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The Consumption—Wealth Ratio and the Japanese Stock Market

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Abstract

Following Lettau and Ludvigson (2001a,b), we examine whether the consumption–wealth ratio can explain Japanese stock market data. We construct the data series cay_t, the residuals from the cointegration relationship between the consumption and the total wealth of households. Unlike the US results, cay_t does not predict future Japanese stock returns. On the other hand, it does help to explain the cross-section of Japanese stock returns of industry portfolios. In the US case, cay_t is used as a scaling variable that explains time variation in the market beta. In the Japanese case, the movement of cay_t is interpreted as the change in the constant terms, hence the change in average stock returns. We also propose to improve cay_t by taking real estate wealth into consideration.

JEL Classification: G12, E21, E23, N12.

Keywords: Consumption–wealth ratio; Cointegration; Cross-section of stock returns.

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1 Introduction

The consumption-based asset pricing model is among the most important benchmarks in financial economics. Yet, its empirical performance with a structural Euler equation of households using aggregate data has been a major disappointment (see Campbell [2003] for a recent survey). Hence, recent studies started looking into other aspects of the consumption-based model. An attractive alternative research strategy is to use disaggregate consumption data, which has been explored by authors such as Mankiw and Zeldes (1991) and Vissing-Jorgensen (2002). More recent studies including Lettau and Ludvigson (2001a,b), Parker and Julliard (2005), and Yogo (2006) examine, using aggregate data, long-run restrictions implied by consumption-based models, and they obtain useful results. In particular, Lettau and Ludvigson (2001a,b) consider the long-run cointegration relationship between consumption and household wealth. They propose to use the "*cay*" variable, which is in essence the consumption–wealth ratio of the household sector, in both predicting aggregate stock returns and explaining cross-sectional patterns of the US stock market.

This paper examines whether Lettau and Ludvigson's framework works with Japanese data. Although there are some other studies examining the same problems in the Japanese market, we take advantage of our familiarity with the data in this paper. We carefully construct a Japanese consumption/financial wealth data set and try to make our definitions of variables as close as possible to the definitions of US data[1]. The most notable feature of the Japanese "*cay*" variable is its high persistence compared with its

[1] There are at least two other unpublished papers examining Lettau and Ludvigson's framework with Japanese data. Gao and Huang (2004) examine mainly the same topics that we are examining in this paper. However, our data construction is much closer to that of Lettau and Ludvigson. Therefore, there are significant differences between our empirical results and those of Gao and Huang. Matsuzaki (2003) uses a Japanese data set that is very similar to ours. However, his definition of consumption is slightly different from ours. This allows him to examine aggregate stock return predictability using a longer data series. He did not find any significant predictability in stock returns for his full sample or the subsample corresponding to our full sample. Therefore, his results on return predictability are consistent with our findings. On the other hand, Matsuzaki does not examine cross-sectional patterns. Neither Gao and Huang (2004) nor Matsuzaki (2003) include real estate wealth in their analysis.

US counterpart. We test whether the "*cay*" variable forecasts future stock returns (Lettau and Ludvigson [2001a]) and whether it helps to explain cross-sectional stock returns (Lettau and Ludvigson [2001b]). We obtain a negative result for the first question but a positive result for the second question. We argue that high persistence of the "*cay*" variable provides explanations for why these Japanese results are different from the US results in some important aspects.

Real estate is often considered an important component of household wealth. However, Lettau and Ludvigson's original works considered only financial and human wealth, perhaps because appropriate nationwide data for real estate wealth is not available for the US. Because owner-occupied housing is a particularly important component of Japanese household wealth, we try to include real estate in the definition of total household wealth. We employ a couple of different real estate variables and use them to calculate consumption–wealth ratios. These augmented "*cay*" variables are even more useful than the original "*cay*" variable in explaining the cross-sectional pattern of Japanese stock returns.

The remainder of this paper is organized as follows. In section 2, we summarize the framework proposed by Lettau and Ludvigson. Section 3 discusses how the data for Japan is constructed. Section 4 presents our main empirical results. Section 5 presents concluding remarks.

2 Analytical framework

Lettau and Ludvigson (2001a,b) use the cointegration among consumption, financial wealth, and human wealth to draw implications for stock returns. This section summarizes their framework. The argument starts from the following general intertemporal budget constraint:

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

where W_t is total wealth and C_t is the consumption of households. Applying the log-linear approximation (Campbell [1991]; Campbell and Shiller [1988]) to (1), we get the following relationship:

$$\Delta w_{t+1} \approx k + r_{w,t+1} + (1 - 1/\rho_w)(c_t - w_t) \tag{2}$$
$$\rho_w \equiv \overline{(W - C)/W},$$

where lowercase letters are natural logs of the variables in equation (1) and $r_{w,t+1} \equiv \ln(1 + R_{w,t+1})$.

The difference equation (2) is solved forward assuming the following "no bubble" condition.

$$lim_{i\to\infty}\rho_w^i(c_{t+i} - w_{t+i}) = 0 \tag{3}$$

After tedious calculations, the following expression for the ex post log consumption–wealth ratio $c_t - w_t$ is obtained.

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

When consistency of investors' expectations is assumed, the following ex ante expression must also hold.

$$c_t - w_t = \mathrm{E}_t\sum_{i=1}^{\infty}\rho_w^i(r_{w,t+1} - \Delta c_{t+i}) \tag{4}$$

To draw empirical implications, Lettau and Ludvigson assume that household total wealth consists of financial wealth a_t and human wealth h_t, i.e., the net present value of future labor income stream.

$$w_t \approx \omega a_t + (1 - \omega)h_t \tag{5}$$

We later extend this definition of household wealth to include real estate.

Therefore, the log return on total household wealth is written as follows.

$$1 + R_{w,t} = \omega_t(1 + R_{a,t}) + (1 - \omega_t)(1 + R_{h,t}) \tag{6}$$
$$r_{w,t} \approx \omega r_{a,t} + (1 - \omega)r_{h,t} \tag{6'}$$

Substituting (6') into ex ante budget constraint (4), we obtain the following.

$$c_t - \omega a_t - (1 - \omega)h_t = \mathrm{E}_t\sum_{i=1}^{\infty}\rho_w^i\left\{[\omega r_{a,t+i} + (1 - \omega)r_{h,t+i}] - \Delta c_{t+i}\right\}$$

Because human wealth h_t cannot be observed, it is assumed to be a linear function of current labor income y_t, so that $h_t = \kappa + y_t + z_t$. Then, h_t can be substituted out from the above expression.

$$\begin{aligned} & c_t - \omega a_t - (1-\omega) y_t \\ = \quad & \mathrm{E}_t \sum_{i=1}^{\infty} \rho_w^i \left\{ [\omega r_{a,t+i} + (1-\omega) r_{h,t+i}] - \Delta c_{t+i} \right\} + (1-\omega) z_t \end{aligned} \tag{7}$$

All right-hand side variables in (7) are stationary, so the sum of the left-hand side should be stationary too. This implies that we have a stationary relationship among $\{c_t, a_t, y_t\}$, which means that they are cointegrated.

Finally, following Lettau and Ludvigson, we define the "*cay*" variable as follows.

$$cay_t \equiv c_t - \omega a_t - (1-\omega) y_t \tag{8}$$

Because ω is not time varying, this is essentially the log consumption–wealth ratio, in which the total wealth of households is defined by the sum of financial and human wealth.

Lettau and Ludvigson estimate cointegration regression (8) to obtain the variable cay_t. In Lettau and Ludvigson (2001a), they use cay_t to forecast future stock returns. Lettau and Ludvigson (2001b) showed that cay_t explains variation in the cross-section of stock returns in the US market.

3 Constructing Japanese data

3.1 Consumption and financial wealth

In this subsection, we summarize the Japanese data used in this paper. For a detailed discussion on the data construction, please refer to Aono and Iwaisako (2006).

Our consumption series is household expenditure on nondurables and services, excluding shoes and clothing. This definition of consumption follows the US benchmark of Wilcox (1992) as well as Lettau and Ludvigson (2001a,b). The data series is taken from the Japanese Cabinet Office's

Annual Report on National Accounts and is seasonally adjusted using X-12. Unfortunately, this definition of Japanese consumption data is available only for after 1970. Then, the data are converted to log real per capita consumption, denoted by c_t.

Financial wealth here is household financial assets measured at the end of each period. It includes the total of all deposits and cash currency, trusts, securities investment trusts, insurance and securities. Data are taken from the Bank of Japan's *Flow of Funds Accounts* data. Using this measure of financial wealth, we construct a log real per capita financial wealth series a_t.

Our labor income is after-tax income reported in the *Annual Report on National Accounts* tabulated by the Japanese Cabinet Office. The log of after-tax labor income, y_t, is also measured in real per capita terms.

3.2 Estimating cointegration relationships

Next, we estimate the cointegration regression by dynamic OLS to obtain the variable cay_t. We follow Lettau and Ludvigson(2001a,b) and use the dynamic least squares technique of Stock and Watson (1993), which specifies a single equation taking the following form:

$$c_t = \alpha + \beta_a a_t + \beta_y y_t + \sum_{i=-k}^{k} b_{a,i} \Delta a_{t-i} + \sum_{i=-k}^{k} b_{y,i} \Delta y_{t-i} + \epsilon_t, \tag{9}$$

where Δ denotes the first difference operator. We denote the estimated trend deviation by $\widehat{cay} = c_t - \hat{\beta}_a a_t - \hat{\beta}_y y_t$, where "hats" denote estimated parameters[2].

As noted in the previous subsection, our consumption series only goes back to 1970. Therefore, our sample period in estimating (9) starts with the first quarter of 1970 and ends with the first quarter of 2004. The following is our full sample estimate.

[2] Following Lettau and Ludvigson(2001), we adopt $k = 8$. However, we obtain a very similar cay_t series even when we use a different number of lags.

$$
\begin{array}{rcllll}
c_t & = & \underset{(17.481)}{3.2565} & + \underset{(9.677)}{0.3125a_t} & + \underset{(3.745)}{0.2242y_t}
\end{array}
\tag{10}
$$

Values reported in parentheses under the parameter estimates are corresponding t-statistics.

We also estimate using the sample after the first oil crisis for a robustness check. The exact period of this subsample is from the first quarter of 1975 to the first quarter of 2004. The following is our subsample estimate.

$$
\begin{array}{rcllll}
c_t & = & \underset{(4.520)}{3.011} & + \underset{(3.383)}{0.2472a_t} & + \underset{(2.108)}{0.3426y_t}
\end{array}
\tag{11}
$$

The estimated subsample parameters are not far from the full sample. In fact, as discussed in Aono and Iwaisako (2006), we obtain very similar values for $\widehat{cay}$ from alternative sample periods. In Aono and Iwaisako (2006), we also considered potential structural breaks in the cointegration relationship and examined various other subsamples. We find that the estimated $\widehat{cay}$ variables behave very similarly to each other, and their predictive abilities for aggregate stock returns are also very similar[3] residuals of the subsample estimation in equation (11). This is mainly because our cross-section data are available only from the second half of the 1970s.

3.3 Including real estate wealth

In modern asset pricing models, investors' market portfolios are the key ingredient in determining asset returns. In estimating the Sharpe–Lintner static CAPM, the market portfolio is typically an aggregate stock market index such as S&P500 or TOPIX. Along with other recent studies, Lettau and Ludvigson's (2001a,b) original framework extended the dimension of investors' market portfolios to include human wealth.

[3]Estimating (9) involves fitting a linear trend to the log consumption–wealth ratio. Therefore, the fitted $\widehat{cay}$ variables' short-run behaviors cannot be very different for the same period, even if the $\widehat{cay}$ variables are calculated using the cointegration regressions for different sample periods.

Another important component of household wealth is real estate. This is particularly true for Japanese households (Iwaisako [2003]; Iwaisako, Mitchell, Piggott [2005]). Hence, we try to include real estate wealth, denoted by rw_t, in our household wealth data. Therefore, equations (5) and (8) are rewritten as follows.

$$w_t \approx \omega_1 a_t + \omega_2 h_t + (1 - \omega_1 - \omega_2) rw_t \tag{5'}$$

$$cay_t \equiv c_t - \omega_1 a_t - \omega_2 h_t - (1 - \omega_1 - \omega_2) rw_t \tag{8'}$$

Unfortunately, there are no quarterly Japanese real estate price data at an aggregate level. We use two alternative methods in calculating cay_t while including real estate wealth. The first method is to use national real estate wealth valuations in GDP statistics. Only annual observations exist for this data series. We fill in missing observations using simple spline interpolations. Admittedly, this is a crude procedure because three out of four observations are interpolated. We add the calculated rw_t and financial wealth to get the series for the log of total nonhuman wealth, $tw_t = a_t + rw_t$. Then, we estimate the cointegration relationship among consumption, nonhuman wealth, and human wealth: $\{c_t, tw_t, y_t\}$. Estimation results are as follows.

$$\begin{array}{ccccccc} c_t & = & 1.9813 & + & 0.1485 tw_t & + & 0.5768 y_t \\ & & (2.459) & & (1.694) & & (2.874) \end{array} \tag{12}$$

As in equation (11), the sample period for equation (12) is the first quarter of 1975 to the first quarter of 2004. Then we calculate $\widehat{cay}$ as we did for equations (10) and (11) in the previous subsection.

In our second approach, we use the urban area land price index (Shigaichi-kakaku-shisu) tabulated by the Japan Real Estate Institute (JREI)[4]. The coverage of JREI's index is narrower than the coverage in the GDP data and is concentrated on urban areas. However, it is reported more frequently on a semiannual basis, as at the end of March and September each year.

[4] The data are available from their website: http://www.reinet.or.jp/jreidata/a_shi/index.htm. They release different types of indexes, and we use the one offering the widest coverage, the index of "*nation wide average*" for "*all purposes.*"

Because it is a price index, it cannot be added to financial wealth to calculate total nonhuman wealth. Therefore, we include rw_t separately as in the cointegration regression, a wealth component independent of both financial and human wealth.

$$\begin{array}{ccccccccc} c_t & = & \underset{(3.272)}{2.7859} & + & \underset{(2.442)}{0.2455a_t} & + & \underset{(1.673)}{0.3702y_t} & - & \underset{(-0.479)}{0.0137rw_t} \end{array} \tag{13}$$

The sample period for (13) is the same as for (11) and (12).

The estimated coefficient of rw_t is negative here. However, this does not necessarily mean that consumption and real estate wealth move in opposite directions, because stock prices and land prices are highly correlated in Japan (Ito and Iwaisako, 1996). In fact, the $\widehat{cay}$ series calculated from (13) most successfully explains the cross-sectional pattern of the Japanese stock market.

4 Empirical results

In this section, we examine whether *cay* helps to forecast future stock returns and whether it helps to explain cross-sectional stock returns with Japanese data.

4.1 Forecasting future stock returns

While Lettau and Ludvigson (2001a) find that $\widehat{cay}$ predicts future stock returns in the US, we find this is not the case for Japan. The results are reported in Table 1. In some specifications, $\widehat{cay}$ seems to predict future stock returns. However, if the lagged returns are included in the regression, its predictive power disappears. Furthermore, the signs of the estimated coefficients of $\widehat{cay}$ are negative in all specifications. This contradicts what the model suggests and the empirical evidence for the US market. Overall, we find very little evidence that $\widehat{cay}$ is useful in predicting future stock returns in the Japanese stock market.

[Table 1 is about here.]

However, we consider this result unsurprising. In the second half of the 1980s, Japan experienced a tremendous stock market boom of historical magnitude (Ito and Iwaisako, 1997). It was followed by a sharp decline in 1990–1992 and prolonged stagnation through the 1990s, known as Japan's lost decade. Because the sample contains such a significant one-time boom and bust in stock prices, any study on Japanese aggregate stock returns including this period faces a major difficulty. We will come back to this issue in subsection 4.3, after we discuss our cross-sectional empirical results.

4.2 Explaining cross-sectional stock returns

Next we examine whether $\widehat{cay}$ helps to explain cross-sectional Japanese stock returns. Here, we use 28 industry portfolio returns tabulated by the *Japan Securities Research Institute* (JSRI). We combine the JSRI data with the Fama–French factors (HML and SMB) available from *Nikkei Media Marketing*, whose data construction closely follows the series of works by Keiichi Kubota and Hitoshi Takehara on Fama–French factors with Japanese data[5]. We convert all asset returns and factor data to a quarterly basis to implement empirical analysis using the $\widehat{cay}$ variable.

Following Lettau and Ludvigson (2001b), we use the Fama–MacBeth two-step approach to examine performances of alternative factors in explaining cross-sectional Japanese stock returns. In the first step, quarterly industry portfolio returns are regressed on alternative sets of factors and conditioning variables.

$$r_{i,t} = \beta_0 + F_t \beta_{1,i} + Z_{t-1} \beta_{2,i} \qquad\qquad i = 1, .., 28$$

Then, in the second step, average returns are regressed on the betas estimated in the first step:

$$\begin{aligned} E[r_{i,t}] &= E[r_{0,t}] + \widehat{\beta}_i \lambda \\ \widehat{\beta} &= \left[\widehat{\beta}_{1,i}, \widehat{\beta}_{2,i}\right], \end{aligned} \tag{14}$$

[5] See, for example, Jagannathan, Kubota and Takehara (1998).

where F_t is the vector of factors including the following variables:

Rvw_t : Market portfolio,
YG_t : Labor income growth,
SMB_t : Fama–French SMB factor,
HML_t : Fama–French HML factor,

and Z_{t-1} includes the following conditioning variables:

$\widehat{cay}_{t-1}$: Consumption–wealth ratio,
$term_{t-1}$: Term premium.

In addition, we also include the scaled market factor, $\widehat{cay}_{t-1} \cdot Rvw_t$, proposed by Lettau and Ludvigson (2001b).

We run and compare various specifications using the Fama–MacBeth two-step approach. Table 2 summarizes estimates of λ_s in the second stage of the Fama–MacBeth regressions. Results for the Japanese data reported in this table exhibit some similarities to and differences from the US results. As with the US results, the market portfolio Rvw_t has almost no explanatory power for the cross-section of stock returns in the Japanese data (Row 1). On the other hand, the Fama–French three-factor model (Row 5) exhibits good performance. Labor income also has some explanatory power when it is included along with the market portfolio, a result also found in the US data (Jagannathan, Kubota and Takehara, 1998). However, the estimated coefficients of labor income growth YG_t have negative signs, which is puzzling and contradicts what theory suggests.

[Table 2 is about here.]

We also examine term premium and $\widehat{cay}_{t-1}$ as conditioning variables. The term premium is statistically significant when it is used along with the market portfolio and/or labor income growth. However, it loses its explanatory power when included with SMB and HML.

The biggest difference between the US results by Lettau and Ludvigson and ours using Japanese data is the role of $\widehat{cay}_{t-1}$. In US results, $\widehat{cay}_{t-1}$ is significant as a scaling variable in the scaled factors models (corresponding to Row 4 in Table 2) but not as a conditioning variable or a risk factor. Table 2 suggests an opposite result for Japan. In the Japanese data, $\widehat{cay}_{t-1}$ is statistically significant as a conditioning variable.

To understand why this is so, in Figure 1, the Fama–French factors (SMB and HML) and $\widehat{cay}_{t-1}$ are plotted. From these graphs, the movements of the $\widehat{cay}$ variables are clearly much more persistent than the movements of SMB and HML. In Table 3, we show autocorrelations of the Japanese and the US. $\widehat{cay}$. It is clear that $\widehat{cay}_t$ for Japan exhibits much higher persistence than for the US. Therefore, the fluctuations of market conditions captured by $\widehat{cay}$ for Japan exhibit much larger swings compared with the US market. This evidence suggests that $\widehat{cay}_{t-1}$ should be characterized as a conditioning variable rather than a risk factor in the Japanese case. Our interpretation is that, for the Japanese data, $\widehat{cay}_{t-1}$ identifies the different phases of market conditions that are best described as regime changes in the constant term. In the US case, on the other hand, Lettau and Ludvigson (2001b) suggest that $\widehat{cay}_{t-1}$ identifies cyclical variation of the market beta or the conditional beta.

[Figure 1 and Table 3 are about here.]

Next, in Table 4, we report the results for real estate wealth. In part (A) of Table 4, $\widehat{cay}_{t-1}$ is tabulated from equation (12), including the ratio of real estate wealth to total wealth. On the other hand, in part (B), calculation of $\widehat{cay}_{t-1}$ is based on (13), which includes the log of the JREI land price index as an independent regressor. These results suggests significant improvement on average in Table 2's results with only financial wealth included in the regression. For example, the specification with Fama–French factors plus $\widehat{cay}_{t-1}$ in Row 6 in Table 2 ($R^2 = 48.2$) corresponds to A1 ($R^2 = 55.0$) and B1 ($R^2 = 52.8$) in Table 4. Therefore, R^2 increases and the sum of squared residuals decreases. Similarly, the specification including the term premium, Row 9 in Table 2 ($R^2 = 48.5$), corresponds to A3 ($R^2 = 56.4$) and

B3 ($R^2 = 53.6$) in Table 4. Hence, we can safely say that $\widehat{cay}_{t-1}$ calculated with real estate wealth explain the cross-sectional pattern of the Japanese stock market better.

[Table 4 is about here.]

A comparison between the case in which real estate is included in total assets and the case where it enters separately in the cointegration regression is difficult (see Tables 4 (A) and (B)). When the scaled factor $\widehat{cay}_{t-1} \cdot Rvw_t$ is included, performance of the regressions in (B) increases significantly (B2 and B4) and outperforms all specifications in part (A). However, the scaled factor is always statistically insignificant. We are in favor of the results reported in Table 4 (B) in which the land price index is included separately in the cointegration regression. However, this is mainly because we are more comfortable with the construction of $\widehat{cay}_{t-1}$ by equation (13) than by (12).

4.3 Discussions

As explained in the two previous subsections, movements in the Japanese consumption–wealth ratio $\widehat{cay}$ are much more persistent than in the US case, reflecting the fact that the Japanese market experienced a large bubble and crash in the sample. Statistically, this means that $\widehat{cay}$ for Japan is close to a unit root.

In a broad sense, the consumption–wealth ratio can be thought of as a type of "financial ratio", along with dividend yields or price earnings ratio. These types of predicting variable are meant to measure the deviation of asset price from its "fundamental value." In the case of the US stock market examined by Lettau and Ludvigson, the deviation measured by $\widehat{cay}$ is relatively short lived, and so $\widehat{cay}$ is useful in predicting stock returns for an investment horizon of one quarter to a year. However, because $\widehat{cay}$ for Japan is much more persistent, it is not particularly useful in explaining short-run asset price dynamics.

The theoretical framework proposed by Lettau and Ludvigson imposes the "no bubble" condition (3) in deriving an expression for log consumption–wealth ratio. This condition rules out rational "bursting bubble"type deviations from the fundamentals (Blanchard [1979]; Blanchard and Watson [1982]), in favor of predictable fluctuations in market conditions or mean reversion. One possible interpretation of the deviation from the long-run cointegration relationship in Japan in the late 1980s is that this "no bubble" condition had been violated in this period.

Alternatively, we may interpret such a deviation as a reflection of the fact that the stock market and household consumption are only loosely connected in Japan. In this respect, Japan is not an outlier among developed economies. Correlations between stock returns and consumption are often very weak and sometimes even negative except for English-speaking countries, in particular Canada, the UK and the US. (see Campbell [2003], Table 4). Therefore, very large deviations measured by $\widehat{cay}$ can occur, and if they occur, the adjustment process will require a much longer time to turn back to the long-run equilibrium. Hence, stock returns may be predictable in Japan too for the very long run, for example, more than a five-year period. Because our sample size is small compared with the size of the fluctuations and the persistence of $\widehat{cay}$, statistical inference on such long-run predictability requires very careful treatment.

Both interpretations of the asset price bubble in the late 1980s provide sensible explanations for why $\widehat{cay}$ is not very helpful in explaining the short-run dynamics of the Japanese stock market. Unfortunately, they cannot be differentiated from the finite sample because both imply that asset prices eventually return to their fundamental values after a certain period of time.

5 Concluding remarks

In this paper, we examine whether the consumption–wealth ratio, more precisely, the deviation from its long-run cointegrating relationship, can explain Japanese stock market data. Following Lettau and Ludvigson (2001a,b), we carefully construct $\widehat{cay}_t$, the residuals from the cointegration relationship between consumption and total household wealth. Unlike the US results, $\widehat{cay}_t$ does not predict future Japanese stock returns. On the other hand, it provides some help in explaining the cross-section of stock returns of industry portfolios. In the US case, $\widehat{cay}_t$ is a scaling variable that explains time variation in the market beta. In the Japanese case, the movement of $\widehat{cay}_t$ is much more persistent and is interpreted as the change in the constant terms, and hence changes in average stock returns. As we discussed extensively in section 4, any empirical study of the Japanese stock market covering the bubble period always faces a fundamental difficulty because the sample contains such a significant one-off boom and bust. This appears as high persistence of the consumption–wealth ratio in our analysis.

We also augment the Japanese $\widehat{cay}$ variable by including real estate wealth. The consumption–wealth ratio including real estate is even more effective in explaining the cross-section of stock returns in Japan. Examining an augmented $\widehat{cay}$ variable with data from other countries will be an interesting subject of future research.

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Table 1

Forecasting quarterly stock returns

	Constant	lag	$c\hat{a}y$	PDR	RREL	TRM	$\bar{R}^2$
			Panel A: Real Returns;1974:1Q–2003:1Q				
1	0.009	0.316***					0.03
	(1.216)	(3.528)					
2	-0.037**		-0.940***				0.07
	(-2.113)		(-3.084)				
3	-0.026	0.258***	-0.671**				0.12
	(-1.491)	(2.822)	(-2.159)				
			Panel B: Excess Returns;1974:1Q–2003:1Q				
4	0.001	0.310					0.09
	(0.081)	(3.509)					
5	-0.035**		-0.694**				0.03
	(-1.981)		(-2.246)				
6	-0.024	0.277***	-0.474				0.10
	(-1.359)	(3.071)	(-1.544)				
			Panel C: Additional Controls; Excess Returns;1974:1Q–2003:1Q				
7	0.047		-1.099***	-0.022			0.09
	(0.449)		(-3.182)	(-1.022)			
8	0.059	0.235**	-0.597*	-0.022	-0.575	0.008	0.12
	(0.713)	(2.508)	(-1.828)	(-1.250)	(-0.774)	(1.146)	

Note: t-statistics are in parentheses.

Table 2

Fama-MacBeth regressions with Equity Only *CAY:* Full Sample

		Conditioning Variables		Factors				Scaled Factor		Sum of
	Constant	cay(−1) (x100)	term	Rvw	YG	SMB	HML	Rvw・cay(−1)	R2	Sq.Resid
1	0.82	0.82		−0.08					0.1	4.70
	(1.07)			(1.19)						
2	0.95			−0.11	−0.28 *				15.6	3.98
	(1.02)			(1.12)	(0.27)					
3	1.16 *	0.74 *		−0.28	−0.35 *				33.8	3.12
	(0.94)	(0.50)		(1.02)	(0.25)					
4	1.03 *	0.74 *			−0.35 *			0.00	34.2	3.10
	(1.00)	(0.50)			(0.25)			(0.06)		
5	2.00 *			−1.04		−0.61 *	−0.43 *		41.0	2.78
	(1.05)			(1.07)		(0.41)	(0.36)			
6	1.97 *	0.48 *		−1.01		−0.67 *	−0.43 *		48.2	2.44
	(1.01)	(0.47)		(1.03)		(0.40)	(0.35)			
7	1.34 *	0.84 *	0.28 *	−0.48	−0.48 *				39.7	2.84
	(0.95)	(0.51)	(0.27)	(1.04)	(0.30)					
8	2.06 *		0.07	−1.11		−0.64 *	−0.43 *		42.0	2.73
	(1.08)		(0.22)	(1.12)		(0.42)	(0.37)			
9	2.00 *	0.49 *	0.09	−1.05		−0.69 *	−0.42 *		48.5	2.42
	(1.05)	(0.49)	(0.21)	(1.08)		(0.41)	(0.36)			

Note: Fama-MacBeth regressions using 28 Industry Portfolios: Second-stage regression

Sample period: 1978:1Q–2003:1Q

Definition of variables

cay(−1)	Consumption-Wealth ratio (residuals from cointegration regression)
term	Term premium (10year JGB − Call rate)
Rvw	Value weighted market index
YG	Labor income growth
SMB	Fama-French SMB factor (return difference between size sorted portfolios)
HML	Fama-French HML factor (return difference between portfolios sorted by Book to Market ratio)

Table 3

Autocorrelations of the consumption–wealth ratio in Japan and the US

No. of quarters	1	2	3	5	10
Japan	0.92	0.88	0.84	0.71	0.46
US	0.85	0.74	0.67	0.52	0.15

Table 4
Fama-MacBeth regressions with *CAY* including Land: Full Sample

	Conditioning Variables						Scaled Factor		Sum of
	Constant	cay(−1) (x100)	term(x100)	Rvw	SMB	HML	Rvw・cay(−1)	R2	Sq.Resid
(A) Land included in Total Asset									
A1	2.24 *	0.45 *		−1.25 *	−0.67 *	−0.57 *		55.0	2.12
	(0.95)	(0.35)		(0.97)	(0.37)	(0.34)			
A2	1.99 *	0.50 *		−1.02	−0.61 *	−0.68 *	0.01	57.9	1.98
	(1.03)	(0.36)		(1.04)	(0.38)	(0.39)	(0.04)		
A3	2.23 *	0.44 *	0.14	−1.24 *	−0.66 *	−0.64 *		56.4	2.05
	(0.96)	(0.35)	(0.31)	(0.98)	(0.37)	(0.39)			
A4	2.01 *	0.49 *	0.12	−1.03	−0.61 *	−0.73 *	0.01	58.8	1.94
	(1.05)	(0.36)	(0.31)	(1.06)	(0.38)	(0.43)	(0.04)		
(B) Land is sepratly in cointegration regression									
B1	2.15 *	0.55 *		−1.18 *	−0.75 *	−0.32		52.8	2.22
	(0.97)	(0.42)		(0.99)	(0.39)	(0.35)			
B2	1.87 *	0.68 *		−0.90	−0.77 *	−0.56	−0.02	61.9	1.79
	(0.93)	(0.40)		(0.95)	(0.36)	(0.38)	(0.03)		
B3	2.14 *	0.51 *	0.09	−1.17 *	−0.73 *	−0.39		53.6	2.18
	(0.98)	(0.44)	(0.33)	(1.01)	(0.40)	(0.42)			
B4	1.85 *	0.64 *	0.11	−0.87	−0.75 *	−0.66 *	−0.02	63.2	1.73
	(0.94)	(0.41)	(0.30)	(0.96)	(0.37)	(0.45)	(0.03)		

Note: Fama-MacBeth regressions using 28 Industry Portfolios: Second-stage regression

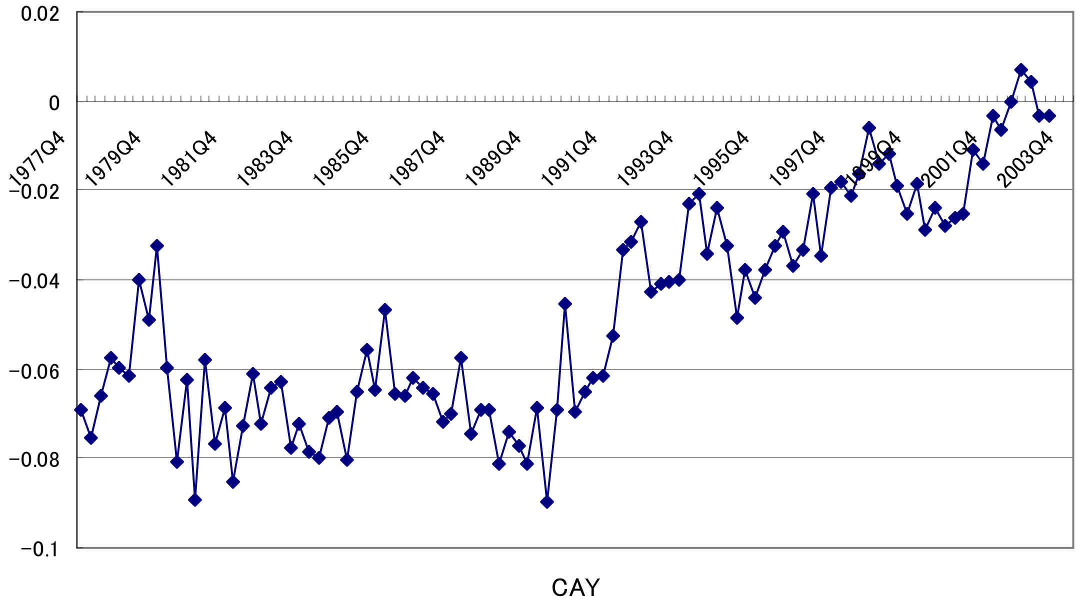
Sample period: 1978:1Q-2003:1Q

Definition of variables

cay(−1)	Consumption-Wealth ratio (residuals from cointegration regression)
term	Term premium (10year JGB − Call rate)
Rvw	Value weighted market index
YG	Labor income growth
SMB	Fama-French SMB factor (return difference between size sorted portfolios)
HML	Fama-French HML factor (return difference between portfolios sorted by Book to Market ratio)

Figure 1　Factors for Cross-section Regression

Panel A



Panel B

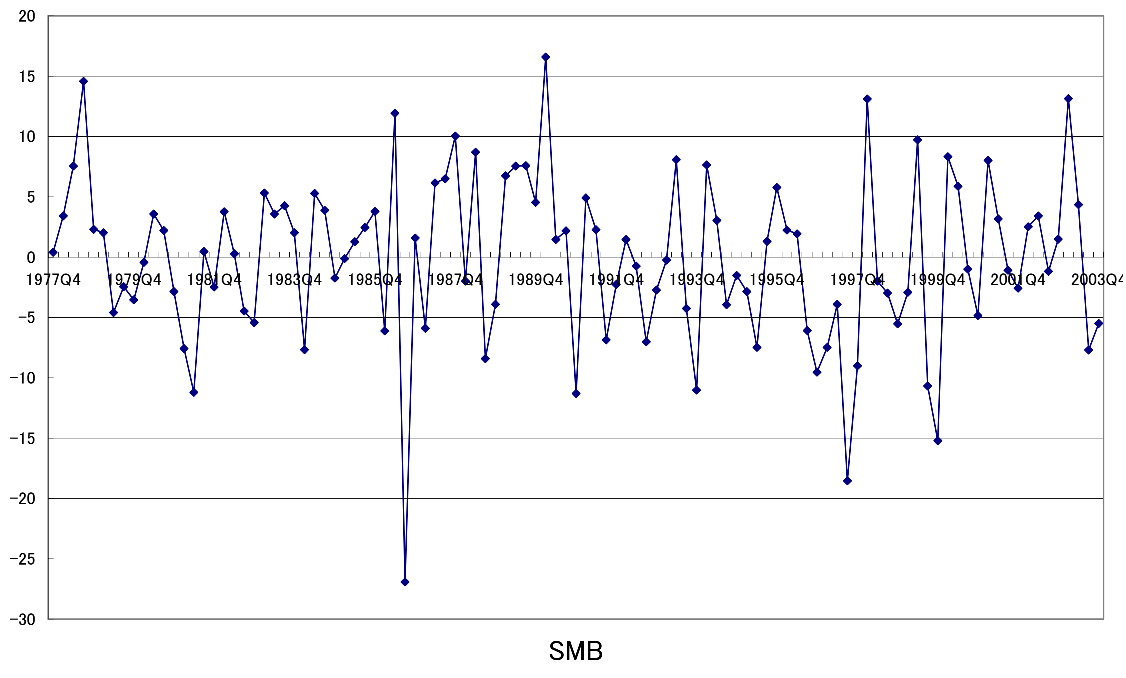
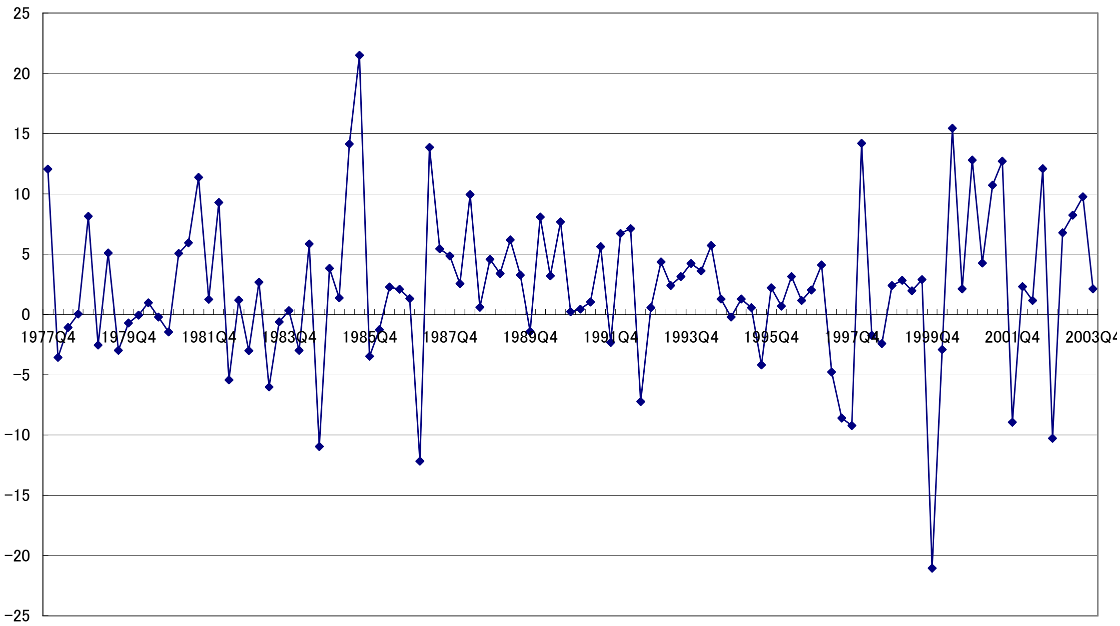


Figure 1 (continued)

Panel C



HML