is defined as $R^2 = 1 - \frac{e'e}{(\overline{R}^i - \sum_j \overline{R}^i)'(\overline{R}^i - \sum_j \overline{R}^i)}$ where $\overline{R}$ denotes a time series average

Figure 3.1: Fit of the CCAPM (returns on currency portfolios)



3.3.3 Consumption growth betas

The first stage of the Fama and MacBeth cross-sectional regression is useful to assess if the notion that high currency returns should be associated with high, positive consumption growth betas pertains in this setting.

Table 3.3 reports the consumption growth betas of the six currency portfolio excess returns. Newey-West (Newey and West, 1987) corrected t-statistics are in parenthesis. Figure 3.2 summarizes the results graphically.

If consumption growth betas were to explain the mean return on the respective currency portfolio exactly, then the points in figure 3.2 would line up perfectly. High consumption betas should be associated with high average returns. This is exactly the pattern that we observe in figure 3.2 with the exception of the portfolio consisting of countries with highest interest rate differential and among those with high stock return differential relative to the U.S.

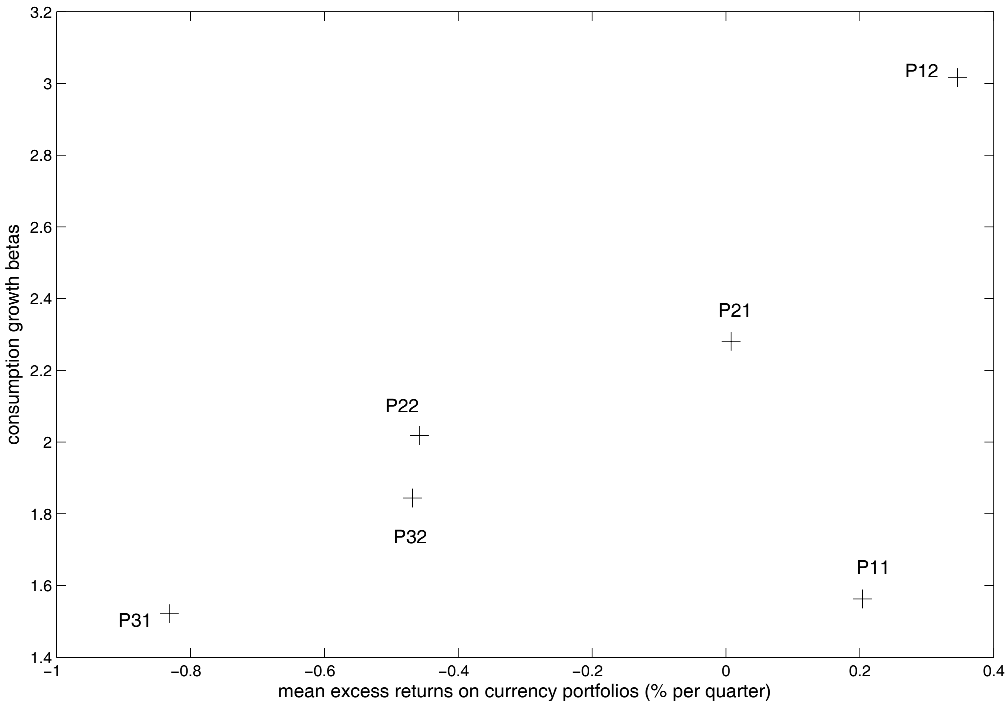
Table 3.3: Consumption growth betas (currency portfolios)

	Mean return (% per quarter)	Constant	$\beta_{\Delta c}$
P11	0.20	−0.53	1.56
		(−0.58)	(1.21)
P12	0.35	−1.07	**3.02**
		(−1.41)	**(2.52)**
P21	0.01	−1.06	2.28
		(−1.31)	(1.94)
P22	-0.46	−1.41	2.02
		(−1.56)	(1.50)
P31	-0.83	−1.55	1.52
		(−1.87)	(1.16)
P32	-0.47	−1.34	1.84
		(−1.53)	(1.33)

Notes: This table presents OLS estimates from the first stage of the FamaMacBeth regression of currency portfolio returns on U.S. consumption growth. The estimate equation takes the following form : $r_t^i = \mu^i + \beta_{\Delta c}^i \Delta \log c_t + \varepsilon_t^i$

T-statistics in parenthesis are computed using Newey-West corrected standard errors with K+1 lags where K is the number of regressors. The results are visualized in figure 3.2.

Figure 3.2: Mean returns on currency portfolios vs. consumption growth betas



But apart from this outlier, figure 3.2 supports the view that high average returns on currencies can be rationalized by their consumption growth betas conditional on instrumental variables (here interest rate *and* stock return differentials). The currencies which are strongest positively related to U.S. consumption growth, and thus destabilize the U.S. investor´s consumption path, have to compensate this riskiness with high average returns.

3.3.4 GMM estimates

The representation (3.9) naturally suggests using pricing errors as moment conditions in the GMM framework of Hansen and Singleton (1982) to estimate the parameter $b_{\Delta c}$ and to judge the quality of the model on the basis of the restrictions on expected discounted returns imposed by equation (3.9). Remember that $b_{\Delta c}$ reflects the coefficient of relative risk aversion. An estimate of the risk aversion parameter will tell us at what cost the success of the CCAPM in capturing the cross-section of currency portfolio returns comes. Note, that GMM chooses the parameter $b_{\Delta c}$ to minimize the pricing errors, i.e. to make the CCAPM explain the cross-section of currency excess returns as well as possible.

In equation (3.7), I follow common practice and set α to unity to consider

$$m_{t+1} = 1 - b_{\Delta c}\Delta c_{t+1} \tag{3.15}$$

as specification of the stochastic discount factor. The moment conditions are

$$g_T(b) = E_T(m_t \boldsymbol{r}_t) = E_T(\boldsymbol{r}_t) - E_T(\boldsymbol{r}_t \boldsymbol{f}_t{}')b \tag{3.16}$$

where $\boldsymbol{r}_t$ is the vector of returns on the six currency portfolios, $\boldsymbol{f}_t$ is a vector of pricing factors, here only consumption growth, and E_T denotes a time series average.

In the first stage of the GMM estimation, I use the identity matrix as weighting matrix and in the second stage the optimal weighting matrix with N+1 lags. N is the number of test assets. The GMM results reported in this paper remain qualitatively the same if I regard only first-stage GMM estimates or if I let the number of lags in the second stage estimation vary around reasonable values.

Table 3.4 presents the estimate of $b_{\Delta c}$ with t-statistics in parenthesis, the price of consumption growth risk implied by the GMM estimate as well as the p-value of the test of overidentifying restrictions for the null that all pricing errors are jointly zero.

Table 3.4: GMM estimates (currency portfolios)

	$b_{\Delta c}$	implied $\lambda_{\Delta c}$	J-Test (p-value)
CCAPM	**134.25** **(5.19)**	0.51	0.46

Notes: This table reports estimates of the coefficient of relative risk aversion from a two-stage GMM estimation of linear versions of the CCAPM. The stochastic discount factor is specified as $m_{t+1} = 1 - b_{\Delta c} \Delta \log c_{t+1}$
I use the identity matrix in the first stage and the optimal weighting matrix in the second stage applying a lag length of N+1, where N is the number of test assets. The implied risk prices, $\lambda_{\Delta c}$, are calculated from $\lambda = E(ff')b$. The J-Test is a χ^2-test of the null that all pricing errors are jointly zero with degrees of freedom equal to number of moments less number of parameters.

The estimate corroborates the well-known observation that the success of consumption-based models in explaining the cross-section of asset returns comes at the cost of implausibly high estimates of the risk aversion coefficient (Mark, 1985; Campbell and Cochrane, 1999; Yogo, 2006). However, the risk price of consumption growth implied by the GMM estimates is not too far away from the Fama-MacBeth regression estimates. Furthermore, we cannot reject the hypothesis of all pricing errors being zero at conventional confidence levels. The GMM results thus complement the estimation of risk prices and underscore that the CCAPM is able to capture the cross-sectional dispersion of returns on foreign currencies.

3.4 Stock return differentials, consumption growth and exchange rate changes: A country-by-country analysis

So far I have argued that the evidence of exchange rate predictability by Cappiello and De Santis (2005) and the persistence of stock return differential shocks for exchange rate changes (Hau and Rey, 2004) qualify lagged foreign stock market returns in excess of the U.S. stock market as explanatory variable for exchange rate changes.

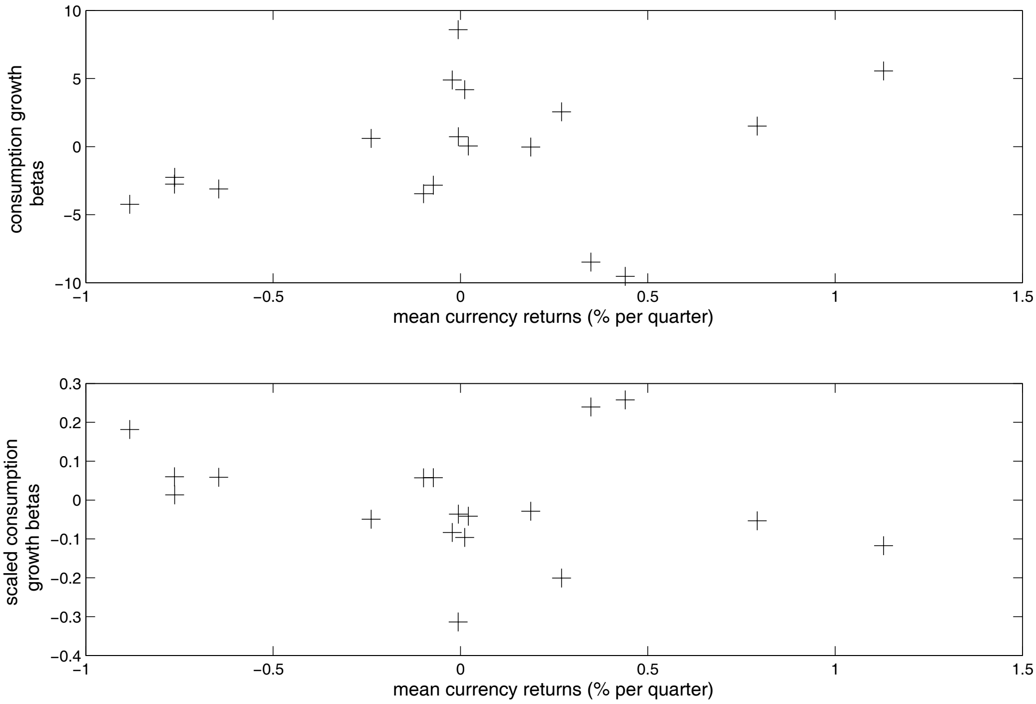
However, I find it useful to assess the relationship of currency returns with past stock return differentials in light of the implications from consumption based asset pricing for this relation. Currencies from countries with high past stock returns compared to the U.S. should depreciate against the U.S. dollar and thus offer low or even negative returns. Hence these currencies should have relatively low or negative consumption growth betas.

To test this conjecture, I run regressions of U.S. dollar exchange rate changes on U.S. consumption growth and control for time variation in relative stock returns by interacting U.S. consumption growth with the respective one-quarter lagged stock return differential. The regression takes the following form

$$\Delta s_{t+1}^{i} = \mu + \beta_{\Delta c} \Delta c_{t+1} + \beta_{\Delta c,r} \Delta c_{t+1} (sr_t^i - sr_t) + \varepsilon_{t+1} \tag{3.19}$$

for U.S. dollar exchange rate changes of country *i* where $sr_t^i - sr_r$ represents the stock market differential of country *i* compared to the U.S. and Δc_{t+1} denotes U.S. consumption growth. The point estimates as well as Newey-West corrected t-statistics are displayed in table A.2.1 in the appendix. Figure 3.3 visualizes the results. The upper panel presents the consumption growth betas, the lower panel shows the betas with respect to consumption growth interacted with lagged stock return differentials.

Figure 3.3: Mean currency returns and consumption growth betas conditional on time variation in stock return differentials



The upper panel of figure 3.3 displays that returns on individual currencies line up with their consumption growth betas. High returns on currencies are associated with high consumption growth betas, whereas low currency returns seem to be the consequence of low covariances with consumption growth. This is exactly what is implied by a consumption-based asset pricing framework.

Furthermore, the lower panel of figure 3.3 provides support for the view that lagged stock return differentials are informative about exchange rate changes in the way as proposed by UEP. Conditioning consumption growth with stock return differentials leaves the impression that high returns on the foreign stock market relative to the U.S. are associated with a foreign currency depreciation and hence with low (negative) returns for a U.S. investor.[16]

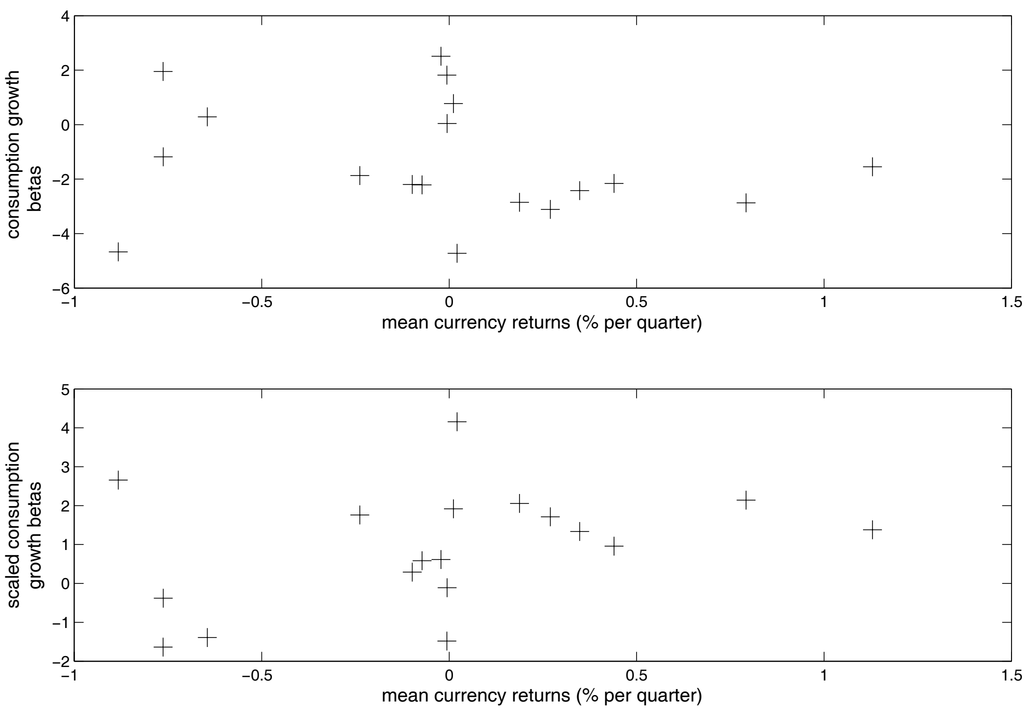
[16] This observation is qualitatively unaffected by considering consumption growth scaled with interest rate differentials as additional regressor. Results are not reported but available upon request.

For comparison, figure 3.4 gives a visual impression of the results from a regression using the corresponding interest rate differential as scaling variable. Table A.2.2 in the appendix displays the OLS estimates of the regression

$$\Delta s_{t+1}^{i} = \mu + \beta_{\Delta c} \Delta c_{t+1} + \beta_{\Delta c,i} \Delta c_{t+1} (i_t^i - i_t) + \varepsilon_{t+1} \tag{3.21}$$

where $i_t^i - i_t$ is the interest rate differential of country i with the U.S. In analogy to the latter finding, high average currency returns should be associated with high consumption growth betas. Betas of the interacted consumption growth term should reflect that currencies from high interest rate countries promise high returns.

Figure 3.4: Mean currency returns and consumption growth betas conditional on time variation in interest rate differentials



Evidence for the latter implication gleams through the lower panel of figure 3.4. But there does not seem to be a systematic relation between average currency returns and their consumption growth betas on a country-by-country basis conditional on time-variation in interest rate differentials.

3.5 Managed returns on foreign currency

Stock return differentials among developed countries are important in explaining quarterly exchange rate changes. Returns on foreign currencies can be rationalized by their sensitivity to consumption growth once one takes time variation in stock return differentials into account. This finding conveys the notion that the conditional implications of the CCAPM for individual currency returns can be tested by regarding returns on foreign currencies scaled with their respective stock return differentials vis-à-vis the U.S.[17] Stock return differentials then serve as conditioning instrument or signal, such that scaled currency returns can be interpreted as investment in an asset managed according to this signal (Cochrane, 1996).

Panel A of table 3.5 presents estimates from a Fama-MacBeth cross-sectional CCAPM regression performed on 15 quarterly currency returns scaled with the respective stock return differential relative to the U.S.

The estimated price of consumption growth risk is statistically significant and near the estimated value when returns on currency portfolios are considered. The cross-sectional R^2 indicates that the CCAPM explains about 44 percent of the cross-sectional dispersion in scaled returns on foreign currencies. Note also that the R^2 is close to the value obtained for currency portfolios

Panel B gives the corresponding GMM estimates. These results largely corroborate the previous results for currency portfolios and show that the CCAPM passes the test of overidentifying restrictions when confronted with scaled currency returns.

[17] In order to maintain the scale of the moments of the currency returns, they are scaled with $1+\frac{sr_t^i - sr_t}{\sigma(sr_t^i - sr_t)}$ where $\sigma(sr_t^i - sr_t)$ is the standard deviation of the (demeaned) stock return differentials.

Table 3.5: Risk price and GMM estimates (scaled currency returns)

Panel A: Fama-MacBeth			
	λ_0	$\lambda_{\Delta c}$	R^2
CCAPM	**− 2.63**	**0.54**	0.44
	(−4.13)	**(3.16)**	
Panel B: GMM			
	$b_{\Delta c}$		J-Test (p-value)
CCAPM	**177.79**		0.95
	(8.06)		

Notes: Panel A of this table gives the risk prices from cross-sectional Fama-MacBeth CCAPM regressions of the form

$$r_t^{e,i} = \lambda_0 + \lambda_C \beta_{\Delta c}^i + \varepsilon_t^i, \forall t$$

of individual currency returns scaled with the respective stock return differential relative to the U.S. Betas are obtained from time-series regressions on consumption growth. As the betas are generated regressors, t-statistics that are corrected for this errors-in-variables issue (Shanken, 1992) are reported in parentheses. R^2 denotes the adjusted cross-sectional R^2.
The sample period is from the first quarter of 1974 to the first quarter of 2005.
Panel B of this table reports estimates of the coefficient of relative risk aversion from a two-stage GMM estimation of linear versions of the CCAPM. The stochastic discount factor is specified as $m_{t+1} = 1 - b_{\Delta c} \Delta \log c_{t+1}$
I use the identity matrix in the first stage and the optimal weighting matrix in the second stage applying a lag length of 12. The qualitative results are not influenced by the choice of lag length. The J-Test is a χ^2-test of the null that all pricing errors are jointly zero with degrees of freedom equal to number of moments less number of parameters.

Figure 3.5 displays the pricing errors of the CCAPM. I present mean actual currency returns in percentage points per quarter (horizontal axis) compared with the returns predicted by the model. The CCAPM seems to fit the data reasonably well also in terms of pricing errors. Figure 3.6 gives the scaled currency returns (horizontal axis) relative to their consumption growth betas (vertical axis). High currency returns are associated with high consumption growth betas and vice versa.

Figure 3.5: Fit of the CCAPM for scaled foreign currency returns

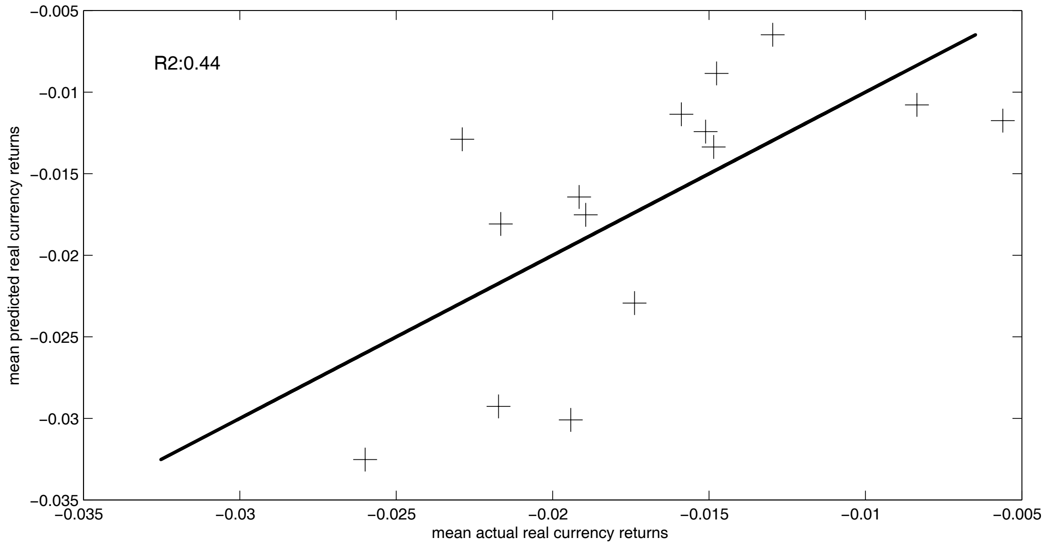
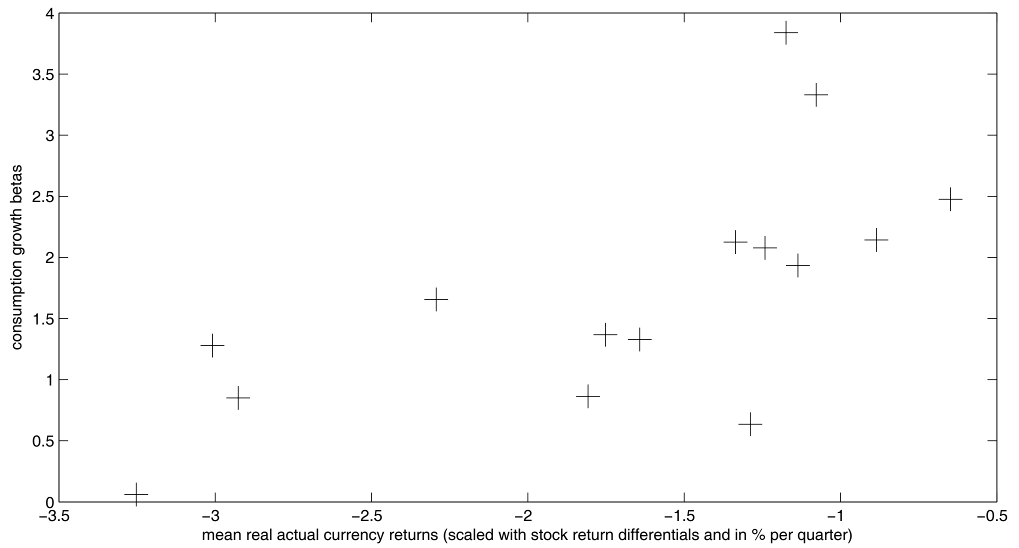


Figure 3.6: Scaled foreign currency returns and consumption growth betas



All in all, this evidence leaves the impression that, conditional on the respective stock market return differentials relative to the U.S., the CCAPM is not only able to explain the cross-sectional variation in currency portfolio but also individual quarterly currency returns. This relation holds if the sample is restricted to countries for which uncovered equity parity holds.

In such a setting prior formation of portfolios does not seem to be necessary in order to test the conditional implications of the CCAPM for returns on foreign currency.

3.6 Conclusions

Lagged foreign stock returns in excess of the U.S. stock market return are informative about quarterly exchange rate movements. A past high foreign stock return relative to the U.S. signals a foreign currency depreciation and hence low returns on the foreign currency.
Conditional on stock return differentials, the consumption-based CAPM (CCAPM) explains the cross-sectional dispersion in U.S. dollar exchange rates. The CCAPM captures more than 40 percent of the variation in foreign currency returns scaled with the respective stock return differential on a country-by-country basis. The performance of the CCAPM for returns on interest rate and stock return differentials sorted currency portfolios is only slightly better.

Chapter 4

The consumption-exchange rate anomaly: An asset pricing perspective

4.1 Introduction

In this chapter, which is joint work with Professor Mathias Hoffmann, we provide evidence that an international variant of the consumption CAPM that takes account of the heterogeneity in international consumption growth rates captures the cross-section of currency returns. The CCAPM with heterogeneity explains more than 60 percent of the cross-sectional variation in quarterly exchange rate changes. Our results are obtained from data of 13 industrialized countries in the post Bretton-Woods period.

The mechanism that explains the empirical success of our CCAPM with heterogeneity is closely related to the solution of the UIP puzzle recently first formalized by Lustig and Verdelhan (2006). At the heart of the mechanism is the insight that investing into foreign currency amounts to a bet on the foreign intertemporal marginal rate of substitution: countries with relatively high consumption growth will tend to have depreciating currencies, those with low consumption growth appreciating ones.

According to the CCAPM, currencies -- like any asset -- should be priced for the exposure to world aggregate consumption growth risk they deliver. In terms of international consumption comovements, this risk has two determinants: first, given the relative volatility of consumption growth rates, the currencies of countries with high consumption correlations offer a better hedge since these currencies will depreciate less (or may even appreciate) when average world consumption growth is low. Secondly, given consumption correlations, the relative volatility of consumption determines the size of the risk premium; if consumption growth rates are positively correlated, countries with more volatile consumption growth rates offer a good hedge against world consumption growth risk since they tend to appreciate when world consumption growth is low. These currencies therefore generate low excess returns. Conversely, currencies of countries whose consumption growth is negatively related to world consumption but more volatile will offer a particularly bad hedge and will therefore show positive expected excess returns.

The logic of this mechanism implies that the international dispersion of consumption growth rates should be a key determinant of the cross-section of exchange rate changes and currency excess returns. This insight forms the point of departure for our analysis in this chapter. While we are not the first to estimate a CCAPM with heterogeneous consumption, very few contributions have actually looked at exchange rate determination (a prominent exception being Sarkissian (2003), but he prices forward rates while we look at exchange rate changes directly, i.e. cash returns). More importantly, the content of this chapter also differs in scope from earlier applications by focussing the analysis on the particular mechanism to which Lustig and Verdelhan have only recently drawn the profession's attention and on which the heterogeneous agent CCAPM formulation we chose here can provide a natural, interesting, and complementary perspective.

As a first important result we find that the relative volatility of consumption growth rates explains most of the cross-sectional variation in exchange rate changes. We do not only explain a large fraction of the cross-sectional variation in exchange rate changes, most of this explanatory power eventually comes from the international consumption dispersion term.

International consumption heterogeneity is also at the heart of another empirical success of our approach: we price currency returns in quarterly data without having to recur to instrumental variables such as stock return or interest rate differentials. In their prominent paper, Lustig and Verdelhan form annually rebalanced currency portfolios using interest differentials as their instrument. The portfolio formation is key in reconciling the various versions of the CCAPM with the data. While the usual justifications for sorting currencies into portfolios certainly also apply to their analysis, we argue that sorting currencies by their interest rate differentials is conceptually very similar to using the cross-sectional dispersion of consumption growth rates as an additional pricing factor.

Furthermore, we explore the implications of our results for the macroeconomic literature on consumption risk sharing. In particular, we shed new light on the consumption-real exchange rate anomaly that so far has mainly been examined in the time-series dimension (Backus and Smith, 1993; Kollmann, 1995 and more recently Ravn, 2001). We argue that the empirical failure of UIP and the difficulty to detect the Backus-Smith link in the data are closely related and that an asset pricing model that can account for departures from UIP should also be able to account for departures from the Backus-Smith relation. In a world with complete financial and goods markets and homothetic preferences, consumption growth rates should equalize. Taking account of traded and non-traded goods, Backus and Smith (1993) show that deviations from purchasing power parity drive a wedge between consumption growth rates. At the same time, real exchange rates should be determined by relative consumption growth rates. However, empirical support for this condition is virtually absent. The correlations of relative consumption growth rates with real exchange rates are extremely low. We show that a somewhat weaker version of the Backus-Smith relation only holds when UIP holds. UIP is grossly violated in the data, but departures from UIP are captured by our model. According to Backus and Smith (1993) cross-sectional differences in price level changes (inflation) should explain cross-sectional differences in (real) exchange rate changes. We therefore decompose the cross-sectional variation in consumption into one component that is due to cross-sectional differences in inflation expectations and a second component that is due to international variation in nominal interest rates (the key to success in the Lustig and Verdelhan (2006) framework). Both factors account to virtually equal parts for the pricing power of the cross-sectional variance of consumption. We thus provide (indirect) evidence for the Backus-Smith condition. Our results hence suggest that the failure of uncovered interest parity and the consumption real exchange rate anomaly are more closely linked than may have commonly been believed.

As a policy application of our framework, we ask how the price of idiosyncratic consumption risk -- as an indicator of the marginal value of buying an additional unit of consumption insurance -- has changed over time, as world financial markets have become more integrated. We find that both the quantity of idiosyncratic risk as well as its marginal disutility have decreased substantially, in particular since 1990. This finding ties in with an important and by now well-established literature that dates the onset of the recent major wave in financial globalization in the first half of the 1990s (Lane and Milesi-Ferretti, 2002, 2003, 2004).

The remainder of this chapter is organized as follows. Section 4.2 presents our theoretical framework and relates it to the Backus-Smith puzzle. We present our main results in section 4.3 and finally conclude in section 4.4.

4.2 A simple framework

In complete currency markets, the change in the real exchange rate can be expressed as the relative intertemporal marginal rate of substitution:

$$\Delta q_{t+1}^{k} = m_{t+1}^{k} - m_{t+1} \tag{4.1}$$

where Δq_{t+1}^{k} is the percentage change in the real exchange rate of country k and m_{t+1} and m_{t+1}^{k} are the logarithmic home and foreign discount factors respectively. In a setting with constant relative risk aversion, (4.1) can be specialized to

$$\Delta q_{t+1}^{k} = \gamma(\Delta c_{t+1}^{k} - \Delta c_{t+1}) \tag{4.2}$$

where γ is the coefficient of risk aversion, Δc denotes consumption growth and superscript k denotes the country k foreign variable. Since Backus and Smith (1993), this relationship has been the focus of much empirical research. Most studies find the link between real exchange rates and consumption tenuous at best, the correlation generally insignificant and wrongly signed. In our analysis here, we build on Lustig and Verdelhan (2006) to explore this link in the cross-section. In so doing, we use a simple version of the consumption capital asset pricing model (CCAPM).

The standard version of the CCAPM assumes the existence of a representative agent which implies that consumption growth rates should be equalized internationally. However, the very notion of equation (4.2) is that international consumption growth rates differ across countries and since Backus, Kehoe and Kydland (1992) low international consumption correlations are a prominently documented feature of the data. We therefore adapt our CCAPM to account for this international consumption heterogeneity. We start from the first-order condition of a country k investor with respect to holdings of foreign currency:

$$E_t(M_{t+1}^k R_{t+1}^j) = 1 \tag{4.3}$$

where R_{t+1}^j is the return of holding one unit of country j currency relative to an arbitrary base currency (We will generally use U.S. dollars as the base denomination). We discuss the exact specification of this return below.

For now we average condition (4.3) over all countries to obtain

$$E_t\left(\frac{1}{K}\sum_{k=1}^{K} M_{t+1}^k R_{t+1}^j\right) = 1 \tag{4.4}$$

where K is the number of countries. We denote the average pricing kernel $\frac{1}{K}\sum_{k=1}^{K} M_{t+1}^k$ with $\overline{M}_{t+1}$. Under CRRA-utility, we can approximate

$$M_{t+1}^k = \beta(1+\Delta c_{t+1}^k)^{-\gamma} \tag{4.5}$$

where Δc_{t+1} is consumption growth, γ is relative risk aversion and β is the subjective discount factor. We then expand $\overline{M}_{t+1}$ around world average consumption growth, $\Delta\bar{c}_{t+1}$ to obtain

$$\begin{aligned}\overline{M}_{t+1} &\approx \beta(1+\Delta\bar{c}_{t+1})^{-\gamma} - \frac{\gamma}{K}\sum_{k=1}^{K}\beta(1+\Delta\bar{c}_{t+1})^{-\gamma-1}(\Delta c_{t+1}^{k}-\Delta\bar{c}_{t+1}) \\ &+ \frac{\gamma(\gamma+1)}{2K}\sum_{k=1}^{K}\beta(1+\Delta\bar{c}_{t+1})^{-\gamma-2}(\Delta c_{t+1}^{k}-\Delta\bar{c}_{t+1})^{2}\end{aligned} \tag{4.6}$$

The first order term is zero by construction whereas for large K, the second term converges to a multiple of the cross-sectional variance of Δc_{t+1}^{k}. We thus obtain

$$\overline{M}_{t+1} \approx \beta(1+\Delta\bar{c}_{t+1})^{-\gamma}\left[1+\frac{\gamma(\gamma+1)\sigma_{K,t}^{2}}{2(1+\Delta\bar{c}_{t+1})^{2}}\right] \tag{4.7}$$

In our empirical implementation, we use the logarithm of $\overline{M}_{t+1}$ so that

$$\bar{m}_{t+1} \approx \kappa - \gamma\Delta\bar{c}_{t+1} + \delta\frac{\sigma_{K,t}^{2}}{(1+\Delta\bar{c}_{t+1})^{2}} \tag{4.8}$$

where $\delta = \gamma(\gamma+1)/2$ and κ is a constant term. Hence, the international heterogeneity of consumption growth rates gives rise to a second pricing factor, the cross-sectional variance of consumption, scaled by the square of the aggregate consumption growth rate. Since the insurance mechanism laid out in the introduction explicitly relies on differences in consumption growth rates across countries, it is natural to use these differences as pricing factor. In light of the mechanism we discussed in the introduction, it would even seem that an empirical implementation of the CCAPM that does not account for international differences in consumption growth rates is misspecified. In fact, our results suggest that the sorting of individual currencies into portfolios according to the size of the interest rate differential (as done in Lustig and Verdelhan (2006)) implicitly amounts to using the cross-sectional variance term as an instrument: using the cross-sectional variance term directly, we price quarterly currency returns in industrialised countries without prior portfolio formation and without using instrumental variables.

4.2.1 Interest Parity and the Backus-Smith condition

If uncovered interest parity holds, regressions of exchange rate changes on the nominal interest rate differential of the form

$$\Delta e_{t+1}^{k} = \alpha_k (i_t^k - i_t) + \varepsilon_t \tag{4.9}$$

where Δe_{t+1}^{k} is the change in the nominal log exchange rate (measured in terms of foreign (country k) currency to home currency) and i_t denotes the nominal exchange rate, should yield a coefficient of unity. The UIP puzzle, first stated by Fama (1984), is the empirical regularity that α_k is typically much smaller than unity and often even negative in the data. The excess return of investing into foreign over domestic bonds is given by $i_t^k - i_t - \Delta e_{t+1}^k$. Hence, a value of $\alpha < 1$ implies that the expected return on investing in foreign currency is predictable.

$$E_t (i_t^k - i_t - \Delta e_{t+1}^k) = (1 - \alpha_k)(i_t^k - i_t) \tag{4.10}$$

so that high interest rate currencies generate a positive excess return.

To highlight the link between the UIP puzzle and the Backus-Smith condition, we find it useful to start from the first-order condition of the country k investor for whom, in analogy to (4.3) above, the risk-free rate is determined by

$$E_t (M_{t+1}^k R_{t+1}^f) = 1 \tag{4.11}$$

Note that the country k bond is only a safe asset for country k investors.[18] Under our maintained assumption of CRRA-utility, a second-order expansion of this condition (and the assumption of log-normality of consumption growth) yields

$$i_{t+1}^k = E_t(\pi_{t+1}^k) + \kappa + \gamma E_t(\Delta c_{t+1}^k) - \frac{\gamma^2}{2} \mathrm{var}_t(\Delta c_{t+1}^k) \tag{4.12}$$

[18] Or, more precisely, the asset without exchange rate risk, since country k investors may still face domestic inflation risk.

where $\kappa = \log(\beta)$ and π_{t+1}^k is the inflation rate in country k. Ignoring the variance term that reflects the precautionary savings motive and plugging in for $i_t^k - i_t$, we can write the excess return of investing into foreign currency bonds as

$$E_t(\pi_{t+1}^k - \pi_{t+1}) + \gamma E_t(\Delta c_{t+1}^k - \Delta c_{t+1}) - E_t(\Delta e_{t+1}^k) \qquad (4.13)$$

Under UIP, this excess return should be zero, so that we obtain

$$E_t(\Delta e_{t+1}^k + \pi_{t+1} - \pi_{t+1}^k) = \gamma E_t(\Delta c_{t+1}^k - \Delta c_{t+1}) \qquad (4.14)$$

This equation is almost identical with the Backus-Smith condition (4.2). It is however, important to note that (4.14) holds in expectations only, whereas (4.2) holds in all states of nature; whereas the latter condition will only be satisfied in complete markets, the former is much more general. However, even tests that have focused on the much weaker version of the condition, e.g. Kollmann (1995), have generally found little evidence to support it. Our derivation here suggests that the failure to identify (4.2) – or even the weaker condition (4.14) - in the data is intimately linked with the failure of UIP. Clearly, condition (4.14) will hold only if UIP holds. Since UIP is grossly violated in the data, there may be little hope to identify a relation such as (4.2) directly.

As an alternative approach, we therefore suggest to explore the link between exchange rates, consumption, inflation and interest rates by exploiting the asset pricing implications of (4.14) using the pricing kernel expansion (4.8).

We first write

$$\sigma_{K,t}^2 = \mathrm{var}_t^K(\Delta c_{t+1}^k) = \mathrm{var}_t^K(E_t(\Delta c_{t+1}^k) + v_{t+1}^k \qquad (4.15)$$

where v_{t+1}^k is an i.i.d. disturbance term. Since, according to (4.14), the cross-sectional variance of consumption should equal the cross-sectional variance of real interest rates, we can write

$$\sigma_{K,t}^2 = \mathrm{var}_t^K(\Delta c_{t+1}^k) = \frac{\mathrm{var}_t^K(i_t^k) - 2\,\mathrm{cov}_t^K(i_t^k, E_t(\pi_{t+1}^k)) + \mathrm{var}_t^K(\pi_{t+1}^k))}{\gamma^2} \qquad (4.16)$$

Plugging this relation into (4.8) and dropping the residual term $\mathrm{var}_t^K(v_{t+1}^k)$, we obtain

$$\overline{m}_{t+1} \approx \kappa - \gamma\Delta\overline{c}_{t+1} + \frac{(\gamma+1)}{2\gamma}\frac{\mathrm{var}_t^K(i_t^k) - 2\mathrm{cov}_t^K(i_t^k, E_t(\pi_{t+1}^k)) + \mathrm{var}_t^K(\pi_{t+1}^k)}{(1+\Delta\overline{c}_{t+1})^2} \qquad (4.17)$$

Note that it is an empirical question whether or not we can neglect $\mathrm{var}_t^K(v_{t+1}^k)$. To the extent that we can, the approximation (4.17) should price exchange rates almost as well as the consumption-based representation 4.8 above. The approximation here implies a model with four pricing factors: the first is world average consumption growth. The second factor is the cross-sectional variance of nominal interest rates. This factor is key in understanding the link between our results here and the ones obtained in the Lustig and Verdelhan study. As Lustig and Verdelhan themselves note, their representative agent CCAPM successfully prices excess returns on a set of currency portfolios that are formed by the size of the interest rate differential, but not the individual currencies. This is because the portfolio formation emphasizes the cross-sectional spread of interest rates and this spread is needed to explain the cross-section of currency excess returns. In our setting, where we set out by explicitly acknowledging the heterogeneity in consumption, the cross-sectional dispersion of interest rates arises naturally as a pricing factor. We believe that it is the feature of our model that is at the root of one of our main results: our ability to price currencies in a CCAPM-context without prior portfolio formation. We call $\mathrm{var}_t^K(i_t^k)$ the Lustig-Verdelhan (LV) factor.

Our framework gives rise to two more pricing factors. Most importantly, the term $\mathrm{var}_t^K(E_t(\pi_{t+1}^k))$ captures idiosyncratic inflation risk. It is this risk which is at the heart of the Backus-Smith condition: international consumption correlations should be perfect only to the extent that relative purchasing powers are equalized. Consumption growth should be high in countries with high domestic purchasing power (low prices) and low in places with high prices (low purchasing power). Differences in real exchange rates should thus be captured by deviations from purchasing power parity. We therefore call $\mathrm{var}_t^K(E_t(\pi_{t+1}^k))$ the Backus-Smith

(BS) factor. Finally, there is the covariance between nominal interest rates and expected inflation – a Fisher parity term- that measures the comovement between idiosyncratic interest rate and inflation risk.

We start our empirical analysis by using the purely consumption based representation (4.8) of the discount factor $\overline{m}_{t+1}$ to price currency returns. We then move on to the decomposition (4.16) of $\overline{m}_{t+1}$ that allows us to ask to what extent cross-sectional variation in interest and inflation rates contributes to the success of $\operatorname{var}_t^K(\Delta c_{t+1}^k)$ as a pricing factor.

We price currency returns in two ways: in a first step, we use pure cash returns, i.e. Δe_{t+1}. We then also consider excess returns on foreign currency bonds, $i_t^k - i_t - \Delta e_{t+1}^k$, and decompose these according to $\Delta e_{t+1}^k = \alpha_k (i_t^k - i_t) + \varepsilon_t$ into what we call a UIP component – measured by $\alpha_k (i_t^k - i_t)$ -- and a residual term. Our results show that irrespective of what formulation we use, what we price is truly the UIP component.

4.3 Data and Empirical Results

4.3.1 Data

Our full sample contains quarterly data on exchange rates as well as real, p.c. consumption of Australia, Austria, Canada, France, Germany, Italy, Japan, Norway, Spain, Sweden, Switzerland, United Kingdom and the United States for the time period from 1971Q1 to 2003Q2.

We calculate quarterly exchange rate changes from end-of-quarter MSCI stock market returns denominated in U.S. dollar (or the respective currency of interest) and local currency freely available on www.mscibarra.com.

We update the Campbell (1999) consumption data set with consumption, cpi and population data from the IFS January 2004 tape to construct world average consumption growth as well as the cross-sectional variance of world consumption growth at each point in time.

Since consumption growth is unitless, average world consumption growth is an equal-weighted average of growth rates from real, p.c. consumption in local currency of the 13 countries under consideration.

In order to illustrate the Lustig and Verdelhan (2006) mechanism in our framework we obtain short-term interest rates (treasury bills or call money market rates) from the IFS January 2004 CD following the recommendation by Campbell (1999).

4.3.2 Empirical Results

4.3.2.1 Pricing currency returns

We follow two different asset pricing approaches to explore the relation between our world consumption-based CAPM with heterogeneity and returns on currencies. First, we ask if one of the two factors in question, world consumption growth and the cross-sectional variance of consumption growth, helps to price currency returns given the presence of the other. Therefore, we use the generalized methods of moments (GMM) framework of Hansen and Singleton (1982) to exploit the testable restrictions imposed by

$$E_t(\overline{m}_{t+1} r^j_{t+1}) = 0$$

with $\overline{m}_{t+1}$ the natural logarithm of the average pricing kernel and r^j_{t+1} the log return on currency j. We simplify the log-linear discount factor to

$$m_{t+1} = 1 - b_1 \Delta \overline{c}_{t+1} - b_2 \sigma^2_{K,t} .^{19} \tag{4.17}$$

The GMM estimator chooses the parameters b_1 and b_2, such that the model fulfils the restrictions in (4.3) best. This procedure amounts to minimizing the errors of our pricing model. Therefore, we use a two-stage procedure as in the previous chapter. We employ the identity matrix as weighting matrix in the first step to obtain estimates of b_1 and b_2 and the optimal weighting matrix suggested by Hansen and Singleton (1982) in the second step to

[19] Scaling the cross-sectional variance term as in the formal derivation does not affect any of the results qualitatively.

compute standard errors. All of our results remain qualitatively the same if we use first-stage GMM estimates only. [20]

Secondly, we assess what risk factor is actually priced in currency returns. We estimate risk prices via the Fama – MacBeth (Fama and MacBeth, 1973) cross-sectional regression. In the first stage we run the following time series regressions

$$r_t^j = \mu + \beta_{\Delta \bar{c}}^j \Delta \bar{c}_t + \beta_{\sigma_K^2} \sigma_{K,t}^2 + \varepsilon_t \tag{4.18}$$

with r_t^j the log return on the country j dollar exchange rate, $\beta_{\Delta \bar{c}}^j$ the sensitivity of currency return j to world consumption growth and $\beta_{\sigma_K^2}^j$ the exposure of the return on currency j to idiosyncratic consumption risk, i.e. the cross-sectional variance of consumption.

In the second step we run the cross-sectional regressions (4.18) at each point in time to obtain estimates of the risk prices, λ.

$$r_{t+1}^j = \mu + \lambda_1 \hat{\beta}_{\Delta \bar{c}}^j + \lambda_2 \hat{\beta}_{\sigma_K^2}^j + e_{t+1}, \forall t \tag{4.19}$$

Table 4.1 presents GMM estimates and risk prices when we consider nominal returns on U.S. dollar exchange rate changes of twelve developed countries at the quarterly frequency.[21] We start with pricing nominal exchange rate changes because the 'true' rational expectations risk premium on foreign currency should be predominantly determined by the covariance of exchange rate changes (currency returns) with consumption growth or other systematic sources of risk as shown by Engel (1996) and Lustig and Verdelhan (2006).

[20] See Cochrane (2005) for an excellent introduction to the GMM framework in asset pricing.
[21] Australia, Austria, Canada, France, Germany, Italy, Japan, Norway, Spain, Sweden, Switzerland, United Kingdom

Table 4.1: GMM estimates and risk prices with Δe_t as test assets

	b_1	b_2	J-Test
GMM	**22.43** (4.31)	**357.35** (7.54)	0.44
	λ_1	λ_2	R^2
Fama-MacBeth	0.87 (1.05)	**0.14** (3.35)	0.62

Notes: The table provides estimates from two-stage GMM estimations of $0 = E_t(\overline{m}_{t+1}\Delta e_{t+1}^k)$ with $\ddot{m}_t = 1 - b_1\Delta\overline{c_t} - b_2 \underset{k}{\mathrm{var}}\Delta c_t^k$

where Δe_{t+1}^k represents U.S. dollar exchange rate changes defined as foreign currency over U.S. dollar, $\Delta\overline{c}_t$ average world consumption growth at time t and $\underset{k}{\mathrm{var}}\Delta c_t^k$ the cross-sectional consumption dispersion at time t. T-statistics occur in parenthesis. In the first stage we use the identity matrix as weighting matrix and in the second stage the Newey-West (Newey and West, 1987) corrected optimal weighting matrix as in Hansen and Singleton (1982). The column J-Test gives the p-value of a test of overidentifying restrictions that tests the null of all pricing errors being zero.

Furthermore, we present estimates of risk prices of the factors under consideration using the Fama-MacBeth cross-sectional regression which takes the following form:

$\Delta e_t^k = \mu + \lambda_1\beta_{\Delta\overline{c}}^k + \lambda_2\beta_{\underset{k}{\mathrm{var}}\Delta c_t^k}^k + v_t, \forall t$

The betas are obtained from time series regressions of exchange rate changes on average world consumption growth and the cross-sectional variance of consumption growth. T-statistics of the risk price estimates, λ, are corrected for generated regressor bias (Shanken, 1992). The sample spans the period from the first quarter of 1971 to the second quarter of 2003 and uses data on exchange rates and consumption from the following countries: Australia, Austria, Canada, France, Germany, Italy, Japan, Norway, Spain, Sweden, Switzerland, United Kingdom and the United States.

The row 'GMM' provides estimates of the parameters b_1 and b_2 with Newey-West (Newey and West, 1987) corrected t-statistics in parenthesis below the estimates. The column 'J-Test' gives the p-value of a test of the null that all pricing errors are jointly zero. Apparently we cannot reject this null at conventional confidence levels. Furthermore, the parameter estimates are both statistically significantly different from zero. Note also that b_1 mirrors the coefficient of constant relative risk aversion (CRRA) since we assume investors maximize a power utility function.

The estimate of the CRRA coefficient of about 22 is considerably lower than in other studies that focused on the pricing of currency returns with a consumption-based CAPM not allowing for consumer heterogeneity (see e.g. Mark, 1985; Lustig and Verdelhan, 2006). This result is due to the cross-sectional variance term that derives immediately from our heterogeneous agent framework. On the other hand, the coefficient on the cross-sectional variance term turns out to be incorrectly signed.[22] This pattern seems to be quite common in models with heterogeneous consumption and it lines up with the findings by Jacobs and Wang (2004) who introduce heterogeneity in domestic consumption into a CCAPM to price domestic stock returns. It is interesting that our results corroborate this pattern on a completely different sample period using currency returns instead of stock returns.

The GMM estimates reflect that both of the risk factors are helpful to price currency returns. However, we are particularly interested if these factors are actually priced and how large the respective price of risk is. The row 'Fama-MacBeth' of table 4.1 presents Fama-MacBeth cross-sectional regression estimates of the risk prices of our two factors in question. The t-statistics in parenthesis are adjusted for the fact that the regressors are generated (Shanken , 1992). R^2 denotes the cross-sectional R^2 proposed by Jagannathan and Wang (1996).

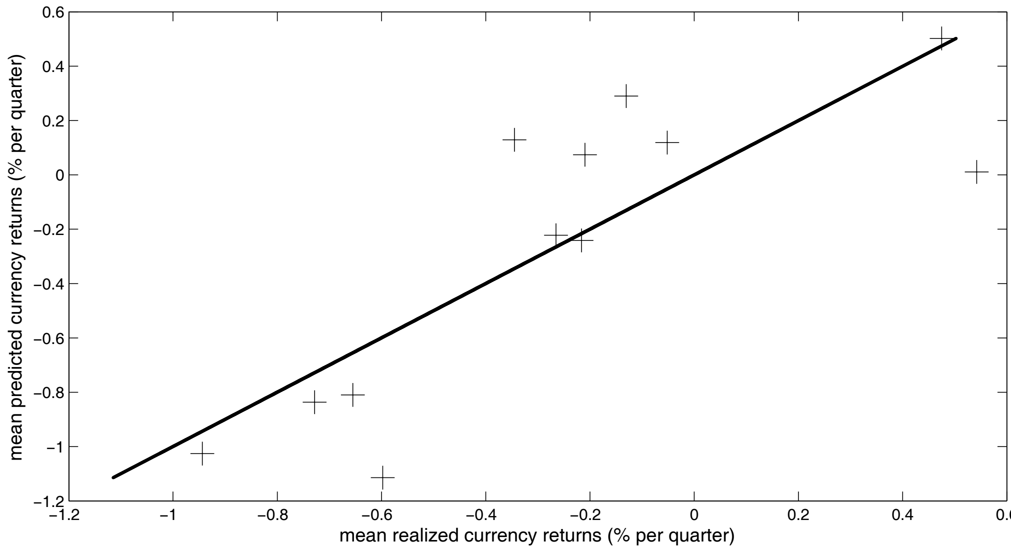
The estimated price of world consumption growth risk is statistically indistinguishable from zero. Though average consumption growth risk seems to be helpful in pricing currencies, it is not priced in average currency returns.

Quite in contrast to this finding, the cross-sectional variation in consumption growth is a significant determinant of returns on currencies. We estimate its price to be around 0.14 percentage points per quarter, i.e. approximately 0.55 percentage points annually. In addition, the heterogeneous agent world consumption CCAPM explains about sixty percent of the cross-sectional dispersion in average returns on currencies. We are thus able to show that international consumption dispersion is an important factor in order to explain the cross-section of (nominal) exchange rate changes. Figure 4.1 plots mean actual currency returns

[22] Mankiw (1986) argues that positive and negative estimates are possible from a theoretical perspective.

against their counterparts predicted by the model to give a visual impression of the fit of the model.

Figure 4.1: Fit of the CCAPM with heterogeneity (currency returns)



Our results suggest that idiosyncratic consumption risk is crucial in order to explain the cross-sectional differences in average exchange rate changes and hence risk premia on foreign currencies. In addition, taking account of international consumption heterogeneity allows to price individual currency returns directly which is interesting given that the seminal results of Lustig and Verdelhan (2006) rely on a set of currency portfolios that are formed with respect to interest rate differentials. As we argue next, one way to interpret this sorting of currencies by the size of their current interest rate differential implicitly amounts to doing what we do here – using the cross-sectional variance of consumption as a pricing factor.

4.3.2.2 Inside the mechanism: consumption heterogeneity and UIP

According to UIP, the only difference in the interest rate of two risk-free assets at home and abroad should be the expected exchange rate change. However, we know since Fama (1984) that interest rate differentials often predict expected exchange rate changes with the wrong sign. Though still a basic ingredient of theoretical macroeconomic models, UIP fails to hold in the data.[23]

Lustig and Verdelhan (2006) show that the failure of UIP is an immediate implication of basic asset pricing theory. Currencies from countries with high interest rates relative to the U.S. are perceived as a relatively risky asset. Hence they have to offer relatively high returns. For a domestic investor, a high return on a currency means a foreign currency appreciation while UIP predicts a foreign currency depreciation for countries whose interest rate is high compared to the home country.

The idea that an extra risk premium for holding foreign assets explains the empirical failure of UIP has attracted considerable attention (see e.g. the survey in Engel, 1996). The risk premium on foreign bonds, ϕ_{t+1}^{k}, implied by UIP obeys

$$\phi_{t+1}^{k} = i_t^k - i_t - \Delta e_{t+1}^{k} \tag{4.20}$$

As argued by Bansal and Dahlquist (2000), basic asset pricing theory should apply to explain the individual UIP risk premia. We thus assess if our framework is useful to price this premium and restrict the sample to the G7 economies.[24]

Panel A of table 4.2 provides the results. Neither world consumption growth nor the cross-sectional dispersion of consumption growth rates seem to help to explain the cross-section of country-by-country UIP risk premia. Accordingly, risk price estimates for the two factors are insignificant. We thus corroborate Sarkissian (2003) who starts from the relation between forward and expected spot exchange rates to price the forward premium rate in a similar framework. Even though consumption heterogeneity seems to help to explain forward premium, risk prices are not statistically significant.

[23] Bansal and Dahlquist (2000) document that UIP holds relatively well among high inflation countries

[24] We had to limit the sample either in the time or in the country dimension due to the limited availability of short-term interest rates.

Table 4.2: GMM estimates and risk prices

	Panel A: $i_t^* - i_t - \Delta e_{t+1}$		
	b_1	b_2	J-Test
GMM	17.72 (1.80)	183.65 (1.72)	0.91
	λ_1	λ_2	R^2
Fama-MacBeth	0.64 (0.74)	0.02 (0.46)	0.68
	Panel B: $\alpha(i_t^* - i_t)$		
	b_1	b_2	J-Test
GMM	**22.77** **(3.16)**	**386.75** **(7.02)**	0.90
	λ_1	λ_2	R^2
Fama-MacBeth	0.01 (0.01)	**0.17** **(3.53)**	0.91
	Panel C: ε_{t+1}		
	b_1	b_2	J-Test
GMM	**23.04** **(4.32)**	**360.39** **(7.53)**	0.50
	λ_1	λ_2	R^2
Fama-MacBeth	-0.00 (-0.00)	0.00 (0.00)	-0.80

Notes: Panel A presents results of cross-sectional pricing exercises for UIP risk premia: $i_t^k - i_t - \Delta e_{t+1}^k$ where i_t^k denotes the foreign short-term interest rate, i_t is the U.S. interest rate and Δe_{t+1}^k represents changes of log U.S. dollar exchange rates defined as foreign currency over U.S. dollar.
Panel B reports cross-sectional pricing estimates for the components of exchange rates that are perfectly correlated with exchange rate changes while panel C displays estimates for the orthogonal complements of exchange rate changes obtained from $\Delta e_{t+1}^k = \mu + \alpha(i_t^k - i_t) + \varepsilon_{t+1}$. The sample spans the period from 1971Q1 to 2003Q2. Countries included are CND, FRA, GER, ITA, JPN, UK. For details on GMM and Fama-MacBeth cross-sectional asset pricing estimates consider the notes to table 4.1.

But then the question remains what actually drives our results displayed in table 4.1? Engel (1996) and Lustig and Verdelhan (2006) emphasize that the 'true' risk premium on investment in foreign bonds should be determined by the covariance of exchange rates with systematic sources of risk. Interest rates should play no role.

In order to answer this question, we split exchange rate changes into the part that is perfectly correlated with interest rate differentials and into the orthogonal complement unrelated to interest rate differentials by running the following regression

$$\Delta e_{t+1}^{k} = \mu + \alpha_k (i_t^k - i_t) + \varepsilon_{t+1}^{k} \tag{4.21}$$

We re-estimate our CCAPM first on the estimated $\alpha_k (i_t^k - i_t)$ and then, separately, on ε_{t+1}^{k}. With some abuse of language, we call the fitted value $\alpha_k (i_t^k - i_t)$ the UIP component of the exchange rate. We do not report our estimates of α_k since they largely corroborate the estimates provided by Bansal and Dahlquist (2000) for the countries under consideration here. The estimates are all negative and in some cases even bigger than unity in absolute values so that these fitted values may be better described as departure from UIP than as UIP component. Panel B of table 4.2 presents the results for $\alpha_k (i_t^k - i_t)$. Our consumption-based pricing model is extremely successful in pricing the mean returns on the UIP components of currencies. The model passes the test of overidentifying restrictions in the GMM framework while entertaining the same level of risk aversion as with pure exchange rate changes.

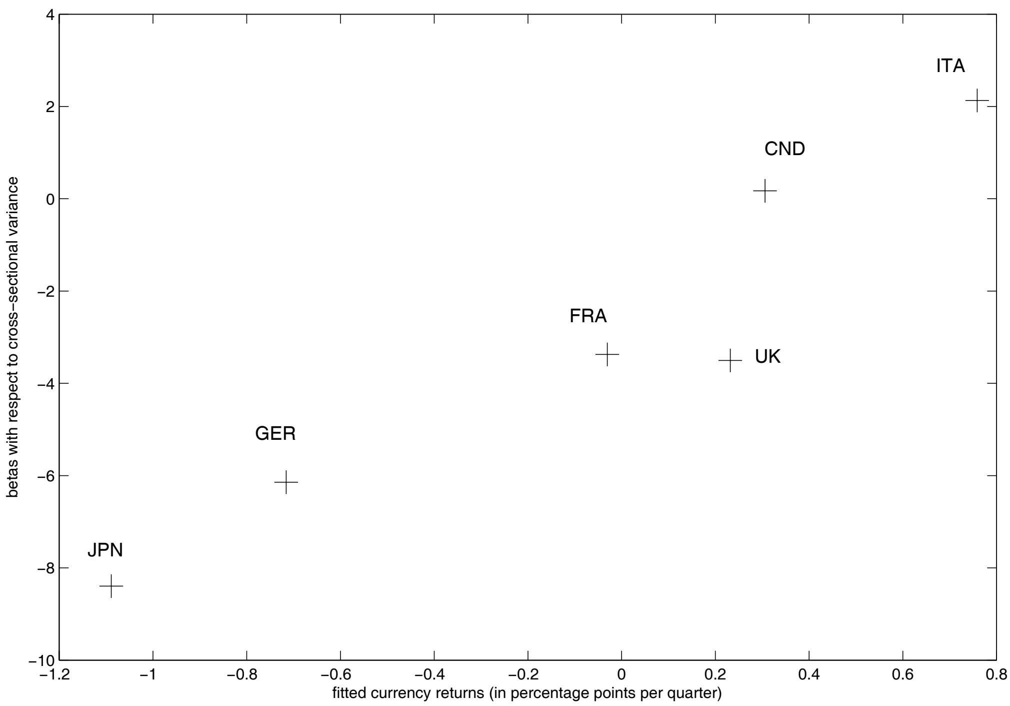
In line with our earlier results for exchange rate changes, the cross-sectional variance of world consumption growth seems to be the main determinant of the average UIP components in exchange rates. Its price is significant, positive and in the same range as the risk price for cash currency returns. Furthermore, the world CCAPM with heterogeneity explains 90 percent of the cross-sectional dispersion in the exchange rate components that are perfectly correlated with the interest rate differentials.

Panel C of table 4.2 displays the corresponding GMM and Fama-MacBeth estimates for the exchange rate components unrelated to interest rate differentials. The results clearly convey the notion that we fail to explain the cross-section of the currency components orthogonal to those predicted by UIP.

Hence, our CCAPM seems to describe the cross-sectional dispersion in currency returns so well because it captures cross-sectional differences in average exchange rate changes as predicted by departures from UIP. And ultimately, it seems to be the differences in the sensitivity of these UIP exchange rate components to idiosyncratic consumption risk that explain the success of our preferred asset pricing model.

Figure 4.2 gives a visual impression of the relation between mean currency returns perfectly correlated with interest rate differentials and their receptiveness to idiosyncratic consumption risk.

Figure 4.2. UIP components and idiosyncratic consumption risk betas



The clear positive relation reflects the fact that assets which are relatively strongly correlated with idiosyncratic, i.e. non-diversifiable, risk have to offer relatively high returns. This reasoning is absolutely in line with Lustig and Verdelhan (2006) who argue that currencies from high interest rate differential countries are perceived as relatively risky assets. We note that, unlike Lustig and Verdelhan, we are able to achieve this result without prior portfolio formation. Furthermore, we are now in the position to give a rationale for this finding. Currencies from countries with high interest rate differentials relative to the U.S. expose the average world investor to a fair amount of idiosyncratic consumption risk, which signals a risky investment.

Our results thus strongly suggest that the portfolio formation with respect to interest rate differentials in Lustig and Verdelhan (2006) summarizes the cross-sectional heterogeneity in consumption growth rates, which according to the mechanism considered here should play a vital role in explaining the cross-section of mean currency returns on a country-by-country basis. If consumption-based asset pricing models contain a grain of truth, then real interest rate differentials should mirror expected relative consumption growth rates as we have derived in (4.12). If we neglect inflation differentials for a moment, then we can write

$$i_t^k - i_t \approx -\gamma E_t(\Delta c_{t+1}^k - \Delta c_{t+1}) \tag{4.22}$$

Hence, if the econometrician estimates a model conditional on interest rate differentials, then he implicitly conditions on expected relative consumption growth rates. According to equation (4.20), the beta of the systematic component in exchange rate changes as given by $\alpha_k(i_t^k - i_t)$ should just capture the comovement of expected consumption growth in country k with world average consumption growth: given the relative variability of consumption growth rates, a high correlation of a country's consumption growth with world consumption growth implies that currencies from these countries will depreciate less or even appreciate when average world consumption growth is low. Given positive consumption correlations, currencies from countries with high volatility of consumption growth compared to world consumption growth offer a particularly good hedge against world consumption growth risk and hence have to offer relatively low returns.

The correlations of consumption growth of the G7 economies with world average consumption growth vary only in a relatively narrow range, between -0.1 for the UK and 0.3

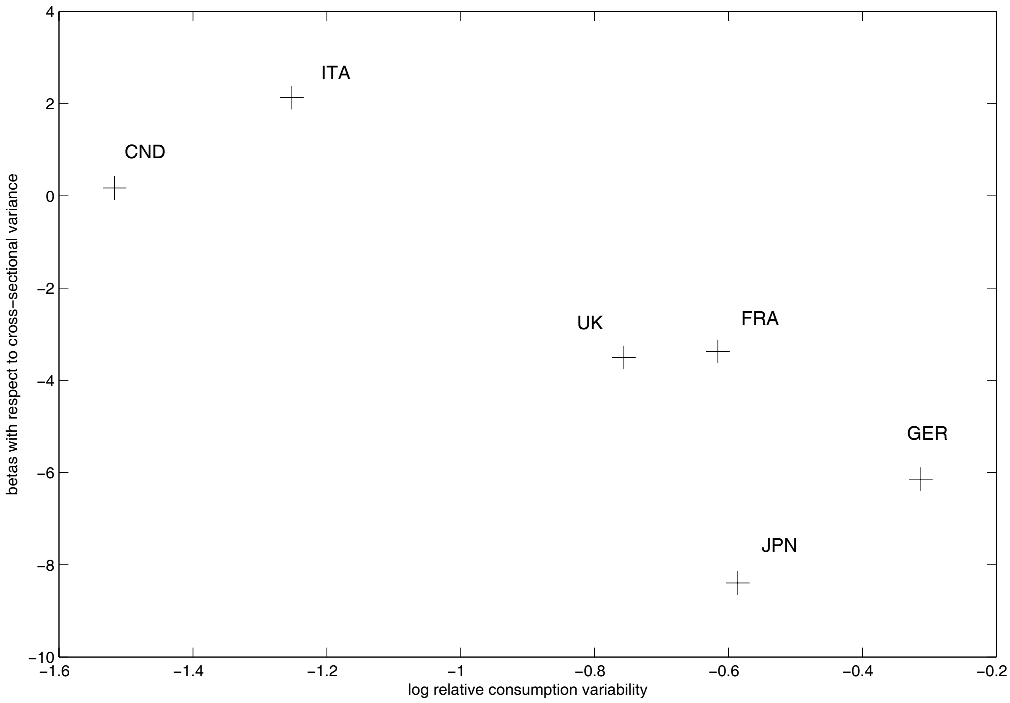
for France. This does not leave the impression that differences in the returns on the currencies of these countries are due to their *correlation* with world consumption growth. Our results support this casual observation; world consumption growth is virtually not priced in currency returns. Rather, given the (mainly) positive consumption correlations, it is the relative variability of consumption growth rates that determines the size of the risk premium on currencies.

This relative variability mechanism should also be reflected in the sensitivity of currency returns to the cross-sectional dispersion in consumption growth – the beta on world consumption dispersion. Figure 4.3 provides a visual impression of this relation.

We plot the logarithm of a country's variance of consumption growth relative to the variance of world consumption growth on the horizontal axis. The vertical axis displays the cross-sectional consumption variance betas of the UIP components of currency returns.

Figure 4.3: Idiosyncratic consumption risk betas (UIP components)

vs. log relative variance of consumption growth rates



High betas with respect to idiosyncratic consumption risk are associated with low relative consumption variability which is tantamount to saying that these currencies provide little insurance against world consumption growth risk. Note also that high cross-sectional consumption variance betas belong to the high interest rate differential currencies. Hence, we provide an empirical corroboration of the relative variability mechanism proposed by Lustig and Verdelhan (2006) for individual exchange rate changes. This mechanism seems to be key in explaining cross-sectional differences in returns on currencies. It is the immediate reflection of the fact that international consumption correlations are far from perfect. To the extent that this is the case, idiosyncratic consumption risk should be priced into currency returns. Our results suggest that it is.

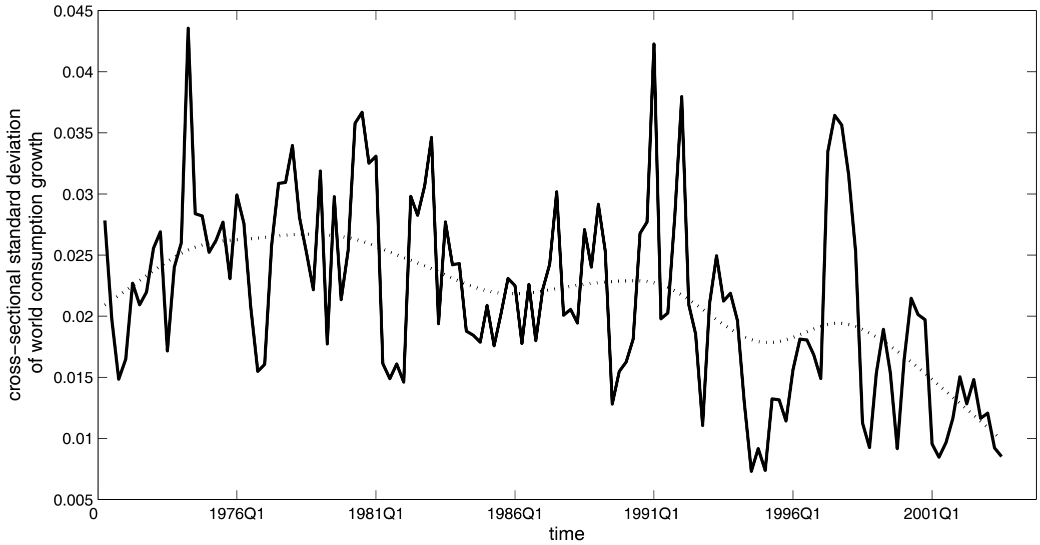
In the next section, we use our framework to provide an impression of how the impact of idiosyncratic consumption risk on exchange rate changes – as it is reflected in the risk price of consumption dispersion - has changed over time.

4.3.2.3 The price of idiosyncratic consumption risk over time

The impact of idiosyncratic consumption risk on currency returns is interesting from a macroeconomic and policy perspective: a growing literature documents that – in spite of two decades of continued financial globalization, international consumption correlations do not seem to have increased (see e.g. Heathcote and Perri, 2004; Artis and Hoffmann, 2004). Does this mean that risk sharing has not increased? Our framework suggests a simple way to get at this issue: if idiosyncratic consumption risk is bad for consumers, it should have a price. Then improvements in risk sharing should be reflected either in a decline in idiosyncratic risk or in a decline of the price of this risk.

Figure 4.4 presents the cross-sectional standard deviation of world consumption growth rates (solid line) as well as the Hodrick-Prescott filtered trend (dashed line) from 1971Q1 to 2003Q2.

Figure 4.4: Cross-sectional standard deviation of world consumption growth



This graph leaves the impression that the cross-sectional dispersion in consumption growth rates decreases over time. Hence, idiosyncratic consumption risk in the 1990s might have been of less importance than in the 1970s since its *level* has decreased over time. This observation inevitably raises the question if the *price* of idiosyncratic consumption risk has decreased over time as well.

We therefore estimate the factor risk prices when we regard simple currency returns on the full cross-section of 13 countries over two subsample periods. The first one ranges from 1970Q1 to 1989Q4 because the drop in the cross-sectional volatility in figure (4.4) at the end of 1980s or beginning of 1990s seems to be especially pronounced. Moreover, Lane and Milesi-Ferretti (2002, 2003, 2004) highlight the tremendous increase in cross-border capital flows since the middle of the 1990s which is suggestive of deeper market integration. The second period is from 1990Q1 to 2003Q2. Table 4.3 summarizes the results.

Table 4.3: Risk prices in subsamples

λ_1	λ_2	R^2
Panel A: 1970Q1-1989Q4		
0.16 (0.20)	**0.13 (2.43)**	0.35
Panel B: 1990Q1-2003Q2		
-0.63 (-0.83)	0.01 (0.40)	0.23

Notes: This table presents estimates of risk prices of average world consumption growth and the cross-sectional variance of consumption growth over subsamples. The risk prices are obtained with Fama-MacBeth cross-sectional regressions using exchange rate changes of the countries mentioned in table 1 as test assets.

Panel A shows that for the first sample period, the 1970s and 1980s, the price of idiosyncratic consumption growth is significant and about as high as in the full sample. It decreases in the 1990s and is not distinguishable from zero in the latter subperiod (Panel B).

All in all, these findings leave the impression that the lower the level of idiosyncratic consumption risk, the lower is also its price. Thus still low international consumption correlations are not necessarily associated with a lack of international risk sharing. Rather, our results suggest that the price of idiosyncratic consumption risk has become negligible in the course of the last 15-20 years, such that consumption risk sharing does no longer appear to be an urgent issue for consumers.

4.3.2.4 The cross-section of currency returns and the Backus-Smith puzzle

In section 4.2.1, we relate our asset pricing framework to the relation between real exchange rates and relative consumption growth as proposed in Backus and Smith (1993). We derive a similar but weaker condition, equation (4.14), that nonetheless constitutes a challenge for empirical research (e.g. Kollmann, 1995). Since this condition is tightly linked with the failure of UIP, we suggested a pricing kernel for exchange rate changes that comprises four factors: world average consumption growth, the Lustig-Verdelhan factor (the cross-sectional variance of nominal interest rates), the Backus-Smith factor (the cross-sectional variance of inflation) and a Fisher parity term (the covariance of interest rates with inflation). We assess the importance of each of the factors for the pricing of currency returns by estimating their prices.

Table 4.4 displays the results from Fama-MacBeth cross-sectional regressions when the full cross-section of currency returns is considered.

Table 4.4: Risk prices of Lustig-Verdelhan and Backus-Smith factor

	λ_1	λ_2	λ_3	λ_4	R^2
Fama-MacBeth	-0.17 (-0.40)	**0.0025** **(2.45)**	**0.01** **(3.03)**	-0.01 (-0.60)	0.59

Notes: Here, we present estimates of risk prices if we split the cross-sectional variance of consumption growth into its components suggested by the relationship between real interest rates and consumption growth implied by the Backus and Smith (1993) framework.
Our cross-sectional regression thus takes the following form:
$$\Delta e_{t+1}^{k} = \mu + \lambda_1 \beta_{\Delta \bar{c}_t}^{k} + \lambda_2 \beta_{\underset{k}{\operatorname{var}} i_t^* - \bar{i}_t}^{k} + \lambda_3 \beta_{\underset{k}{\operatorname{var}} p_{t+1}^* - \bar{p}_{t+1}}^{k} + \lambda_4 \beta_{\underset{k}{\operatorname{cov}}(i_t^* - \bar{i}_t, p_{t+1}^* - \bar{p}_{t+1})}^{k} + e_{t+1}, \forall t$$
The betas are obtained from time series regressions of exchange rate changes on average world consumption growth and the cross-sectional variance of short-term interest rates, inflation and the cross-sectional covariance of interest rates with inflation. T-statistics of the risk price estimates, λ, are corrected for generated regressor bias (Shanken, 1992).
The sample spans the period from the first quarter of 1971 to the second quarter of 2003 and uses data on exchange rates and consumption from the following countries: Australia, Austria, Canada, France, Germany, Italy, Japan, Norway, Spain, Sweden, Switzerland, United Kingdom and the United States.

World average consumption growth still is not a significant pricing factor. The cross-sectional variance of interest rates as well as inflation rates explain average exchange rate changes. Their risk price estimates are comfortably significantly different from zero whereas the Fisher parity term does not seem to be priced. These results show clearly that the Backus-Smith (BS) as well as the Lustig-Verdelhan (LV) factor jointly explain the behaviour of currency returns in the cross-section.

In the preceding subsection we have argued that relative nominal interest rates proxy for relative consumption growth rates which suggests to interpret the LV factor as mirror image of the relative variability mechanism that drives most of our results. But how to interpret the BS factor? We know from Backus and Smith (1993) that deviations from purchasing power parity should account for imperfect consumption correlations. Thus it is tempting to interpret the BS factor as reflection of the correlation mechanism proposed by Lustig and Verdelhan (2006) and outlined earlier in this chapter. Then, our findings suggest that i) the forward premium puzzle is driven by both dispersion in consumption correlations and the relative variability in consumption growth rates and ii) the failure of UIP as well as the lack of empirical support for the Backus-Smith condition are tightly linked.

Moreover, note that the explanatory power of this specification of the heterogeneous agent CCAPM for exchange rate changes is approximately the same as when we consider the cross-sectional consumption dispersion as pricing factor.

Furthermore, subsample analysis corroborates the results reported in section 4.3.2.3. Table 4.5 reports the results. The LV and the BS factor are statistically significant determinants of average currency returns in the period from 1971Q1 to 1989Q4 but not in the more recent period from 1990Q1 to 2003Q2.

Table 4.5: Risk prices of Lustig-Verdelhan and Backus-Smith factor (subsamples)

Panel A: 1971Q1 – 1989Q4					
	λ_1	λ_2	λ_3	λ_4	R^2
Fama-MacBeth	0.33 (1.10)	**0.0022** **(2.07)**	**0.01** **(2.62)**	-0.01 (-1.57)	0.46
Panel B: 1990Q1 – 2003Q2					
	λ_1	λ_2	λ_3	λ_4	R^2
Fama-MacBeth	-0.31 (-1.25)	0.0008 (1.12)	0.001 (1.20)	-0.00 (-0.31)	0.65

Notes: This table presents estimates of risk prices if we split the cross-sectional variance of consumption growth into its components suggested by the relationship between real interest rates and consumption growth implied by the Backus and Smith (1993) framework.
Our cross-sectional regression thus takes the following form:

$$\Delta e_{t+1}^{k} = \mu + \lambda_1 \beta_{\Delta \bar{c}_t}^{k} + \lambda_2 \beta_{\underset{k}{\mathrm{var}}\, i_t^* - \bar{i}_t}^{k} + \lambda_3 \beta_{\underset{k}{\mathrm{var}}\, p_{t+1}^* - \bar{p}_{t+1}}^{k} + \lambda_4 \beta_{\underset{k}{\mathrm{cov}}(i_t^* - \bar{i}_t,\, p_{t+1}^* - \bar{p}_{t+1})}^{k} + e_{t+1}, \forall t$$

The betas are obtained from time series regressions of exchange rate changes on average world consumption growth and the cross-sectional variance of short-term interest rates,

inflation and the cross-sectional covariance of interest rates with inflation. T-statistics of the risk price estimates, λ, are corrected for generated regressor bias (Shanken, 1992).
Panel A reports the results for the sample period from the first quarter of 1971 to the fourth quarter of 1989, panel B the estimates for the period from 1990Q1 to 2003Q2. We use data on exchange rates and consumption from the following countries: Australia, Austria, Canada, France, Germany, Italy, Japan, Norway, Spain, Sweden, Switzerland, United Kingdom and the United States.

One can think of the LV and BS factor as representing idiosyncratic risk on financial and goods markets respectively. The risk price estimates thus indicate that the importance of idiosyncratic risks on both markets has declined in the past two decades. This finding underscores our earlier drawn conclusion that the Backus-Smith puzzle and the lack of empirical support for UIP are closely related. Furthermore, it highlights that low international consumption correlations do not seem to signal low international risk sharing.

4.3.2.5 Robustness Checks

Before we conclude, we briefly report on a range of robustness checks with respect to alternative data specifications: we tested our CCAPM with heterogeneity using U.S. dollar exchange rate changes deflated with realized inflation. The results in panel A of table 4.6 reveal that the qualitative results remain unaltered.
The same is true if we change the base currency of our returns and regard Yen (panel B) as well as Deutschmark/Euro (panel C) returns. Furthermore, the estimate of the coefficient of relative risk aversion as well as the price of idiosyncratic consumption growth risk have the same order of magnitude as in the U.S. dollar case.

Table 4.6: robustness checks

Panel A: $\Delta e_t - \Delta p_t$ from U.S. perspective			
	b_1	b_2	J-Test
GMM	**25.04** **(4.70)**	**409.27** **(10.72)**	0.17
	λ_1	λ_2	R^2
Fama-MacBeth	0.88 (1.01)	**0.16** **(3.37)**	0.57
Panel B: Δe_t (Yen)			
	b_1	b_2	J-Test
GMM	**23.34** **(4.64)**	**366.31** **(8.41)**	0.33
	λ_1	λ_2	R^2
Fama-MacBeth	0.86 (0.99)	**0.13** **(2.24)**	0.71
Panel C: Δe_t (Deutschmark/Euro)			
	b_1	b_2	J-Test
GMM	**23.08** **(4.63)**	**362.42** **(8.28)**	0.36
	λ_1	λ_2	R^2
Fama-MacBeth	0.42 (0.46)	**0.13** **(2.99)**	0.61

Notes: The table provides estimates from two-stage GMM estimations of $m_t = 1 - b_1 \Delta \overline{c_t} - b_2 \operatorname{var}_k \Delta c_t^k$
where Δe_t represents either Yen or Deutschmark/Euro exchange rate changes, $\Delta \overline{c}_t$ average world consumption growth at time t and $\operatorname{var}_k \Delta c_t^k$ the cross-sectional consumption dispersion at time t. T-statistics occur in parenthesis. Furthermore, we present estimates of risk prices of the factors under consideration using the Fama-MacBeth cross-sectional regression.

4.4 Conclusions

In this chapter we have provided an asset pricing perspective on the link between real exchange rates and consumption. This link, first explored by Backus and Smith (1993), suggests that relative consumption growth rates should be inversely related to fluctuations in real exchange rates: a country´s consumption growth should be high when local prices are low and vice versa. While this link is central in most modern macro-models, it has found little support in the data. Here we have argued that the empirical failure of this link and the failure of uncovered interest parity may be two sides of the same medal.

In a first step we recognize that if consumption growth rates differ across countries, then this heterogeneity should be reflected in standard pricing relations for currency risk. We find that idiosyncratic consumption risk – measured by the cross-sectional variance of real consumption growth – explains more than 60 percent of the cross-sectional variation in exchange rates in a standard consumption CAPM adjusted for heterogeneous consumers. In a second step we further decompose idiosyncratic consumption risk into one component that is due to heterogeneity in inflation rates – a Backus-Smith factor – and another that reflects international heterogeneity in the price of intertemporal substitution and that is measured by the cross-sectional variance of interest rate differentials. We find both factors significant and successful in pricing the cross-section of currency returns and in particular in explaining departures from UIP.

As a policy application of our framework, we have examined how the price of idiosyncratic consumption risk has changed over time as financial globalization has accelerated in pace. We find that, even though international consumption correlations do not at first sight seem to have increased since the beginning of the 1990s, the cross-sectional dispersion of consumption risk has clearly been trending downward as has the world price of idiosyncratic consumption risk.

Summary

This thesis contributes to the growing literature that documents the tight link between financial markets and the macroeconomy in an international context.

Chapter one shows that the literature on stock return predictability has the potential to shed light on issues related to the long-term integration and comovement of national stock markets. Short-run fluctuations of the U.S. consumption-wealth ratio predict risk premia on foreign stock markets at the business cycle frequency. This evidence of predictability leaves the impression of an important, common temporary component in national stock markets. This common component helps to capture the covariation among the G7 stock markets that has increased tremendously during the last two decades.

Chapter two takes into question if European value stocks offer higher average returns than the corresponding growth stocks because of their sensitivity to news about future fundamentals of national market returns. Therefore, I exploit that a common, temporary component in national stock markets is mirrored in the explanatory power of the U.S. consumption-wealth ratio and other U.S. variables for foreign stock market returns. These variables are used to decompose the national market beta of European value and growth stock returns into a variety that mirrors news about the national market return's cashflows and into a variety that reflects discount rate news. Receptiveness to cashflow news should be rewarded with a higher price of risk than discount rate news since high expected discount rates decrease the present value of future fundamentals but at the same time reflect better investment opportunities. I find high average returns to be associated with relatively high cashflow components in market betas. Furthermore, two-beta versions of national CAPMs capture cross-sectional differences in European stock returns as implied by basic asset pricing theory. One national asset pricing model should explain all asset returns if capital markets are sufficiently integrated.

Investment in foreign securities exposes an investor to the risk that the exchange rate between the two countries could change. The question if basic asset pricing theory explains returns on foreign currencies and thus exchange rate changes in the cross-section is addressed in **chapter three**. I exploit recent evidence for an intimate link between stock return differentials and exchange rate changes. Shocks to relative stock returns persistently affect exchange rate changes such that lagged stock return differentials should be useful instruments in order to price currency returns. Consumption-based asset pricing underscores this intuition. Currencies from countries with past high stock returns compared to the U.S. are perceived as relatively riskless assets. Hence, they offer only low returns. I show that conditional on lagged stock return differentials the the CCAPM explains up to 50 percent of the cross-sectional variation in quarterly currency returns. This finding pertains to average returns on interest rate and stock return differential sorted currency portfolios as well as to U.S. dollar exchange rate changes on a country-by-country basis conditional on relative stock returns.

Recent evidence suggests that the failure of UIP is the immediate implication of basic asset pricing theory. Key to success is the formation of portfolios of currency excess returns, i.e. exchange rate changes less interest rate differentials, with respect to interest rate differentials. Currencies from countries with relatively high interest rates are perceived as risky investments that have to promise relatively high returns (a foreign currency appreciation as opposed to a depreciation predicted by UIP). In **chapter four**, we pay particular attention to the mechanisms that underlie the asset pricing based explanation for departures from UIP. We argue that forming currency portfolios with respect to interest rate differentials is conceptually very similar to exploiting the heterogeneity in international consumption growth rates. We show that a CCAPM that accounts for this heterogeneity successfully explains the cross-section of currency returns not only without portfolio formation but also without using conditional information in the form of instrumental variables. We use the heterogeneous agent framework to dissect the consumption-exchange rate anomaly. In a world with complete financial and goods markets and homothetic preferences, consumption growth rates should equalize. Deviations from purchasing power parity drive a wedge between consumption growth rates as shown by Backus and Smith (1993). At the same time, real exchange rates should be determined by relative consumption growth rates. However, empirical support for this condition is virtually absent. The correlations of relative consumption growth rates with real exchange rates are extremely low. We show that a somewhat weaker version of the Backus-Smith relation only holds when UIP holds. UIP is grossly violated in the data, but departures from UIP are captured by our model. According to Backus and Smith (1993) cross-sectional differences in price level changes (inflation) should explain cross-sectional differences in (real) exchange rate changes. We therefore decompose the cross-sectional variation in consumption into one component that is due to cross-sectional differences in inflation expectations and a second component that is due to international variation in nominal interest rates (the key to success in the Lustig and Verdelhan (2006) framework). Both factors account to virtually equal parts for the pricing power of the cross-sectional variance of consumption. Our results thus suggest that the failure of uncovered interest parity and the consumption real exchange rate anomaly are intimately linked. Furthermore, our main findings suggest that the level as well as the price of idiosyncratic consumption risk, measured as cross-sectional dispersion in consumption growth rates, has declined over time. This result highlights that low consumption correlations do not necessarily imply a lack of international risk sharing.

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Appendix Chapter one

A.1 Data sources and definitions :

U.S. household stock market wealth includes directly held equity shares at market value and indirectly held equity shares namely life insurance companies´ holdings, private pension fund holdings, state and local government as well as federal government fund holdings and households´ mutual fund holdings. Annual data is published in the supplemental table B.100e in the Z1 Flow of Funds Accounts of the Federal Reserve Board. I construct quarterly stock market wealth from Flow of Funds tables L.213 and L.214 to calculate foreign equity holdings at quarterly frequency.

U.S. household foreign equity holdings are calculated with help of Z1 Flow of Funds table L.213 which provides details about equity issues and holdings at market value. Corporate equity issues at market value include holdings of foreign issues by U.S. residents inclusive American Depositary Receipts. I assume that the share of rest-of-the-world equity holdings in total corporate equity holdings is the same as the share of rest-of-the-world equity holdings in U.S. households´ equity holdings because U.S. households hold either directly or indirectly roughly 90% of total corporate equity issues.

U.S. household domestic asset wealth is the difference between household net worth, Z1 Flow of Funds table B.100, line 42, and U.S. household foreign equity holdings.

U.S. consumption is consumption expenditure on non-durable goods and services excluding clothing and shoes published by the Bureau of Economic Analysis in NIPA table 2.3.5. Data on total personal consumption expenditure is also taken from NIPA table 2.3.5.

U.S. labour income is defined as in Lettau and Ludvigson (2001a, 2004). The data source is NIPA table 2.1 published by the Bureau of Economic Analysis.

Real variables are obtained by deflating with the CPI of total personal consumption expenditure in chain-weighted (2000 = 100) seasonally adjusted U.S. dollars published by the Bureau of Economic Analysis in NIPA table 2.3.4.

Per capita figures are obtained with population figures from NIPA table 2.1 published by the Bureau of Economic Analysis.

Returns on Morgan Stanley Capital International indexes are defined as natural logarithm of the respective index value at the end of period t+1 minus the natural logarithm of the index value at the end of period t. Excess returns are obtained by subtracting the risk-free rate from the returns. The risk-free rate is the 3-month U.S. treasury bill published by the Federal Board of Governor´s. Data on the the Morgan Stanley Capital International indexes is freely available at www.mscibarra.com. Since logarithmic approximations of net returns are used, the *h*-period return is the sum of one period returns over *h* periods.

Appendix Chapter three

Table A.1: Exchange rate changes and (conditional) consumption growth betas (stock return differentials)

country	constant	$\beta_{\Delta c}^{i}$	$\beta_{\Delta c,i}^{i}$
AUS	0.05	0.05	**−0.03**
	(0.78)	(0.61)	**(−1.99)**
AUT	0.05	0.33	0.01
	(0.48)	(1.68)	(0.40)
BEL	0.02	−0.0651	−0.0395
	(0.22)	(−0.4125)	(−1.7650)
CND	0.0006	−0.01	**−0.02**
	(0.0214)	(−0.16)	**(−3.27)**
DK	−0.06	0.01	**−0.05**
	(−0.57)	(0.04)	**(−3.35)**
FRA	0.09	−0.09	**−0.03**
	(1.30)	(−0.92)	**(−2.88)**
GER	0.07	0.08	−0.03
	(0.89)	(0.65)	(−1.75)
GRE	−0.62	2.18	−0.13
	(−0.42)	(0.77)	(−1.57)
IRL	−2.08	2.29	−0.12
	(−1.70)	(0.97)	(−0.66)
ITA	−0.03	0.06	**−0.05**
	(−0.24)	(0.29)	**(−2.49)**
JPN	−0.04	−0.02	−0.02
	(−0.52)	(−0.16)	(−1.31)
NL	0.04	−0.11	**−0.03**
	(0.85)	(−1.34)	**(−2.43)**
NZ	−1.04	−0.83	**−0.29**
	(−1.12)	(−0.52)	**(−2.88)**
NOR	0.20	−0.22	**−0.05**
	(1.31)	(−0.92)	**(−2.49)**
POR	−1.86	3.06	−0.08
	(−1.45)	(1.23)	(−0.82)
ESP	−0.02	0.13	**−0.05**
	(−0.25)	(0.83)	**(−2.55)**
SWE	0.07	0.03	0.00
	(1.03)	(0.35)	(0.51)

CH	0.02	0.07	− 0.01
	(0.34)	(0.56)	(−0.84)
UK	− 0.03	0.03	**− 0.03**
	(−1.42)	(0.78)	**(−3.79)**

Notes: This table presents results from regressions of the form

$$\Delta s_{t+1}^{i} = \mu + \beta_{\Delta c}\Delta c_{t+1} + \beta_{\Delta c,r}\Delta c_{t+1}(sr_t^i - sr_t) + \varepsilon_{t+1}$$

with Δs_{t+1}^{i} the U.S. dollar exchange rate change of country *i*, Δc_{t+1} log U.S. consumption growth, $sr_t^i - sr_t$ the stock return differential of country *i* relative to the U.S.

Table A.2: exchange rate changes and (conditional) consumption growth betas (interest rate differentials)

country	constant	$\beta^{i}_{\Delta c}$	$\beta^{i}_{\Delta c,i}$
AUS	0.06	0.02	0.03
	(0.86)	(0.26)	(0.33)
AUT	0.02	**0.42**	0.39
	(0.24)	**(2.02)**	(1.75)
BEL	− 0.02	− 0.01	0.06
	(−0.17)	(−0.08)	(0.44)
CND	0.01	− 0.07	0.10
	(0.51)	(−1.28)	(1.55)
DK	− 0.06	− 0.02	0.14
	(−0.65)	(−0.12)	(0.48)
FRA	0.08	**− 0.25**	**0.18**
	(1.20)	**(−2.09)**	**(2.52)**
GER	0.01	**0.34**	**0.45**
	(0.09)	**(2.08)**	**(2.22)**
GRE	− 0.58	−1.10	1.33
	(−0.44)	(−0.36)	(1.24)
IRL	− 2.16	1.73	2.12
	(−1.83)	(0.66)	(1.05)
ITA	− 0.02	− 0.37	0.29
	(−0.13)	(−1.78)	(1.66)
JPN	− 0.03	− 0.03	− 0.07
	(−0.44)	(−0.25)	(−0.69)
NL	0.04	− 0.11	0.10
	(0.68)	(−0.97)	(0.82)
NZ	−1.43	1.94	−1.69
	(−1.55)	(0.98)	(−0.95)
NOR	0.22	− 0.65	0.46
	(1.47)	(−1.69)	(1.18)
POR	−1.98	2.88	0.67
	(−1.51)	(1.13)	(0.48)
ESP	− 0.05	− 0.17	**0.39**
	(−0.74)	(−0.99)	**(1.98)**
SWE	0.07	− 0.01	0.07
	(1.00)	(−0.09)	(1.02)
CH	− 0.02	0.27	0.18
	(−0.22)	(1.34)	(1.50)

UK	− 0.00	− 0.03	− 0.00
	(−0.09)	(−0.58)	(−0.07)

Notes: This table presents results from regressions of the form

$$\Delta s^i_{t+1} = \mu + \beta_{\Delta c}\Delta c_{t+1} + \beta_{\Delta c,i}\Delta c_{t+1}(i^i_t - i_t) + \varepsilon_{t+1}$$

with Δs^i_{t+1} the U.S. dollar exchange rate change of country *i*, Δc_{t+1} log U.S. consumption growth, $i^i_t - i_t$ the interest differential of country *i* relative to the U.S.

Erklärung

Ich versichere hiermit, dass ich diese Dissertation selbständig verfasst habe. Bei der Erstellung der Arbeit habe ich mich ausschliesslich der angegebenen Hilfsmittel bedient. Die Dissertation ist nicht bereits Gegenstand eines erfolgreich abgeschlossenen Promotions- oder sonstigen Prüfungsverfahrens gewesen.

Zürich, den 9. März 2007