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Consumption, Wealth, Stock and Government Bond Returns: International Evidence

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Consumption, Wealth, Stock and Government Bond Returns: International Evidence[*]

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Abstract

In this paper, we show, from the consumer's budget constraint, that the residuals of the trend relationship among consumption, aggregate wealth, and labour income should predict both stock returns and government bond yields. We use data for several OECD countries and find that when agents expect future stock returns to be higher, they will temporarily allow consumption to rise. Regarding government bond yields, when bonds are seen as a component of asset wealth, then investors react in the same way. If, however, the increase in the yields is perceived as signalling a future rise in taxes, then they will temporarily reduce their consumption.

Contents

1. Introduction

Differences in expected returns across assets are explained by differences in risk, and the risk premium is generally considered as reflecting the ability of an asset to insure against consumption fluctuations (Sharpe, 1964). Despite this belief, a measure such as the covariance of returns across portfolios and contemporaneous consumption growth did not prove to be sufficient to explain the differences in expected returns (Breeden et al., 1989). In fact, the literature on asset pricing has concluded that inefficiencies of financial markets[1] and the rational response of agents to time-varying investment opportunities[2] help justifying why expected excess returns appear to vary with the business cycle.

In addition, different macro-financially motivated variables that capture time-variation in expected returns have been developed. For instance: the consumption-wealth ratio (Lettau and Ludvigson, 2001); the long-run risk (Bansal and Yaron, 2004; Bansal et al., 2005); the labour income risk (Julliard, 2004); the housing collateral risk (Lustig and van Nieuwerburgh, 2005);[3] the ultimate consumption risk (Parker and Julliard, 2005); and the composition risk (Yogo, 2006; Piazzesi et al., 2007). Additional variables include the (adjusted) dividend or cash flow yield (Goyal and Welsh, 2003; Robertson and Wright, 2006; Boudoukh et al., 2007); the ratio of excess consumption (i.e. consumption in excess of labour income) to observable assets (Whelan, 2008); and the wealth composition risk (Sousa, 2010a).[4]

In contrast with the literature on the predictability of stock returns, only a few studies tried to explain the factors behind sovereign bond risk premia. Among these: Fama and Bliss (1987) focus on the spread between the *n*-year forward rate and the one-year yield; Campbell and Shiller (1991) emphasize the Treasury yield spreads; Campbell and Cochrane (1999) and Wachter (2006) stress the role of shocks to aggregate consumption, while Brandt and Wang (2003) address the importance of shocks to inflation; Cochrane and Piazzesi (2005) highlight a single factor, a single tent-shaped linear combination of forward rates; and Ludvigson and Ng (2009) find marked countercyclical variation in bond risk premia.

The current paper argues that the wealth and macroeconomic data can be combined to address the issue of predictability of asset returns. More specifically, we derive an

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[1] See Fama (1998), Fama and French (1996), and Farmer and Lo (1999).
[2] See Sundaresan (1989), Constantinides (1990), Campbell and Cochrane (1999), Duffee (2005), and Santos and Veronesi (2006).
3 In the same spirit, Sousa (2007a) shows that housing can be used as an hedge against wealth shocks.
[4] Sousa (2007b) provides a comparison of the forecasting power of several empirical proxies, but distinguishes between their expected and unexpected components.

equilibrium relation between the transitory deviation from the common trend in consumption, aggregate wealth and labour income, labelled as *cay* (*cday* when disaggregate wealth is used) and stock returns as well as government bond yields.

The above-mentioned common trends summarize agent's long-term expectations of stock returns, government bond yields and/or consumption growth. In particular, when forward-looking investors expect future stock returns to be higher, they will allow consumption to rise above its common trend with aggregate wealth and labour income. As in Lettau and Ludvigson (2001) and Sousa (2010a), in this way, investors insulate future consumption from fluctuations in stock returns. As for government bond returns, one needs to understand the way government debt is perceived by agents. If government bonds are seen as a component of asset wealth, then investors allow consumption to rise above its equilibrium relationship with aggregate wealth and labour income when they have expectations of higher government bond yields. However, if the issuance of government debt is seen as a symptom of deteriorating public finances, then investors will allow consumption to fall below its common trend with aggregate wealth and labour income.

Using data for a set of 16 OECD countries, we show that the predictive power of the *cay* and *cday* measures for real stock returns is particularly important for horizons from three to four quarters for Australia, Belgium, Canada, Denmark, Finland, the UK and the US. As for Germany, Ireland, and Spain, those proxies do not seem to track well time-variation in stock returns.

In what concerns government bond returns, our analysis suggests that we can cluster the countries in two groups. In the first group (which includes Australia, Finland and the Netherlands), both *cay* and *cday* have an associated coefficient with positive sign in the forecasting regressions. Therefore, this corroborates the idea that government debt is a component of asset wealth. In the second group (which includes Canada, Denmark, Germany, Ireland, Italy and the US), the forecasting regressions show that both *cay* and *cday* have an associated negative coefficient. Consequently, agents in these countries understand the rise in government bond returns rather as signalling an increase in future taxation.

Finally, assessing the robustness of our results, we show that: (i) additional control variables do not change the predictive power of *cay* and *cday*; and (ii) models that include *cay* and *cday* perform better than other benchmark models.

The paper is organized as follows. Section two describes the theoretical framework and presents the empirical methodology. Section three provides the estimation results of

the forecasting regressions for asset returns. Section four provides the robustness analysis. Section five concludes.

2. Theoretical framework and empirical methodology
2.1 Theory

Consider the case of a representative consumer. The budget constraint can be written as

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

where W_t represents aggregate wealth, C_t denotes private consumption, and $R_{w,t+1}$ corresponds to the return on aggregate wealth between period t and $t+1$.

Campbell and Mankiw (1989) show that, under the assumption that the consumption-aggregate wealth is stationary and that $\lim_{i \to \infty} \rho_w{}^i (c_{t+i} - w_{t+i}) = 0$, [5] one can approximate equation (1) by a Taylor expansion as follows

$$c_t - w_t = \sum_{i=1}^{\infty} \rho_w^i r_{w,t+i} - \sum_{i=1}^{\infty} \rho_w^i \Delta c_{t+i} + k_w, \tag{2}$$

where $c \equiv logC$, $w \equiv logW$, and k_w is a constant. This equation shows that deviations of consumption from its equilibrium relationship with aggregate wealth can reflect changes in the returns on aggregate wealth or in consumption growth.

The aggregate return on wealth can be decomposed as

$$R_{w,t+1} = \omega_t R_{a,t+1} + (1 - \omega_t) R_{h,t+1}, \tag{3}$$

where ω_t is a time varying coefficient and $R_{a,t+1}$ is the return on asset wealth. Campbell (1996) shows that the last expression can be approximated as

$$r_{w,t} = \omega_t r_{a,t} + (1 - \omega_t) r_{h,t} + k_r, \tag{4}$$

where k_r is a constant, and $r_{w,t}$ is the log return on asset wealth. Similarly, the log total wealth can be approximated as

$$w_t = \omega a_t + (1 - \omega) h_t + k_a, \tag{5}$$

where a_t is the log asset wealth, h_t is the log human wealth, ω is the mean of ω_t, and k_a is a constant.

Campbell (1996) and Jagannathan and Wang (1996) suggest that labour income, Y_t, can be thought as the dividend on human capital, H_t. Consequently, the return to human capital can be defined as:

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[5] The authors assume that $\rho_w \equiv (W - C)/W < 1$.

$$1 + R_{h,t+1} = \frac{H_{t+1} + Y_{t+1}}{H_t}. \tag{6}$$

Log-linearizing this relation around the steady state,[6] one gets

$$r_{h,t+1} = (1 - \rho_h)k_h + \rho_h(h_{t+1} - y_{t+1}) - (h_t - y_t) + \Delta y_{t+1}, \tag{7}$$

where $r \equiv log(1+R)$, $h \equiv logH$, $y \equiv logY$, k_h is a constant of no interest, and the variables without time subscript are evaluated at their steady state value. Assuming that $\lim_{i \to \infty} \rho_h^{\,i}(h_{t+i} - y_{t+i}) = 0$, the log human capital income ratio can be rewritten as a linear combination of future labour income growth and future returns on human capital:

$$h_t - y_t = \sum_{i=1}^{\infty} \rho_h^{i-1}(\Delta y_{t+i} - r_{h,t+i}) + k_h. \tag{8}$$

Replacing equation (4), (7) and (8) into (2), one obtains

$$c_t - \omega a_t - (1-\omega)(y_t + \sum_{i=1}^{\infty} \rho_h^{i-1}\Delta y_{t+i}) =$$

$$= \omega \sum_{i=1}^{\infty} \rho_w^i r_{a,t+i} + (1-\omega)\sum_{i=1}^{\infty} (\rho_w^i - \rho_h^{i-1})r_{h,t+i} - \sum_{i=1}^{\infty} \rho_w^i \Delta c_{t+i} + k, \tag{9}$$

where k is a constant. This equation holds ex-post as a direct consequence of agent's budget constraint, but it also has to hold ex-ante. Taking time t conditional expectation of both sides gives

$$\underbrace{c_t - \omega a_t - (1-\omega)y_t}_{cay_t} = \omega E_t \sum_{i=1}^{\infty} \rho_w^i r_{a,t+i} + (1-\omega)E_t \sum_{i=1}^{\infty} \rho_h^{i-1}\Delta y_{t+i} - E_t \sum_{i=1}^{\infty} \rho_w^i \Delta c_{t+i} + \eta_t + k,$$

$$\tag{10}$$

where $\eta_t \equiv (1-\omega)\sum_{i=1}^{\infty} (\rho_w^i - \rho_h^{i-1})r_{h,t+i}$, is a stationary component.

When the left hand side of equation (10) is high, consumers expect high future returns on market wealth. Based on that equation, cay_t may carry relevant information about market expectations of future asset returns, $r_{a,t+i}$, future labour income growth, Δy_{t+i}, and future consumption growth, Δc_{t+i}. In particular, when future stock returns are expected to be higher, forward-looking investors will allow consumption to rise above its common trend with aggregate wealth and labour income. In what concerns government bond returns: (i) investors behave in the same fashion when government bonds are seen as

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6 This is true under the assumption that the steady state human capital-labour income ratio is constant, that is, $Y/H = \rho_h^{-1} - 1$, where $0 < \rho_h < 1$.

a component of asset wealth; but (ii) investors will reduce consumption when higher government bond returns are perceived as signalling a deterioration of public finances.

Finally, Sousa (2010a) highlights the importance of wealth composition in pricing risk premium.[7] By disaggregating wealth, a_t, into its major components (financial wealth, f_t, housing wealth, u_t) and aggregate returns, $r_{a,t}$, into returns on financial assets, $r_{f,t}$, and returns on housing assets, $r_{u,t}$, one can link the trend deviation $cday_t$ with the market expectations of future financial and housing asset returns as follows:

$$\underbrace{c_t - \omega_f f_t - \omega_u u_t - (1 - \omega_f - \omega_u) y_t =}_{cday_t}$$

$$= \omega_f E_t \sum_{i=1}^{\infty} \rho_w^i r_{f,t+i} + \omega_u E_t \sum_{i=1}^{\infty} \rho_w^i r_{u,t+i} + (1 - \omega_f - \omega_u) E_t \sum_{i=1}^{\infty} \rho_h^{i-1} \Delta y_{t+i} - E_t \sum_{i=1}^{\infty} \rho_w^i \Delta c_{t+i} + \eta_t + k,$$

$$(11)$$

where $\eta_t \equiv (1 - \omega_f - \omega_u) \sum_{i=1}^{\infty} (\rho_w^i - \rho_h^{i-1}) r_{h,t+i}$.[8]

2.2. Long-run relationship between consumption, wealth and income

In order to assess the long-run relationship between consumption, (dis)aggregate wealth and labour income, we start by using the augmented Dickey-Fuller and the Phillips-Perron tests. This allows one to determine the existence of unit roots in the series and the tests suggest that all series are first-order integrated, *I(1)*. Next, we analyze the existence of cointegration among the series, using the methodology of Engle-Granger and Phillips-Ouliaris, and find evidence that supports that hypothesis. Finally, we estimate the trend relationship among consumption, wealth and labour income following Davidson and Hendry (1981) and Blinder and Deaton (1985).

Since the impact of different assets' categories on consumption can vary (Poterba and Samwick, 1995; Sousa, 2008), we also disaggregate wealth into its main components: financial wealth and housing wealth. Following Stock and Watson (1993), we use a dynamic least squares (DOLS) technique, specifying the following equation

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[7] Sousa (2010b) also shows that monetary policy can indeed have a strong impact on the composition of wealth in the euro area as a whole.

[8] From a theoretical point of view, some authors argue that housing wealth effects should be small. For instance, Buiter (2008) sustains that an increase in value of housing leads to higher housing consumption costs, which offset the housing wealth effect on non-housing consumption. Muellbauer (2008) suggests that the positive effect on non-housing consumption from an increase in housing prices is counterbalanced by a fall on housing consumption. Calomiris et al. (2009) emphasize that changes in housing wealth are typically correlated with changes in expected permanent income.

$$c_t = \mu + \beta_a a_t + \beta_y y_t + \sum_{i=-k}^{k} b_{a,i} \Delta a_{t-i} + \sum_{i=-k}^{k} b_{y,i} \Delta y_{t-i} + \varepsilon_t \qquad (12)$$

where the parameters β_a and β_y represent, respectively, the long-run elasticities of consumption with asset wealth and labour income, Δ denotes the first difference operator, μ is a constant, and ε_t is the error term.

In the estimation of the long-run relationships among consumption, (dis)aggregate wealth and labour income, we use quarterly data, post-1960, for 16 countries (Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, Spain, Sweden, the UK, the US).

The consumption data are the private consumption expenditure and were taken from the database of the NiGEM model of the NIESR Institute, the Main Economic Indicators of the OECD and DRI International. The labour income data correspond to the compensation series of the NIESR Institute. In the case of the US, labour income series was constructed following Lettau and Ludvigson (2004). The wealth data were taken from the national central banks or Eurostat. For the G-8 countries, the wealth series were compared with alternative sources, namely, Bertaut (2002), Pichette and Tremblay (2003), Tan and Voss (2003), Catte et al. (2004), and the Bank of Japan.

The stock return data were computed using the share price index provided by the International Financial Statistics of the IMF and the dividend yield ratio provided by Datastream. The 10-year government bond yield data were obtained from the International Financial Statistics of the IMF.

The government finance data normally refers to the Central Government, therefore, with the exclusion of the Local and/or the Regional Authorities. It is typically disseminated through the monthly publications of the General Accounting Offices, Ministries of Finance, National Central Banks and National Statistical Institutes of the respective countries. The latest figures are also published in the Special Data Dissemination Standard (SDDS) section of the International Monetary Fund (IMF) website.

Finally, the population series were taken from the OECD's Main Economic Indicators and interpolated (from annual data), and all series were deflated with consumption deflators and expressed in logs of per capita terms. The series were seasonally adjusted using the X-12 method where necessary and the time frames were chosen based on the availability of reliable data for each country.

Table 1.1 shows the estimates for the shared trend among consumption, asset wealth, and income. It can be seen that, despite some heterogeneity, the long-run elasticities of consumption with respect to aggregate wealth and labour income imply roughly a share of one third for asset wealth and two thirds for human wealth in aggregate wealth, in accordance with the values that one would expect in a production function with Cobb-Douglas technology. This is particularly true for Australia, Canada, Finland, France, Ireland, the UK and the US. Moreover, the disaggregation between asset wealth and labour income is statistically significant for all countries (with the exceptions of Finland and Italy). The table also presents the unit root tests to the residuals of the cointegration relationship based in the Engle and Granger (1987) methodology and reveals their stationarity.

Table 1.1 - Long-run relationship between consumption, aggregate wealth, and labour income. $cay_t = c_t - \beta_1 a_t - \beta_2 y_t$.

	a	Y	ADF t-statistic	Critical values	
			Lags: 1	5%	10%
Australia	0.35***	0.54***	-1.45	-1.95	-1.61
	(13.39)	(8.03)			
Austria	-0.08***	1.46***	-2.30	-1.95	-1.61
	(-5.10)	(23.48)			
Belgium	0.16***	0.56***	-4.53	-1.95	-1.61
	(8.02)	(13.01)			
Canada	0.36***	0.56***	-2.27	-1.95	-1.61
	(13.16)	(10.82)			
Denmark	0.09***	0.65***	-1.88	-1.95	-1.61
	(6.12)	(19.10)			
Finland	0.38***	0.13	-1.78	-1.95	-1.61
	(6.88)	(0.98)			
France	0.25***	0.55***	-2.09	-1.95	-1.61
	(16.95)	(18.03)			
Germany	0.13*	1.16***	-1.64	-1.95	-1.61
	(1.71)	(35.01)			
Ireland	0.36***	0.46***	-3.33	-1.95	-1.61
	(9.17)	(10.03)			
Italy	-0.02	1.49***	-1.07	-1.95	-1.61
	(-0.20)	(11.32)			
Japan	0.08***	0.89***	-3.27	-1.95	-1.61
	(3.74)	(25.99)			
Netherlands	0.17***	0.53***	-3.00	-1.95	-1.61
	(12.92)	(10.30)			
Spain	0.06*	0.76***	-2.19	-1.95	-1.61
	(1.67)	(16.10)			
Sweden	-0.13**	1.12***	-2.06	-1.95	-1.61
	(-2.45)	(9.06)			
UK	0.32***	0.66***	-2.45	-1.95	-1.61
	(13.84)	(12.84)			
US	0.28***	0.79***	-2.90	-1.95	-1.61
	(17.14)	(35.75)			

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

Table 1.2 reports the estimates of the long-run elasticities of consumption with respect to financial wealth, f, housing wealth, u, and labour income. First, it shows that the disaggregation between financial and housing wealth is statistically significant for almost

all countries. Moreover, consumption is, in general, more sensitive to financial wealth than to housing wealth, as the elasticities of consumption with respect to financial wealth are larger in magnitude. Second, it tells us that consumption is very responsive to financial wealth in the case of Belgium (0.11), Canada (0.30), Finland (0.14), Germany (0.31), Italy (0.24), Sweden (0.12) and the UK (0.17). Third, the long-run elasticity of consumption with respect to housing wealth is particularly strong for Australia (0.27), France (0.10), Ireland (0.13) and the Netherlands (0.10). Finally, the cointegration tests suggest that the residuals of the common trend between consumption, financial wealth, housing wealth and labour income are stationary.

Table 1.2 - Long-run relationship between consumption, financial wealth, housing wealth, and labour income. $cday_t = c_t - \beta_1 f_t - \beta_2 u_t - \beta_3 y_t$.

	f	U	y	ADF t-statistic	Critical values	
				Lags: 1	5%	10%
Australia	0.07***	0.27***	0.59***	-1.66	-1.95	-1.61
	(10.26)	(9.63)	(10.44)			
Austria	-0.05***	-0.02	1.44***	-2.31	-1.95	-1.61
	(-2.67)	(-1.21)	(19.10)			
Belgium	0.11***	-0.06**	0.80***	-3.73	-1.95	-1.61
	(8.98)	(-2.13)	(11.45)			
Canada	0.30***	0.06***	0.49***	-2.77	-1.95	-1.61
	(14.43)	(2.98)	(11.37)			
Denmark	0.02**	0.02	0.71***	-1.83	-1.95	-1.61
	(2.45)	(0.84)	(18.88)			
Finland	0.14***	-0.04	0.69***	-1.73	-1.95	-1.61
	(12.09)	(-1.00)	(6.53)			
France	0.08***	0.10***	0.62***	-2.32	-1.95	-1.61
	(17.22)	(4.23)	(22.74)			
Germany	0.31***	0.09***	0.33***	-2.66	-1.95	-1.61
	(22.10)	(3.41)	(9.60)			
Ireland	0.13***	0.13***	0.53***	-4.11	-1.95	-1.61
	(9.11)	(3.06)	(9.97)			
Italy	0.24***	-0.03	0.74***	-2.87	-1.95	-1.61
	(15.87)	(-1.10)	(9.57)			
Japan	0.07***	-0.03	0.86***	-3.27	-1.95	-1.61
	(3.25)	(-1.52)	(13.95)			
Netherlands	0.08***	0.10***	0.53***	-3.04	-1.95	-1.61
	(14.80)	(0.02)	(7.39)			
Spain	0.08***	0.02	0.67***	-2.58	-1.95	-1.61
	(5.60)	(0.93)	(13.80)			
Sweden	0.12***	-0.15***	0.61***	-3.43	-1.95	-1.61
	(10.28)	(-8.11)	(9.43)			
UK	0.17***	0.06***	0.76***	-3.94	-1.95	-1.61
	(18.87)	(2.99)	(16.14)			
US	0.04***	-0.02	1.21***	-2.55	-1.95	-1.61
	(2.66)	(-0.46)	(22.53)			

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

3. Empirical results

3.1. Forecasting stock returns

Equation (10) shows that transitory deviations from the long-run relationship among consumption, aggregate wealth and income, cay_t, mainly reflect agents' expectations of future changes in asset returns.

Moreover, given the disaggregation of asset wealth into its main components (financial and housing wealth), we argue that $cday_t$ should provide a better forecast for asset returns than a variable like cay_t in Lettau and Ludvigson (2001). We look at real stock returns (denoted by SR_t) for which quarterly data are available and should provide a good proxy for the non-human component of asset wealth.

Table 2.1 summarizes the forecasting power of cay_t at different horizons. It reports estimates from OLS regressions of the H-period real stock return, $SR_{t+1} + \ldots + SR_{t+H}$, on the lag of cay_t.

Table 2.1 – Forecasting real stock returns: estimated effect of *cay*.
$SR_{t+1} + SR_{t+2} + \ldots + SR_{t+H} = f(cay_{t-1})$, H=1, 2, 3, 4, 8.

	Forecast Horizon H						Forecast Horizon H				
	1	2	3	4	8		1	2	3	4	8
Australia	0.64*	1.07**	1.44**	1.88***	3.03***	Ireland	0.32	0.21	-0.26	-0.51	-1.06
	(1.73)	(2.02)	(2.35)	(2.61)	(3.71)		(0.56)	(0.22)	(-0.19)	(-0.33)	(-0.57)
	[0.04]	[0.05]	[0.06]	[0.07]	[0.10]		[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Austria	0.51	1.01	1.45*	1.99*	2.92*	Italy	0.23	0.43	0.67	0.91*	1.69**
	(1.42)	(1.59)	(1.66)	(1.87)	(1.82)		(0.97)	(1.23)	(1.59)	(1.82)	(2.48)
	[0.02]	[0.03]	[0.03]	[0.04]	[0.04]		[0.01]	[0.01]	[0.02]	[0.02]	[0.04]
Belgium	1.56***	2.72***	3.39***	3.31***	-0.12	Japan	0.86**	1.40**	1.58**	1.76*	-0.51
	(4.37)	(4.31)	(3.63)	(2.94)	(-0.07)		(2.32)	(2.31)	(1.95)	(1.84)	(-0.33)
	[0.11]	[0.11]	[0.12]	[0.08]	[0.00]		[0.05]	[0.05]	[0.04]	[0.04]	[0.00]
Canada	0.72***	1.27***	1.51***	1.50***	0.68	Netherlands	0.65*	1.15**	1.91***	2.62***	3.47***
	(3.31)	(3.55)	(3.08)	(2.52)	(0.95)		(1.80)	(2.23)	(2.86)	(3.37)	(2.78)
	[0.07]	[0.11]	[0.07]	[0.17]	[0.01]		[0.02]	[0.02]	[0.04]	[0.05]	[0.04]
Denmark	0.38***	0.82***	1.31***	1.77***	3.57***	Spain	0.27	-0.02	0.35	0.47	-2.06
	(2.84)	(3.58)	(4.09)	(4.43)	(5.24)		(0.25)	(-0.02)	(0.28)	(0.29)	(-1.10)
	[0.07]	[0.11]	[0.14]	[0.15]	[0.25]		[0.00]	[0.00]	[0.00]	[0.00]	[0.01]
Finland	0.74**	1.67***	2.67***	3.64***	6.30***	Sweden	0.05	0.09	-0.04	-0.41	-1.81*
	(2.09)	(3.06)	(3.83)	(4.16)	(4.87)		(0.19)	(0.19)	(-0.07)	(-0.59)	(-1.68)
	[0.04]	[0.07]	[0.11]	[0.14]	[0.18]		[0.00]	[0.00]	[0.00]	[0.00]	[0.04]
France	1.51***	3.24***	4.75***	6.05***	10.51***	UK	0.94***	1.88***	2.66***	3.27***	3.42***
	(2.76)	(3.94)	(4.85)	(5.34)	(7.00)		(3.81)	(4.19)	(4.41)	(4.54)	(4.27)
	[0.07]	[0.13]	[0.18]	[0.20]	[0.29]		[0.09]	[0.15]	[0.15]	[0.18]	[0.12]
Germany	-0.25	-0.47	-0.68	-0.94	-1.85**	US	0.71**	1.55***	2.21***	2.77***	5.69***
	(-0.99)	(-1.20)	(-1.33)	(-1.55)	(-2.36)		(2.24)	(2.89)	(3.13)	(3.24)	(4.86)
	[0.01]	[0.01]	[0.02]	[0.02]	[0.04]		[0.03]	[0.06]	[0.07]	[0.08]	[0.15]

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

It shows that cay_t is statistically significant for almost all countries and the point estimate of the coefficient is large in magnitude. Moreover, its sign is positive. These results are in line with the theoretical framework presented in Section 3, suggesting that investors will temporarily allow consumption to rise above its equilibrium level in order to smooth it and insulate it from an increase in real stock returns. Therefore, deviations in the long-term trend among c_t, a_t and y_t should be positively related to future stock returns.

Moreover, they explain an important fraction of the variation in future real returns (as described by the adjusted R-square), in particular, at horizons spanning from three to four quarters. In fact, at the four quarter horizon, cay_t explains 20% (France), 18% (UK), 17% (Canada), 15% (Denmark), 14% (Finland), 8% (Belgium and US) and 7% (Australia)

of the real stock return. In contrast, its forecasting power is poor for countries such as Germany, Ireland, Spain and Sweden.

Table 2.2 summarizes the forecasting power of $cday_t$ at different horizons. It reports estimates from OLS regressions of the H-period real stock return, SR_{t+1} + ... + SR_{t+H}, on the lag of $cday_t$.

Table 2.2 – Forecasting real stock returns: estimated effect of *cday*.
$SR_{t+1} + SR_{t+2} + ... + SR_{t+H} = f(cday_{t-1})$, H=1, 2, 3, 4, 8.

	Forecast Horizon H						Forecast Horizon H				
	1	2	3	4	8		1	2	3	4	8
Australia	0.66*	1.08*	1.53**	2.18**	3.72***	Ireland	0.37	0.34	-0.08	-0.30	0.24
	(1.67)	(1.92)	(2.19)	(2.47)	(3.55)		(0.63)	(0.35)	(-0.06)	(-0.20)	(0.11)
	[0.03]	[0.04]	[0.05]	[0.07]	[0.10]		[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Austria	0.39	0.75	1.04	1.44	2.21	Italy	0.65	1.27	1.76*	2.21*	5.50***
	(1.14)	(1.24)	(1.26)	(1.43)	(1.39)		(1.09)	(1.46)	(1.78)	(1.95)	(3.06)
	[0.01]	[0.01]	[0.02]	[0.02]	[0.02]		[0.01]	[0.02]	[0.02]	[0.03]	[0.09]
Belgium	2.25***	4.18***	5.74***	6.67***	6.96***	Japan	0.80**	1.33**	1.58**	1.77*	-0.31
	(4.50)	(5.97)	(7.15)	(7.28)	(4.27)		(2.24)	(2.21)	(1.96)	(1.89)	(-0.18)
	[0.18]	[0.24]	[0.28]	[0.26]	[0.12]		[0.05]	[0.05]	[0.04]	[0.04]	[0.00]
Canada	1.21***	2.34***	3.08***	3.40***	3.44***	Netherlands	0.69*	1.29**	2.13***	2.90***	4.09***
	(4.71)	(5.48)	(4.95)	(4.22)	(3.14)		(1.90)	(2.48)	(3.17)	(3.69)	(3.05)
	[0.07]	[0.19]	[0.19]	[0.17]	[0.10]		[0.02]	[0.03]	[0.05]	[0.06]	[0.05]
Denmark	0.36***	0.76***	1.21***	1.62***	3.18***	Spain	-0.40	0.04	1.76	3.84	4.91
	(2.61)	(3.27)	(3.74)	(4.07)	(4.80)		(-0.30)	(0.02)	(0.81)	(1.45)	(1.46)
	[0.06]	[0.10]	[0.12]	[0.13]	[0.21]		[0.00]	[0.00]	[0.01]	[0.03]	[0.02]
Finland	0.92*	1.68**	2.27**	2.74**	4.01*	Sweden	0.87	2.20***	3.45***	3.98***	4.65***
	(1.94)	(2.06)	(2.18)	(2.11)	(1.86)		(1.58)	(2.77)	(3.60)	(3.68)	(3.73)
	[0.03]	[0.03]	[0.04]	[0.04]	[0.03]		[0.03]	[0.08]	[0.11]	[0.11]	[0.08]
France	1.83***	3.99***	5.91***	7.70***	13.51***	UK	0.95***	2.08***	3.13***	4.13***	5.68***
	(2.87)	(4.13)	(5.20)	(6.04)	(8.12)		(2.49)	(4.25)	(5.30)	(5.65)	(4.66)
	[0.07]	[0.14]	[0.19]	[0.22]	[0.32]		[0.06]	[0.12]	[0.17]	[0.22]	[0.22]
Germany	-1.52***	-2.16***	-2.15**	-2.06	1.19	US	0.50	1.14	1.93*	2.41*	5.42***
	(-3.10)	(-2.65)	(-1.99)	(-1.51)	(0.68)		(0.88)	(1.38)	(1.79)	(1.93)	(3.68)
	[0.06]	[0.05]	[0.03]	[0.02]	[0.00]		[0.01]	[0.02]	[0.03]	[0.04]	[0.09]

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets. *, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

In accordance with the findings for cay_t, it shows that $cday_t$ is statistically significant for almost all countries, the point estimate of the coefficient is large in magnitude and its sign is positive. Therefore, deviations in the long-term trend among c_t, f_t, h_t and y_t should be positively linked with future stock returns.

In addition, it can be seen that the trend deviations explain a substantial fraction of the variation in future real returns. At the four quarter horizon, $cday_t$ explains 26% (Belgium), 22% (France and UK), 17% (Canada), 13% (Denmark), 7% (Australia), 6% (Netherlands), 4% (Finland and US) of the real stock return. However, it does not seem to exhibit forecasting power for countries such as Germany, Ireland, and Spain.

Noticeably, it is important to emphasize that, in general, $cday_t$ performs better than cay_t, also in accordance with the findings of Sousa (2010a), reflecting the ability of $cday_t$ to track the changes in the composition of asset wealth. Portfolios with different compositions of assets are subject to different degrees of liquidity, taxation, or transaction

costs. For example, agents who hold portfolios where the exposure to housing wealth is larger face an additional risk associated with the (il)liquidity of these assets and the transaction costs involved in trading them. Wealth composition is, therefore, an important source of risk that $cday_t$ – but not cay_t – is able to capture.

3.2. Forecasting government bond returns

We now look at the power of cay_t (Table 3.1) and $cday_t$ (Table 3.2) in predicting bond returns (proxied by the government bond yields and denoted by BR_t) for which quarterly data are available. As mentioned before, one needs to keep in mind that, in contrast with stocks, an increase in the government bond return may not be seen as a rise in wealth, but may be perceived as a signal of a future increase in taxes. Therefore: (i) when agents see government debt as a component of wealth, one should observe a positive point coefficient for cay_t and/or $cday_t$ in the forecasting regressions; and (ii) when agents interpret the rise in government debt as a signal of deterioration of public finances, then cay_t and $cday_t$ should be negatively related to future government bond returns.

Table 3.1 shows that cay_t is statistically significant for almost all countries and the point estimate of the coefficient is large in magnitude. It can also be seen that the trend deviations strongly predict future real government bond yields, in particular, at horizons spanning from three to four quarters. Indeed, at the four quarter horizon, cay_t explains 64% (Italy), 31% (Sweden), 33% (Australia), 27% (Canada), 23% (Germany), 13% (Belgium), 11% (Denmark), 10% (Ireland) and 8% (Japan) of the real government bond returns. As for France, Spain and the UK, the forecasting power of cay_t is virtually nil.

Table 3.1 – Forecasting real bond returns: estimated effect of cay.

$$BR_{t+1}+ BR_{t+2}+...+ BR_{t+H} = f(cay_{t-1}),\ H=1, 2, 3, 4, 8.$$

	Forecast Horizon H						Forecast Horizon H				
	1	2	3	4	8		1	2	3	4	8
Australia	0.54***	1.08***	1.62***	2.15***	4.26***	Ireland	-0.48***	-1.02***	-1.61***	-2.17***	-3.72***
	(7.81)	(8.61)	(9.23)	(9.29)	(9.20)		(-2.96)	(-3.24)	(-3.43)	(-3.52)	(-3.06)
	[0.30]	[0.32]	[0.33]	[0.33]	[0.33]		[0.07]	[0.08]	[0.10]	[0.10]	[0.08]
Austria	-0.08	-0.17	-0.27*	-0.36*	-0.84**	Italy	-0.44***	-0.91***	-1.38***	-1.86***	-3.86***
	(-1.08)	(-1.39)	(-1.70)	(-1.86)	(-2.25)		(-11.10)	(-12.34)	(-13.02)	(-13.61)	(-15.52)
	[0.01]	[0.02]	[0.03]	[0.04]	[0.05]		[0.57]	[0.60]	[0.62]	[0.64]	[0.70]
Belgium	-0.47***	-1.00***	-1.60***	-2.17***	-4.11***	Japan	-0.09	-0.26	-0.58	-0.98***	-1.72***
	(-3.06)	(-3.58)	(-4.01)	(-4.26)	(-4.03)		(-0.27)	(-0.60)	(-1.19)	(-3.52)	(-3.70)
	[0.08]	[0.11]	[0.12]	[0.13]	[0.12]		[0.00]	[0.01]	[0.03]	[0.08]	[0.07]
Canada	-0.42***	-0.86***	-1.33***	-1.80***	-3.48***	Netherlands	0.33***	0.57***	0.80***	0.99***	1.31***
	(-5.15)	(-5.54)	(-5.97)	(-6.32)	(-6.76)		(4.79)	(4.44)	(4.47)	(4.10)	(2.71)
	[0.21]	[0.23]	[0.26]	[0.27]	[0.28]		[0.10]	[0.09]	[0.08]	[0.07]	[0.03]
Denmark	-0.40***	-0.81***	-1.25***	-1.79***	-4.11***	Spain	0.10	0.17	0.23	0.24	0.58
	(-3.72)	(-4.06)	(-4.05)	(-4.23)	(-4.61)		(0.46)	(0.40)	(0.36)	(0.28)	(0.31)
	[0.08]	[0.09]	[0.10]	[0.11]	[0.15]		[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Finland	0.39**	0.75**	1.09**	1.36**	2.13*	Sweden	-0.30***	-0.56***	-0.94***	-1.31***	-2.70***
	(2.36)	(2.44)	(2.43)	(2.32)	(1.86)		(-2.66)	(-3.79)	(-5.25)	(-7.31)	(-8.28)
	[0.06]	[0.07]	[0.07]	[0.06]	[0.04]		[0.06]	[0.13]	[0.21]	[0.31]	[0.34]
France	0.22*	0.40*	0.54	0.65	0.79	UK	-0.07	-0.18	-0.30	-0.41	-1.01**
	(1.91)	(1.73)	(1.55)	(1.39)	(0.85)		(-0.97)	(-1.34)	(-1.54)	(-1.58)	(-2.01)
	[0.03]	[0.02]	[0.02]	[0.01]	[0.01]		[0.01]	[0.01]	[0.02]	[0.02]	[0.03]
Germany	-0.17***	-0.33***	-0.51***	-0.73***	-1.59***	US	-0.27**	-0.54**	-0.82**	-1.14**	-2.16**
	(-4.39)	(-4.52)	(-4.98)	(-5.93)	(-7.51)		(-2.13)	(-2.15)	(-2.23)	(-2.34)	(-2.24)
	[0.11]	[0.15]	[0.18]	[0.23]	[0.32]		[0.04]	[0.04]	[0.04]	[0.04]	[0.04]

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets. *, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

Interestingly the results suggest that the sign of the coefficient of cay_t is positive for Australia, Finland, and the Netherlands and negative for Belgium, Canada, Denmark, Germany, Ireland, Italy and the US. This piece of evidence corroborates the idea that agents allow consumption to rise above its equilibrium relationship with asset wealth and labour income when they expect government bond yields to increase in the future. As for the second set of countries, agents perceive the rise in government bond returns as a deterioration of the public finances and an increase in future taxation. In practice, these results largely reflect higher sustainability of public finances in the first set of countries.[9] As for the second set of countries, they characterize well the relatively frequent swings in public deficits and government debt and the concerns about the long-term sustainability of public finances.

Table 3.2 describes the results from forecasting regressions of $cday_t$ at different horizons.

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[9] Interestingly, Afonso (2005) finds that it is not possible to reject the hypothesis of cointegration between government spending and revenue for Austria, Finland, and the Netherlands.

Table 3.2 – Forecasting real bond returns: estimated effect of *cday*.
$$BR_{t+1}+ BR_{t+2}+...+ BR_{t+H} = f(cday_{t-1}), H=1, 2, 3, 4, 8.$$

	Forecast Horizon H						Forecast Horizon H				
	1	2	3	4	8		1	2	3	4	8
Australia	0.67***	1.37***	2.07***	2.76***	5.56***	Ireland	-0.14	-0.35	-0.60	-0.78	-0.67
	(7.60)	(8.46)	(9.20)	(9.28)	(9.48)		(-0.60)	(-0.76)	(-0.88)	(-0.89)	(-0.45)
	[0.32]	[0.35]	[0.36]	[0.37]	[0.39]		[0.01]	[0.01]	[0.01]	[0.01]	[0.00]
Austria	-0.10	-0.21*	-0.32**	-0.43**	-0.95***	Italy	-0.01	-0.06	-0.16	-0.29	-1.30
	(-1.37)	(-1.75)	(-2.06)	(-2.21)	(-2.64)		(-0.05)	(-0.18)	(-0.30)	(-0.43)	(-1.01)
	[0.02]	[0.03]	[0.05]	[0.05]	[0.06]		[0.00]	[0.00]	[0.00]	[0.00]	[0.01]
Belgium	0.45***	0.90***	1.25**	1.62**	3.35***	Japan	-0.00	-0.03	-0.25	-0.63**	-1.23***
	(2.72)	(2.74)	(2.44)	(2.39)	(2.55)		(-0.01)	(-0.07)	(-0.50)	(-2.46)	(-3.13)
	[0.06]	[0.07]	[0.06]	[0.06]	[0.07]		[0.00]	[0.00]	[0.01]	[0.04]	[0.04]
Canada	-0.52***	-1.08***	-1.66***	-2.32***	-3.91***	Netherlands	0.48***	0.88***	1.27***	1.63***	2.68***
	(-5.45)	(-5.94)	(-6.42)	(-6.79)	(-5.87)		(6.84)	(6.27)	(6.17)	(6.09)	(5.14)
	[0.21]	[0.24]	[0.26]	[0.27]	[0.22]		[0.21]	[0.21]	[0.20]	[0.19]	[0.13]
Denmark	-0.36***	-0.71***	-1.08***	-1.52***	-3.37***	Spain	0.11	0.04	-0.29	-0.81	-2.41
	(-3.43)	(-3.65)	(-3.58)	(-3.71)	(-3.99)		(0.30)	(0.05)	(-0.27)	(-0.58)	(-0.86)
	[0.07]	[0.07]	[0.08]	[0.09]	[0.11]		[0.00]	[0.00]	[0.00]	[0.00]	[0.01]
Finland	0.80***	1.58***	2.32***	2.92***	4.56***	Sweden	0.07	0.30	0.29	0.34	1.03
	(4.63)	(5.40)	(5.67)	(5.53)	(4.27)		(0.30)	(1.20)	(0.88)	(0.98)	(1.55)
	[0.12]	[0.14]	[0.14]	[0.18]	[0.09]		[0.00]	[0.01]	[0.01]	[0.01]	[0.02]
France	0.54***	1.03***	1.46***	1.84***	3.02***	UK	0.02	-0.00	-0.02	-0.00	0.68
	(3.62)	(3.43)	(3.23)	(3.07)	(2.54)		(0.17)	(-0.02)	(-0.07)	(-0.01)	(0.88)
	[0.10]	[0.09]	[0.08]	[0.08]	[0.05]		[0.00]	[0.00]	[0.00]	[0.00]	[0.01]
Germany	-0.04	-0.00	-0.01	-0.08	-.17	US	0.01	0.01	-0.00	-0.07	-1.62
	(-0.37)	(-0.00)	(-0.06)	(-0.23)	(-0.25)		(0.05)	(0.04)	(-0.01)	(-0.09)	(-1.03)
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]		[0.00]	[0.00]	[0.00]	[0.00]	[0.01]

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

In accordance with the findings for cay_t, Table 3.2 shows that $cday_t$ is statistically significant for a reasonable set of countries. At the four quarter horizon, $cday_t$ explains 37% (Australia), 27% (Canada), 19% (Netherlands), 18% (Finland), 9% (Denmark), 8% (France), 6% (Belgium), 5% (Austria) and 4% (Japan) of the real government bond returns. As for Germany, Ireland, Italy, Spain, Sweden, the UK and the US, the forecasting power of $cday_t$ is negligible.

The results also suggest that the sign of the coefficient of $cday_t$ is positive for Australia, Belgium, Finland, France and the Netherlands, therefore, supporting the idea that agents behave in a non-Ricardian manner. As for Austria, Canada, Denmark and Japan, the sign of the coefficient of $cday_t$ is negative and, therefore, indicates that agents behave in a Ricardian fashion. In fact, a rise in government bond yields is perceived as a signal of deterioration of public finances.[10]

Under the Ricardian Equivalence hypothesis, forward-looking consumers save the proceeds from a debt-financed fiscal stimulus in anticipation of the future tax increase needed to repay the additional government debt. Such Ricardian behaviour implies that consumers' net wealth is invariant to a debt-financed government expenditure increase.

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[10] Afonso (2008) also reports empirical evidence regarding the existence of Ricardian fiscal regimes in the European Union. By its turn, Sousa (2010c) shows that investors in the euro area as a whole seem to consider government debt as a component of asset wealth.

However, some rather restrictive assumptions need to be in place, notably: infinitely living households; lump-sum taxes; efficient capital markets; and absence of credit constraints. [11]

Nevertheless, if fiscal expansions are perceived as permanent, and lead to expectations of much higher government debt, the importance of Ricardian behaviour may actually increase. In addition, the possible negative reactions of financial markets to sizeable increases in government debt may undermine the expected positive economic effect from a fiscal expansion. Indeed, an increased risk of government debt default and the potential rise in interest rates will dampen or even offset the related economic stimulus.

Finally, we assess the existence of a potential bias in the coefficients associated with *cay* and *cday* in the forecasting regressions of stock returns and government bond yields. Stambaugh (1986, 1999) suggest that when the empirical proxies *cay* and *cday* – used as regressors in the forecasting equations – are autocorrelated and the shocks to regressors are correlated with shocks to returns, the dependent variable is not independent of all leads and lags of the error terms. This, in turn, generates an upward bias in the estimations, which is, approximately, equal to $\gamma(1+3\rho)/T$ under the normality assumption, where γ is the coefficient obtained from regressing the residual in the returns regression on the residual from an AR(1) regression for the forecasting variables (*cay* or *cday*), ρ is the AR coefficient for the forecasting variables, and T is the sample size.

The magnitudes of the bias are shown in Table 3.3 and suggest that it does not seem to affect the predictive power of both *cay* and *cday*. In fact, the bias is small in the forecasting regressions at different horizons. Consequently, even after making this adjustment, the empirical proxies *cay* and *cday* are statistically significant and important predictors of stock returns and government bond yields. This is also in accordance with the works Lettau and Ludvigson (2001) and Sousa (2010a). Similarly, Whelan (2008) finds that the bias does not impact on the predictive ability of the excess consumption to observable assets.

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[11] See notably Barro (1976) and Seater (1993).

Table 3.3 – Forecasting asset returns: checking for potential bias

Stock returns, cay	Forecast Horizon H					Stock returns, cday	Forecast Horizon H				
	1	2	3	4	8		1	2	3	4	8
Australia	0.03	0.10	0.13	0.13	0.24	Australia	0.03	0.07	0.08	0.06	0.16
Austria	0.05	0.06	0.07	0.07	0.15	Austria	0.05	0.06	0.07	0.07	0.13
Belgium	-0.03	0.05	0.13	0.19	0.23	Belgium	-0.03	0.03	0.09	0.15	0.27
Canada	0.00	0.04	0.09	0.13	0.18	Canada	-0.00	0.03	0.08	0.12	0.18
Denmark	0.01	0.01	0.02	0.04	0.05	Denmark	0.01	0.01	0.03	0.05	0.05
Finland	-0.10	-0.16	-0.17	-0.14	0.08	Finland	-0.02	0.03	0.09	0.13	0.15
France	0.00	-0.01	0.04	0.09	0.27	France	-0.01	-0.01	0.04	0.08	0.27
Germany	0.03	0.03	0.04	0.05	0.02	Germany	0.00	-0.03	-0.05	-0.05	-0.04
Ireland	-0.01	0.03	0.04	-0.01	0.04	Ireland	-0.04	-0.00	0.01	-0.05	0.03
Italy	0.11	0.12	0.14	0.21	0.24	Italy	-0.03	-0.02	0.06	0.11	0.02
Japan	0.02	0.05	0.07	0.07	0.05	Japan	0.01	0.04	0.06	0.06	0.04
Netherlands	0.02	0.04	0.04	0.05	0.09	Netherlands	0.02	0.04	0.04	0.05	0.10
Spain	-0.06	-0.09	-0.07	-0.08	0.20	Spain	-0.02	-0.08	-0.09	-0.07	0.14
Sweden	0.04	0.04	0.08	0.17	0.24	Sweden	-0.06	-0.09	-0.05	0.07	0.18
UK	-0.03	-0.02	0.02	0.05	0.16	UK	-0.04	-0.03	-0.01	0.01	0.07
US	-0.03	-0.03	-0.01	0.01	0.06	US	0.01	0.02	0.02	0.04	0.10

Gov. bond returns, cay	Forecast Horizon H					Gov. bond returns, cday	Forecast Horizon H				
	1	2	3	4	8		1	2	3	4	8
Australia	0.80	0.04	0.06	0.08	0.14	Australia	0.02	0.03	0.05	0.06	0.12
Austria	0.04	0.00	-0.00	-0.00	-0.01	Austria	0.00	-0.00	-0.00	-0.01	-0.02
Belgium	0.57	0.02	0.02	0.02	0.12	Belgium	0.02	0.04	0.06	0.07	0.12
Canada	0.06	-0.00	-0.00	-0.01	-0.05	Canada	-0.00	-0.01	-0.02	-0.03	-0.08
Denmark	0.04	0.00	0.01	0.02	0.01	Denmark	-0.00	-0.00	-0.00	0.01	0.00
Finland	0.91	0.04	0.06	0.08	0.14	Finland	0.05	0.07	0.09	0.14	0.28
France	0.42	0.02	0.03	0.04	0.05	France	0.02	0.03	0.04	0.05	0.08
Germany	-0.10	-0.01	-0.01	-0.01	-0.03	Germany	-0.00	-0.00	-0.00	-0.00	-0.01
Ireland	0.52	0.03	0.03	0.03	0.01	Ireland	0.03	0.05	0.06	0.07	0.09
Italy	-0.72	-0.04	-0.05	-0.07	-0.11	Italy	-0.02	-0.02	-0.02	-0.02	-0.00
Japan	-0.21	-0.00	-0.01	-0.03	-0.06	Japan	-0.01	-0.01	-0.00	-0.02	-0.04
Netherlands	0.20	0.02	0.02	0.03	0.04	Netherlands	0.01	0.02	0.03	0.04	0.07
Spain	-1.33	-0.07	-0.09	-0.11	-0.23	Spain	-0.01	-0.02	-0.03	-0.04	-0.17
Sweden	-0.70	-0.04	-0.01	-0.02	-0.03	Sweden	-0.00	-0.03	0.00	-0.01	-0.03
UK	0.28	0.02	0.02	0.02	0.00	UK	0.00	0.01	0.01	0.00	0.00
US	-0.07	-0.00	-0.01	-0.01	-0.02	US	0.00	0.00	0.01	0.01	-0.01

Notes: the magnitude of the bias is, approximately, equel to $\gamma/(1+3\rho)/T$, under the normality assumption; γ is the coefficient from regressing the residual in the returns regression on the residual from an AR(1) regression for the forecasting variables; ρ is the AR coefficient for the forecasting variables; T is the sample size. (Stambaugh, 1986, 1999).

3.3. Forecasting consumption growth

In principle, *cay* and *cday* could be a proxy for expected future consumption growth, asset returns, or both. Table 4.1 present the results of the regressions of consumption growth over horizons spanning 1 to 8 quarters, on the lag of trend deviation *cday*. Table 4.2 provides a summary of the findings when the lag of *cday* is used as the explanatory variable. In the estimation of the regressions of consumption growth, the dependent variable is, therefore, the *H*-period consumption growth rate $\Delta c_{t+1} + ... + \Delta c_{t+H}$.

Consistent with the findings of Lettau and Ludvigson (2001), the results shown in Table 4.1 suggest that, in general, cay_t has no predictive power for future consumption growth. The individual coefficients are not statistically significant and the adjusted R-square is close to zero. A few exceptions include the cases of Canada, France, Netherlands, Spain, Sweden and the UK, where the trend deviation cay_t exhibits some predictive power. Nevertheless, one should note that the coefficients are still very

small in magnitude. As a result, the residuals from the contegrating relationship among consumption, asset wealth and labour income can be generally described as a good predictor of asset returns (and not of future consumption growth).

Table 4.1 – Forecasting consumption growth: estimated effect of *cay*.
$$\Delta c_{t+1} + \Delta c_{t+2} + \dots + \Delta c_{t+H} = f(cay_{t-1}), H = 1, 2, 3, 4, 8.$$

	Forecast Horizon *H*						Forecast Horizon *H*				
	1	2	3	4	8		1	2	3	4	8
Australia	-0.05*	-0.07*	-0.09*	-0.08	-0.02	Ireland	-0.03	-0.08	-0.12	-0.12	0.00
	(1.71)	(-1.71)	(-1.69)	(-1.57)	(-0.31)		(-0.40)	(-0.67)	(-0.67)	(-0.58)	(0.02)
	[0.03]	[0.03]	[0.03]	[0.02]	[0.00]		[0.00]	[0.00]	[0.01]	[0.00]	[0.00]
Austria	-0.05	-0.06	-0.08	-0.05	-0.03	Italy	-0.00	-0.00	0.00	0.01	0.04
	(-1.00)	(-1.08)	(-1.39)	(-0.81)	(-0.42)		(-0.28)	(-0.15)	(0.12)	(0.32)	(0.65)
	[0.01]	[0.01]	[0.02]	[0.01]	[0.00]		[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Belgium	-0.01	-0.07	-0.15**	-0.18*	-0.04	Japan	0.04	0.09	0.10	0.15	0.15
	(-0.26)	(-0.94)	(-2.16)	(-1.94)	(-0.46)		(1.10)	(1.57)	(1.22)	(1.36)	(0.86)
	[0.00]	[0.02]	[0.05]	[0.05]	[0.00]		[0.01]	[0.02]	[0.03]	[0.03]	[0.01]
Canada	0.13***	0.27***	0.39***	0.51***	0.85***	Netherlands	-0.13***	-0.25***	-0.34***	-0.41***	-0.67***
	(5.03)	(7.45)	(8.56)	(8.69)	(9.35)		(-2.72)	(-3.60)	(-4.09)	(-4.05)	(-4.04)
	[0.18]	[0.33]	[0.41]	[0.46]	[0.44]		[0.07]	[0.12]	[0.14]	[0.14]	[0.12]
Denmark	0.03	0.07	0.07	0.10	0.03	Spain	0.12***	0.26***	0.41***	0.60***	1.20***
	(0.65)	(1.08)	(0.91)	(1.22)	(0.21)		(2.64)	(3.37)	(4.04)	(5.06)	(7.71)
	[0.00]	[0.01]	[0.01]	[0.01]	[0.00]		[0.09]	[0.16]	[0.22]	[0.30]	[0.49]
Finland	-0.01	0.00	0.04	0.09	0.46***	Sweden	0.04***	0.07***	0.11***	0.14***	0.31***
	(-0.40)	(0.01)	(0.59)	(1.18)	(3.44)		(2.59)	(2.97)	(3.17)	(3.36)	(4.84)
	[0.00]	[0.00]	[0.00]	[0.01]	[0.10]		[0.05]	[0.06]	[0.07]	[0.07]	[0.13]
France	-0.05	-0.14***	-0.20***	-0.24***	-0.34***	UK	0.10***	0.20***	0.28***	0.39***	0.73***
	(-1.41)	(-2.53)	(-2.93)	(-2.92)	(-2.72)		(3.72)	(4.91)	(5.73)	(6.80)	(8.04)
	[0.02]	[0.06]	[0.07]	[0.06]	[0.06]		[0.07]	[0.12]	[0.15]	[0.20]	[0.29]
Germany	0.01	0.06	0.10**	0.13**	0.31***	US	0.02	0.03	0.02	0.02	-0.02
	(0.44)	(1.44)	(1.95)	(2.23)	(3.33)		(0.44)	(0.58)	(0.08)	(0.16)	(-0.11)
	[0.00]	[0.02]	[0.04]	[0.05]	[0.10]		[0.00]	[0.00]	[0.00]	[0.00]	[0.00]

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

In contrast, Table 4.2 shows that *cday*$_t$ contains some relevant information about future consumption growth. The coefficients associated with the deviations of consumption from its trend relationship with financial wealth, housing wealth, and labour income are statistically significant for Canada, France, Germany, Ireland, Italy, Netherlands and Sweden. In the case of Canada, this is also reflected in a relatively large adjusted R-square. These findings may be related with the fact that *cday*$_t$ tracks changes in the composition of asset wealth and, in particular, in housing wealth. Given that housing wealth changes tend to have small but also very persistent effects on consumption,[12] *cday*$_t$ may, therefore, be capturing time-variation in expected returns *and* in consumption growth. Finally, one should emphasize that despite this, the coefficients associated with *cday*$_t$ in the regressions of consumption growth are quite

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[12] See Sousa (2010d) for a review of the topic and, in particular, an application of the wealth effects on consumption to the euro area as a whole. Similarly, Sousa (2009) finds evidence of important wealth effects from monetary policy.

small, which is in contrast with the findings in the regressions of real stock returns and real government bond yields. Consequently, $cday_t$ mainly forecasts asset returns.

Table 4.2 – Forecasting consumption growth: estimated effect of *cday*.
$$\Delta c_{t+1} + \Delta c_{t+2} + \ldots + \Delta c_{t+H} = f(cday_{t-1}),\ H=1,\ 2,\ 3,\ 4,\ 8.$$

	Forecast Horizon H						Forecast Horizon H				
	1	2	3	4	8		1	2	3	4	8
Australia	-0.07**	-0.10**	-0.12**	-0.12*	-0.09	Ireland	-0.18**	-0.40***	-0.61***	-0.78***	-1.24***
	(-2.09)	(-2.09)	(-1.96)	(-1.92)	(-1.33)		(-2.12)	(-2.58)	(-2.81)	(-3.02)	(-3.78)
	[0.04]	[0.05]	[0.05]	[0.04]	[0.01]		[0.09]	[0.12]	[0.13]	[0.13]	[0.12]
Austria	-0.04	-0.05	-0.07	-0.04	-0.02	Italy	-0.07***	-0.14***	-0.17***	-0.17***	-0.04
	(-0.98)	(-1.04)	(-1.29)	(-0.73)	(-0.21)		(-3.49)	(-3.92)	(-3.31)	(-2.55)	(-0.31)
	[0.01]	[0.01]	[0.01]	[0.00]	[0.00]		[0.07]	[0.08]	[0.06]	[0.04]	[0.00]
Belgium	-0.04	-0.10*	-0.19***	-0.25***	-0.21	Japan	0.05	0.12**	0.15*	0.22**	0.28
	(-0.92)	(-1.82)	(-2.71)	(-2.74)	(1.62)		(1.39)	(1.94)	(1.81)	(2.06)	(1.57)
	[0.01]	[0.04]	[0.07]	[0.08]	[0.03]		[0.01]	[0.04]	[0.05]	[0.06]	[0.05]
Canada	0.15***	0.30***	0.44***	0.56***	0.83***	Netherlands	-0.13***	-0.26***	-0.36***	-0.44***	-0.76***
	(4.75)	(7.04)	(8.07)	(7.31)	(6.42)		(-2.90)	(-3.78)	(-4.24)	(-4.21)	(-4.48)
	[0.14]	[0.26]	[0.34]	[0.36]	[0.26]		[0.07]	[0.14]	[0.16]	[0.16]	[0.16]
Denmark	0.02	0.04	0.02	0.03	-0.18	Spain	-0.23***	-0.34***	-0.40**	-0.33	0.10
	(0.43)	(0.67)	(0.24)	(0.35)	(-1.33)		(-2.68)	(-2.83)	(-2.30)	(-1.35)	(0.24)
	[0.00]	[0.00]	[0.00]	[0.00]	[0.01]		[0.12]	[0.10]	[0.07]	[0.03]	[0.00]
Finland	-0.05	-0.08	-0.08	-0.10	0.04	Sweden	-0.11***	-0.21***	-0.31***	-0.38***	-0.47***
	(-1.36)	(-1.19)	(-0.94)	(-0.87)	(0.26)		(-2.74)	(-3.38)	(-3.66)	(-3.79)	(-2.80)
	[0.01]	[0.02]	[0.01]	[0.01]	[0.00]		[0.10]	[0.14]	[0.15]	[0.15]	[0.08]
France	-0.07	-0.18***	-0.27***	-0.33***	-0.61***	UK	0.05	0.07	0.07	0.14	0.28**
	(-1.50)	(-2.76)	(-3.29)	(-3.36)	(-3.99)		(1.42)	(1.27)	(0.93)	(1.56)	(2.17)
	[0.02]	[0.07]	[0.08]	[0.08]	[0.12]		[0.01]	[0.01]	[0.01]	[0.02]	[0.03]
Germany	-0.18**	-0.26***	-0.34***	-0.50***	-0.68***	US	-0.01	-0.02	-0.07	-0.10	-0.32
	(-1.96)	(-2.79)	(-3.19)	(-5.04)	(-3.56)		(-0.15)	(-0.29)	(-0.63)	(-0.67)	(-1.41)
	[0.04]	[0.06]	[0.06]	[0.11]	[0.08]		[0.00]	[0.00]	[0.00]	[0.01]	[0.02]

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

4. Robustness analysis

4.1. Additional control variables

Campbell and Shiller (1988), Fama and French (1988) and Lamont (1998) show that the ratios of price to dividends or earnings or the ratio of dividends to earnings have predictive power for stock returns. More recently, Goyal and Welsh (2003) argue that because the dividend yield follows a random walk it cannot predict stock prices. However, Robertson and Wright (2006) and Boudoukh et al. (2007) mention that a change in tax legislation in the US in 1983 that legalised share buybacks implies an adjustment of the dividend yield for these and similar effects. Consequently, this adjusted statistic is mean reverting and a good predictor of stock returns.

Table 5.1 reports the estimates from forecasting regressions that include additional variables shown to contain predictive power for real stock returns, in particular, the dividend yield ratio ($DivYld_t$). We also add the lag of the real stock returns (SR_{t-1}) as a control variable.

The results show that both the point coefficient estimates of *cay* and *cday* and their statistical significance do not change with respect to the findings of Tables 2.1 and

2.2 where only *cay* and *cday* were included as explanatory variables. Moreover, the lag of the dependent variable is not statistically significant, a feature that is in accordance with the forward-looking behaviour of stock returns. Finally, the dividend yield ratio ($DivYld_t$) seems to provide relevant information about future asset returns since it is statistically significant in practically all regressions and it improves the adjusted R-square.

Table 5.1 – Forecasting real stock returns: additional control variables.
$SR_{t+1} = f(cay_{t-1},...); SR_{t+1} = f(cday_{t-1},...).$

	SR_{t-1}	cay_{t-1}	$DivYld_{t-1}$	Adj. R-square	SR_{t-1}	$cday_{t-1}$	$DivYld_{t-1}$	Adj. R-square
Australia	0.08	0.73**	7.36**	[0.11]	-0.05	0.75**	7.48***	[0.08]
	(0.79)	(2.40)	(2.22)		(-0.44)	(2.02)	(2.86)	
Austria	0.12	0.36		[0.03]	0.13	0.22		[0.02]
	(1.18)	(0.95)			(1.30)	(0.60)		
Belgium	-0.03	2.01***	3.85**	[0.20]	0.01	2.59***	-2.80	[0.23]
	(-0.26)	(6.08)	(1.89)		(0.05)	(4.75)	(-1.32)	
Canada	0.06	1.03***	7.89***	[0.13]	0.09	1.34***	5.54**	[0.17]
	(0.68)	(4.42)	(2.91)		(1.00)	(5.36)	(2.10)	
Denmark	0.43***	0.23**		[0.23]	0.43***	0.20*		[0.22]
	(2.86)	(2.00)			(2.90)	(1.71)		
Finland	0.13	1.00**	4.58	[0.07]	0.12	1.29**	4.29	[0.05]
	(1.23)	(2.42)	(1.41)		(1.15)	(0.60)	(1.20)	
France	-0.06	1.90***	3.64**	[0.10]	-0.03	1.97***	2.34	[0.08]
	(-0.69)	(3.60)	(2.04)		(-0.31)	(3.22)	(1.29)	
Germany	0.02	0.58	5.22	[0.05]	-0.00	-1.36**	6.97*	[0.07]
	(0.19)	(1.32)	(1.33)		(-0.01)	(-2.49)	(1.89)	
Ireland	-0.07	0.24		[0.01]	-0.07	0.25		[0.01]
	(-0.72)	(0.39)			(-0.70)	(0.39)		
Italy	-0.03	0.81***	29.41***	[0.16]	0.06	1.16*	23.33***	[0.10]
	(-0.32)	(3.27)	(4.09)		(0.56)	(1.89)	(2.98)	
Japan	-0.06	0.72**	6.99	[0.07]	-0.04	0.58	6.02	[0.06]
	(-0.62)	(1.99)	(1.61)		(-0.39)	(1.54)	(1.30)	
Netherlands	0.19	-3.63	3.67	[0.10]	0.32*	4.86	-12.46	[0.12]
	(1.05)	(-1.33)	(0.21)		(1.80)	(0.97)	(-0.57)	
Spain	0.12	0.20		[0.02]	0.13	0.09		[0.02]
	(0.88)	(0.20)			(0.77)	(0.05)		
Sweden	-0.16	1.07***	29.68***	[0.19]	-0.03	0.57	14.88***	[0.13]
	(-1.61)	(2.66)	(3.96)		(-0.04)	(1.07)	(3.26)	
UK	-0.21*	0.60***	4.73**	[0.13]	-0.18	0.71***	4.31**	[0.12]
	(-1.80)	(3.20)	(2.44)		(-1.51)	(2.65)	(2.11)	
US	0.05	0.76**	2.32	[0.04]	0.05	0.47	0.91	[0.01]
	(0.57)	(2.42)	(1.32)		(0.66)	(0.81)	(0.53)	

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

On the other hand, Table 5.2 reports the estimates from forecasting regressions that include additional variables shown to contain predictive power for long-term interest rates, in particular, the inflation rate (*Inflation*) and the deficit-to-GDP ratio (*Deficit*).

Gale and Orszag (2003) argue that there are two important reasons why government budget deficits may raise nominal interest rates: (i) budget deficits reduce aggregate savings when private savings do not increase by the same amount and there

are no compensating foreign capital inflows; and (ii) budget deficits increase the stock of government debt and, consequently, the outstanding amount of government bonds. In this case, there is a "portfolio effect", as a higher interest rate on government bonds would be required in order for investors to hold the additional bonds.

Table 5.2 – Forecasting real bond returns: additional control variables.
$$BR_{t+1} = f(cay_{t-1},..\ );\ BR_{t+1} = f(cday_{t-1},..\ ).$$

	cay_{t-1}	$Inflation_{t-1}$	$Deficit_{t-1}$	Adj. R-square	$cday_{t-1}$	$Inflation_{t-1}$	$Deficit_{t-1}$	Adj. R-square
Australia	0.60***	0.00***	0.01	[0.35]	0.71***	0.00**	-0.01	[0.35]
	(9.45)	(3.20)	(0.20)		(8.66)	(2.23)	(-0.13)	
Austria	-0.07	0.00***		[0.11]	-0.09	0.00***		[0.12]
	(-1.03)	(2.90)			(-1.32)	(2.91)		
Belgium	-0.41***	-0.00	0.13***	[0.52]	0.15	-0.00	0.13***	[0.46]
	(-3.83)	(-0.25)	(7.35)		(1.08)	(-0.42)	(6.61)	
Canada	-0.17**	0.00*	-0.86	[0.10]	-0.48***	0.00	-1.26**	[0.33]
	(-2.04)	(1.91)	(-1.27)		(-5.92)	(1.09)	(-2.13)	
Denmark	-0.27**	0.01***		[0.17]	-0.26**	0.01***		[0.17]
	(-2.27)	(3.81)			(-2.40)	(4.12)		
Finland	0.35**	-0.00	-0.18	[0.08]	0.59**	-0.00	-0.11	[0.14]
	(2.01)	(-1.50)	(-1.22)		[2.36]	(-1.29)	(-1.06)	
France	0.61***	0.01***	0.01	[0.38]	0.87***	0.01***	0.00	[0.46]
	(5.95)	(6.71)	(0.54)		(7.59)	(7.09)	(0.03)	
Germany	-0.16***	0.00***	0.07	[0.16]	-0.09	0.00**	-0.12	[0.06]
	(-4.44)	(3.79)	(0.63)		(-0.96)	(2.29)	(-1.31)	
Ireland	-0.48***			[0.07]	-0.14			[0.01]
	(-2.96)				(-0.60)			
Italy	-0.32***	0.01***	0.06	[0.75]	-0.06	0.02***	-0.23***	[0.65]
	(-8.47)	(6.33)	(1.45)		(-0.59)	(7.37)	(-4.40)	
Japan	0.02	0.01***	-2.78**	[0.36]	0.23	0.01***	-3.47**	[0.32]
	(0.05)	(4.72)	(-2.35)		(0.57)	(4.73)	(-2.17)	
Netherlands	0.18**	0.00	-0.19***	[0.24]	0.36***	0.00	-0.14***	[0.30]
	(2.24)	(1.31)	(-3.76)		(4.01)	(1.28)	(-2.73)	
Spain	0.35**	0.02***	-0.43**	[0.30]	0.25	0.02***	-0.39**	[0.28]
	(2.03)	(2.74)	(-2.32)		(0.71)	(2.61)	(-2.00)	
Sweden	-0.38***	-0.00	-0.18*	[0.10]	0.07	0.00	-0.10	[0.01]
	(-3.15)	(0.07)	(-1.81)		(0.29)	(0.37)	(-0.89)	
UK	-0.18***	0.00***	0.01	[0.18]	-0.18**	0.00***	-0.07	[0.16]
	(-3.63)	(2.51)	(0.23)		(-2.25)	(3.01)	(-1.37)	
US	0.03	0.01***	-0.00	[0.11]	0.16	0.02***	-0.66***	[0.34]
	(0.21)	(3.79)	(-0.03)		(1.17)	(5.77)	(-3.23)	

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets.
*, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

While some studies find that interest rates tend to increase after a rise in the deficit, others do not (Barro and Sala-i-Martin, 1990; Engen and Hubbard, 2004). The empirical findings seem to depend on whether expected or current budget deficits are used as explanatory variables (Upper and Worms, 2003; Laubach, 2009), and also on whether yield differentials in Europe with respect to Germany (Codogno et al., 2003) or interest rate swap spreads are used as the dependent variable (Goodhart and Lemmen, 1999; Afonso and Strauch, 2007).

For Europe, the existing evidence points either to a significant (although small) effect (Bernoth et al., 2003; Codogno et al., 2003; Afonso and Strauch, 2007; Faini,

2006; Afonso and Rault, 2010), or to the absence of impact (Heppke-Falk and Hüfner, 2004). For the US, the effect seems to be substantially larger (Gale and Orszag, 2002). For OECD countries, Barro and Sala-i-Martin (1990) suggest that fiscal variables are unimportant determinants of interest rates, while Ardagna (2009) shows that long-term government bond rates fall in periods of budget consolidation and rise when the fiscal position deteriorates.

Our results corroborate the findings of Tables 3.1 and 3.2. In addition, the lag of the deficit-to-GDP ratio, in general, does not help forecasting bond returns. This is particularly so when the lag of *cay* is considered in the set of explanatory variables: the lag of the deficit-to-GDP ratio is not statistically significant and its coefficient is small in magnitude. This evidence seems to support the work of Barro and Sala-i-Martin (1990) and, consequently, it highlights the importance of Ricardian equivalence. In contrast, the lag of the inflation rate is statistically significant in roughly all regressions, which suggests that investors use government bonds to hedge against the risk of inflation.

4.2. Nested forecast comparisons

As a final robustness exercise, we make nested forecast comparisons, in which we compare the mean-squared forecasting error from a series of one-quarter-ahead out-of-sample forecasts obtained from a prediction equation that includes either *cay* or *cday* as the only forecasting variables, to a variety of forecasting equations that do not include either *cay* or *cday*.

We consider two benchmark models: the *autoregressive benchmark* and the *constant expected returns benchmark*. In the *autoregressive benchmark*, we compare the mean-squared forecasting error from a regression that includes just the lagged asset return as a predictive variable to the mean-squared error from regressions that include, in addition, *cay* or *cday*. In the *constant expected returns benchmark*, we compare the mean-squared forecasting error from a regression that includes a constant to the mean-squared error from regressions that include, in addition, *cay* or *cday*.

A summary of the nested forecast comparisons for the equations of the real stock returns and the government bond yields using respectively *cay* and *cday* is provided in Tables 6.1 and 6.2. In general, including *cay* in the forecasting regressions improves vis-à-vis the benchmark models. This is especially true in the case of the of the *constant expected returns benchmark*, supporting the evidence that reports time-variation in expected returns.

In addition, the models that include *cday* generally have a lower mean-squared forecasting error. Moreover, the ratios are smaller that the ones presented in Table 6.1, reflecting the better predicting ability for stock returns and government bond yields of *cday* relative to *cay*.

Table 6.1 – One-quarter ahead forecasts of returns: a comparison with benchmark models. *cay* model vs. constant/AR models.

	Real stock returns		Real bond returns	
	$MSE_{cay}/MSE_{constant}$	MSE_{cay}/MSE_{AR}	$MSE_{cay}/MSE_{constant}$	MSE_{cay}/MSE_{AR}
Australia	0.975	0.984	0.837	0.995
Austria	1.001	1.006	1.003	1.002
Belgium	0.956	0.942	0.955	0.990
Canada	0.990	0.987	0.892	1.038
Denmark	0.971	0.993	0.968	1.003
Finland	0.991	0.998	0.974	1.003
France	0.980	0.997	0.995	0.990
Germany	1.064	1.069	0.973	1.030
Ireland	1.010	1.008	0.964	0.976
Italy	1.001	1.002	0.662	1.001
Japan	0.878	0.877	0.745	0.815
Netherlands	0.995	0.996	0.952	1.002
Spain	0.976	0.984	0.965	0.539
Sweden	1.005	1.005	0.973	1.006
UK	0.997	1.003	1.018	1.047
US	0.991	0.990	0.984	1.006

Notes: MSE represents the mean-squared forecasting error.
*, **, *** denotes statistical significance at the 10, 5, and 1%percent level, respectively.

Table 6.2 – One-quarter ahead forecasts of returns: a comparison with benchmark models. *cday* model vs. constant/AR models.

	Real stock returns		Real bond returns	
	$MSE_{cday}/MSE_{constant}$	MSE_{cday}/MSE_{AR}	$MSE_{cday}/MSE_{constant}$	MSE_{cday}/MSE_{AR}
Australia	0.981	0.990	0.825	0.997
Austria	1.005	1.008	0.999	1.002
Belgium	0.916	0.924	0.965	0.992
Canada	0.960	0.957	0.889	1.035
Denmark	0.973	0.996	0.974	1.005
Finland	0.997	1.001	0.944	1.004
France	0.981	0.997	0.955	0.989
Germany	1.037	1.041	1.034	1.030
Ireland	1.010	1.010	0.998	0.978
Italy	0.999	1.001	1.006	1.003
Japan	0.879	0.874	0.746	0.797
Netherlands	0.993	0.994	0.890	1.000
Spain	0.977	0.986	0.965	0.553
Sweden	0.991	0.990	1.005	1.006
UK	1.012	1.024	1.021	1.044
US	0.959	0.985	1.039	1.088

Notes: MSE represents the mean-squared forecasting error.
*, **, *** denotes statistical significance at the 10, 5, and 1% level, respectively.

5. Conclusion

This paper uses the representative consumer's budget constraint to derive an equilibrium relation between the trend deviations among consumption, (dis)aggregate wealth and labour income (summarized by the variables *cay* and *cday*) and expected

future asset returns, and explores the predictive power of the empirical counterpart of these trend deviations for both future stock returns and government bond yields.

This strategy follows from the fact that the above-mentioned common trends summarize agent's expectations of stock returns, government bond yields and/or consumption growth. In particular, when stock returns are expected to be higher (lower) in the future, forward-looking investors allow consumption to rise (decrease) above (below) its equilibrium level.

As for government bond yields, if bonds are seen as a component of asset wealth, then investors have a similar reaction. However, if the increase in bond returns is perceived as a signal of public finance deterioration, then investors will allow consumption to fall below its equilibrium relationship with wealth and labour income.

Using data for 16 OECD countries, we show that the predictive ability of *cay* and *cday* for real stock returns is stronger for Australia, Belgium, Canada, Denmark, Finland, the UK and the US. In the case of Germany, Ireland, and Spain, the evidence suggests that those proxies do not capture well the time-variation in stock returns.

Regarding government bond returns, the analysis reveals that we can group the countries in two sets. In the first set (which includes Australia, Finland and the Netherlands), the forecasting regressions show that both *cay* and *cday* have an associated coefficient with positive sign. As a result, government debt is considered as part of the investor's asset wealth. In the second set (which includes Canada, Denmark, Germany, Ireland, Italy and the US), *cay* and *cday* have an associated coefficient that is negative. Consequently, agents in these countries see the rise in government bond returns as a future increase in taxation.

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