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MONETARY POLICY AND ASSET VALUATION

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**ABSTRACT**

This paper presents evidence of infrequent shifts, or "breaks," in the mean of the consumption-wealth variable $cay_t$ that are strongly associated with low frequency fluctuations in the real value of the Federal Reserve's primary policy rate, with low policy rates associated with high asset valuations, and vice versa. By contrast, there is no evidence that infrequent shifts to high asset valuations and low policy rates are associated with higher expected economic growth or lower economic uncertainty; indeed the opposite is true. Additional evidence shows that low interest rate/high asset valuation regimes coincide with significantly lower equity market risk premia.

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# 1  Introduction

Asset values fluctuate widely around measures of economic fundamentals. Some of this variation is concentrated at high frequencies, as is apparent from the daily volatility in the stock market. One might reasonably attribute the origins of this variation to market "noise" around a more stable aggregate economic state. But what if a significant fraction of this variation is attributable to low frequency, decades-long shifts in these relative relationships? Such a phenomenon, were it to exist, could not be readily attributed to short-term volatility in the stock market, but would instead raise questions about the role of structural changes in the macroeconomy that govern how far and how persistently asset values can deviate from their historical relationship to measures of fundamental value.

This paper presents empirical evidence of just such a low frequency phenomenon. The consumption-wealth variable $cay_t$ is an estimated asset market valuation ratio that uses data on total household net worth (asset wealth), but its variation is driven primarily by movements in the stock market relative to two key macroeconomic fundamentals: consumer spending and labor income. In a 2001 published paper, Lettau and Ludvigson (2001) (LL hereafter) introduced this variable and found that it had strong forecasting power for U.S. stock returns. We refer to this asset valuation variable interchangeably as a "wealth ratio," since it measures how high or low asset values are relative to a linear combination of consumption and labor income.

In the years since the 2001 paper was published, the statistical properties of the estimated $cay_t$ series have shifted in some fundamental ways. Notably, the measured value of $cay_t$ has become more persistent over time, resulting in forecasting power for stock market returns increasingly concentrated at longer horizons and making it difficult, according to some statistical tests, to distinguish $cay_t$ from a unit root process. Mechanically, the reason for this has to do with the persistently high asset valuations of the post-millennial period, which have resulted in observations on $cay_t$ that have remained well below the variable's pre-2000 mean even in the aftermath of two large stock market crashes and one large housing market crash. Similar findings have been documented for other stock market valuation ratios long used as predictor variables for stock returns, including price-dividend or price-earnings ratios.[1] Despite these findings, a literal unit root interpretation for these variables is unappealing because it implies that stock prices or asset values could wander arbitrarily far from measures of fundamental value *indefinitely*.[2] An arguably more appealing interpretation is that there are instead infrequent

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[1] The near-unit root statistical properties of these ratios and their implications for return forecasting have been the subject of empirical work by Lewellen (2004), Campbell and Thompson (2008), Lettau, Ludvigson, and Wachter (2008), Lettau and Van Nieuwerburgh (2008), van Binsbergen and Koijen (2010), and Koijen and Van Nieuwerburgh (2011).

[2] This is unappealing even with presumed departures from conventional notions of market efficiency. Even theories that postulate "bubbles" almost always imply that the bubble will eventually burst, restoring a pre-

shifts in certain moments of the stationary distribution that–when not taken into account–make distinguishing a stationary from a unit root variable difficult in a small sample.

This paper presents evidence that these persistent changes in asset values relative to fundamentals as measured by $cay_t$, including the persistently high asset valuations of the post-2000 period, are largely driven by infrequent but stationary shifts, or "breaks," in the mean of $cay_t$. In addition, these shifts are found to coincide with quantitatively large structural changes in the long-run expected value of the real federal funds rate, and with evidence for changes in risk-taking behavior in equity markets.

To establish this evidence, we first adjust $cay_t$ for the infrequent shifts in its mean by estimating a Markov-switching version of the variable, denoted $cay_t^{MS}$. Formal model comparison tests based on the $BIC$ criterion show that $cay_t^{MS}$ describes the data far better than a model with fixed coefficients and no switches in the constant. We find that the sample is divided into three clear subperiods characterized by two regimes for the mean of $cay_t$: a low asset valuation regime that prevails from 1976:Q2 to 2001:Q2, and a high asset valuation regime that prevails in two subperiods at the beginning and end of our sample, namely 1952:Q1-1976:Q1, and the post-millennial period 2001:Q2-2013:Q3. Though persistent, the fluctuations captured by our estimated regime switches are not permanent. Unlike the conventional $cay_t$, which presumes a constant mean, $cay_t^{MS}$ does not exhibit increasing persistence as estimates are updated over the sample to include the post-millennial period (the original estimation used data through 1998). Moreover, evidence in favor of stationarity for $cay_t^{MS}$ is much stronger in current samples than it is for $cay_t$. This implies that infrequent shifts in the mean of $cay_t$ largely explain why its statistical properties have shifted over time.

We find that the forecasting power of $cay_t^{MS}$ for future stock market returns is superior to that of $cay_t$, even if no forward-looking data are used in the construction of $cay_t^{MS}$. This remains true out-of-sample, at least for some forecast horizons. Forecasts are improved because adjustments in the mean imply that estimates of conditional expectations do not mix data across regimes characterized by very different structural relationships between the level of $cay_t$ and future asset returns.

We then direct our attention to the key question of what these infrequent mean shifts represent economically. Any estimated statistical relationship is subject to possible structural change as the number of years over which the relationship is measured rises. But structural shifts in the economy are also likely to play a role, as suggested by evidence that other stock market valuation measures have also experienced "breaks" in the mean values of their distributions (e.g., Lettau, Ludvigson, and Wachter (2008); Lettau and Van Nieuwerburgh (2008)), and we confirm here that five other valuation ratios exhibit similar high-low-high valuation patterns over the same subperiods that characterize the $cay_t^{MS}$ regimes and at frequencies that roughly

bubble relationship between prices and fundamentals.

line up with the persistence of those regimes.

We are interested in the macroeconomic origins of these lower frequency shifts in valuation ratios. We therefore estimate a macroeconomic Markov-switching vector autoregression (MS-VAR) for output growth, inflation, investment growth, research-and-development (R&D) growth, and the federal funds rate in which we allow the parameters of the MS-VAR to potentially undergo structural breaks. Importantly, we force these breaks to coincide with the periods corresponding to the shifts identified from our estimates of $cay^{MS}$, but the parameters characterizing the different regimes as well as the transition matrix are freely estimated. We then compute conditional moments of the macro variables implied by the MS-VAR, taking into account the possibility of regime changes. In doing so, we find evidence of strikingly large breaks in the expected real federal funds rate that coincide with the breaks in the mean of $cay$, with low wealth ratios (low asset valuations or high $cay$) associated with an expectation of persistently high values for the real federal funds rate, and high wealth ratios (high asset valuations or low $cay$) associated with an expectation of persistently low values for the real federal funds rate. The post-millennial period in particular, characterized by high asset valuations according to a number of indicators, is marked by expectations of prolonged (but not permanent) low values for the real federal funds rate, in contrast to the middle subperiod where asset valuations were low and expected policy rates were high.

We find no evidence, however, that these low frequency shifts to high asset valuations and persistently low policy rates are associated with higher expected economic growth over any horizon, or lower economic uncertainty; indeed the opposite is true. The high asset valuation subperiods are times of relative weakness in GDP growth, investment growth, and R&D growth, and relatively high macroeconomic uncertainty. The findings therefore run counter to the idea that high asset valuations associated with a persistently low interest rate environment are the result of a positive outlook for economic growth, or lower uncertainty about that growth. Moreover, evidence from a second MS-VAR that employs data on the labor compensation share of GDP and stock market dividend growth indicates that the high valuation subperiods are associated with an expectation of persistent declines in the labor share, as opposed to the low valuation subperiod in which the share is not expected to fall. The one exception to these findings of broad-based economic weakness during high asset valuation regimes is the stock market itself: high asset valuation subperiods are characterized by expectations of persistently higher dividend growth on publicly traded shares.

These findings raise a question: why is the Central Bank's expected policy rate associated with such pronounced low frequency shifts in asset valuations? Given that the findings are so closely tied to the behavior of the expected real interest rate, one theoretical answer is that the results at least partly reflect corresponding regime shifts in discount rates, and indeed our evidence is consistent with this interpretation. But even if we restrict attention to interpreta-

tions based on changing discount rates, theories differ on the reasons discount rates change with interest rates. Some theories tie low and declining discount rates entirely to the behavior of the risk-free rate, which is presumed to be driven endogenously downward by shocks that increase the fraction of wealth held in the hands of more risk averse or more pessimistic investors (e.g., Barro and Mollerus (2014); Caballero and Farhi (2014); Hall (2016)). In these theories, risk premia rise as the risk-free rate declines, implying that asset valuations can only be higher if the decline in the risk-free rate exceeds the rise in risk premia. Other theories imply that shifts downward in the risk-free rate coincide with shifts downward in return premia, as in theories that can be broadly characterized as having a "reaching for yield" channel (e.g., Rajan (2006); Rajan (2013); Drechsler, Savov, and Schnabl (2014); Piazzesi and Schneider (2015); Acharya and Naqvi (2016)).

Our empirical findings imply that the high asset valuation episodes we document are not well described as periods during which lower interest rates coincided with higher risk premia. On the contrary, we present evidence from equity markets that low interest rate regimes coincide with *lower* rather than higher risk premia, consistent with reaching for yield. Specifically, we find that, in a switch from a high to low interest rate regime, the estimated risk premia of 14 out of 17 portfolios that carry a positive risk premium on average simultaneously fall to lower levels. Moreover, both the risk premia and the book-market ratios (adjusted for expected earnings) of evidently riskier, higher Sharpe ratio portfolios, such as those that go long in value stocks or stocks that have recently appreciated the most, fall much more than those of evidently less risky, lower Sharpe ratio portfolios, such as those that go long in growth stocks or stocks that have recently appreciated the least. For several portfolios, the estimated risk premia reach lows or near-lows early in the post-2000 period and, after a brief spike upward in the 2007-08 financial crisis, again in the post-2009 period when interest rates entered the zero-lower-bound range. Our findings for stock market returns in this regard are reminiscent of recent evidence of reaching for yield in the Treasury market (e.g., Hanson and Stein (2015)), by U.S. prime money funds (e.g., Di Maggio and Kacperczyk (2015)), and by U.S. corporate bond mutual funds (Choi and Kronlund (2015)). The evidence in these papers pertains to heavily intermediated asset classes. By contrast, our evidence pertains to equity market portfolios, an asset class ostensibly held by retail investors and households, as well as intermediaries.

A growing empirical literature documents a linkage between monetary policymaking activities and financial returns in high frequency data, using either formal event studies and daily data (Cook (1989); Bernanke and Kuttner (2005); Gürkaynak, Sack, and Swanson (2005)) or by studying the timing of when premia in the aggregate stock market are earned in weeks related to FOMC-cycle time (Lucca and Moench (2015); Cieslak, Morse, and Vissing-Jorgensen (2015)) or by directly forecasting weekly stock returns using weekly observations on federal funds futures implied rates (Neuhierl and Weber (2016)). To the best of our knowledge, the

findings presented in this paper are the first formal statistical evidence that lower frequency structural shifts in equity market return premia and asset values relative to economic fundamentals are strongly associated with persistent regimes for the primary policy instrument under direct control of the central monetary authority.

We take the federal funds rate to be a core tool of conventional monetary policy over which the Federal Reserve has direct control, one with the benefit that it is readily observed over most of the post-war period. Implicit in our use of the term "monetary policy" with regard to this tool is the assumption that the Central Bank can affect expected real rates for extended periods of time, even if it cannot do so forever.[3] Our approach makes no attempt to identify unexpected policy "shocks" that are independent of observed economic conditions. The Central Bank's predictable reactions to economic fluctuations are among the elements of deliberate policymaking that we wish to capture with our lower frequency analysis. For example, a persistent disturbance that generates secular stagnation could lead both private agents and the monetary authority to expect a prolonged period of low growth. In which case, the Central Bank might be expected to accommodate this state by keeping policy rates low for an extended time. Yet the choice of how long and how extensively to accommodate a persistent economic disturbance is itself important discretionary tool of the monetary authority, especially when there is significant uncertainty or disagreement about the economic outlook. Consistent with this hypothesis, we find that one measure of the stance of monetary policy–namely the long-term response of the nominal federal funds rate to a permanent change in inflation or output growth–differs significantly across the two regimes, with the high interest rate/low asset valuation regime characterized by a funds rate that is more responsive to inflation, and the low interest rate/high asset valuation regime characterized by a funds rate that is more responsive to output growth. The results here, spanning three clear subperiods over 50 years, suggest that persistent interest rate policies aimed at achieving specific macroeconomic objectives for inflation and/or growth may have unplanned consequences for asset valuations and risk premia.

The rest of the paper is organized as follows. The next section discusses the empirical model and the estimation of a Markov-switching $cay_t$. Section 3 presents results from this estimation, including evidence of breaks in the mean of $cay$, evidence on the persistence of $cay$ once corrected for regime shifts in its mean, and a comparison of the forecasting power of $cay^{MS}$ and $cay$ for U.S. stock market returns. These investigations are all designed to address the question of whether or not infrequent shifts in the mean of $cay_t$ can econometrically

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[3]The persistently high values of the real federal funds rate in the late 1970s and into 1980s were arguabley engineered to build credibility for lower inflation, while the opposite goal–building credibility for higher inflation– might be labeled an objective of the post-millennial period, especially after the Great Recession. Indeed, the assumption that the Central Bank can and in some cases should influence real rates for extended periods underlies at least one ostensive objective of another monetary policy tool, namely forward guidance, one use of which is to re-anchor expectations over the extended future around a new future path for inflation and therefore real rates (e.g., Eggertsson and Woodford (2003)).

account for its changing statistical properties over the last 15 years. This is an important first step in understanding the macroeconomic sources of highly persistent fluctuations in asset market valuations, since if we cannot pinpoint the statistical mechanisms behind the changing behavior, we cannot hope to explain it. With this evidence in hand, section 4 turns to the key economic question of what macroeconomic forces lie behind the regime shifts in the mean of $cay_t$. Section 5 adds to this evidence by empirically evaluating the hypothesis that some of the shift to high asset valuation regimes is attributable to a decline in risk premia in equity market assets. Section 6 briefly remarks on the behavior of other stock market valuation ratios over our estimated regime subperiods and Section 7 concludes. A large amount of additional material and test results have been placed in an Appendix for online publication.

## 2  Econometric Model of $cay^{MS}$

The variable studied by LL, denoted $cay_t$, is a stationary linear combination of log consumer spending, $c_t$, log asset wealth, $a_t$, and log labor income, $y_t$, all measured on an aggregate basis. Under assumptions described in LL and elaborated on in Lettau and Ludvigson (2010), $cay_t$ may be interpreted as a proxy for the log consumption- aggregate (human and non-human asset) wealth ratio, and its relationship with future growth rates of $a_t$ (highly correlated with stock market returns in quarterly data) and/or future growth rates of $c_t$ and $y_t$, may be motivated from an aggregate household budget constraint. Specifically, if labor income is the dividend paid to human capital, we can derive an approximate expression linking $c_t$, $a_t$, and $y_t$ to expected future returns to asset wealth, consumption growth, and labor income growth:

$$cay_t \equiv c_t - \gamma_a a_t - \gamma_y y_t \approx \mathbb{E}_t \sum_{i=1}^{\infty} \rho_w^i \left( (1 - \nu) \, r_{a,t+i} - \Delta c_{t+i} + \nu \Delta y_{t+1+i} \right),$$

where $\nu$ is the steady state ratio of human wealth to asset wealth and $r_{a,t}$ is the log return to asset wealth. Theory implies that $c_t$, $a_t$, and $y_t$ should be cointegrated, or that $cay_t$ should be covariance stationary. Asset wealth returns on the right-hand-side can be tautologically decomposed into the sum of a risk premium component and risk-free rate component, i.e., $\mathbb{E}_t r_{a,t+i} \equiv \mathbb{E}_t \left( r_{a,t+i} - f_{t+i} \right) + \mathbb{E}_t f_{t+i}$, where $f_t$ is the risk-free rate. Historically, almost all of the short-medium frequency variation in the traditional measure of $cay_t$ on the left-hand-side has been driven not by future consumption, labor income growth or risk-free rates, but by the first component of this decomposition, namely the asset wealth risk-premium.

This section presents the econometric model of regime switches in the mean of $cay_t$. In the standard estimation without regime shifts in any parameters, the stationary linear combination of $c_t$, $a_t$, and $y_t$ can be written

$$c_t = \alpha + \gamma_a a_t + \gamma_y y_t + \epsilon_t, \tag{1}$$

where the parameters to be estimated are $\alpha$, $\gamma_a$, and $\gamma_y$. The residual $\epsilon_t$ is the stationary linear combination of these data, referred to as the cointegrating residual. In this paper, we estimate a Markov-switching version of this cointegrating relationship, analogously written as

$$c_t = \alpha_{\xi_t^\alpha} + \beta_a a_t + \beta_y y_t + e_t,$$

where the notation $\alpha_{\xi_t^\alpha}$ indicates that the value of the constant depends on the existence of a latent state variable, $\xi_t^\alpha$, presumed to follow a two-state Markov-switching process with transition matrix $\boldsymbol{H}^\alpha$. This implies that the constant term $\alpha_{\xi_t^\alpha}$ can assume one of two discrete values, $\alpha_1$ or $\alpha_2$. The parameters $\beta_a$ and $\beta_y$ are the slope coefficients in this cointegrating relationship, analogous to $\gamma_a$ and $\gamma_y$ in the fixed coefficient regression (1). The residual $e_t$ is again a stationary variable by assumption.

The standard approach to estimating a single cointegrating relation such as (1) is to run a dynamic least squares regression (DLS–Stock and Watson (1993)) that controls for leads and lags of the right-hand-side variables in order to adjust for the asymptotic inefficiencies attributable to regressor endogeneity. These inefficiencies are only relevant in finite samples. Let $\boldsymbol{z}_t$ be a $3 \times 1$ vector of data on $c_t$, $a_t$, and $y_t$, and $k$ leads and $k$ lags of $\Delta a_t$ and $\Delta y_t$ and let $\boldsymbol{Z}_t = (\boldsymbol{z}_t, \boldsymbol{z}_{t-1}, ..., \boldsymbol{z}_1)$ be a vector containing all observations obtained through date $t$. To estimate the parameters of this stationary linear combination we modify the standard fixed coefficient DLS regression to allow for shifts in the intercept $\alpha_{\xi_t^\alpha}$:

$$c_t = \alpha_{\xi_t^\alpha} + \beta_a a_t + \beta_y y_t + \sum_{i=-k}^{k} b_{a,i} \Delta a_{t+i} + \sum_{i=-k}^{k} b_{y,i} \Delta y_{t+i} + \sigma \varepsilon_t \tag{2}$$

where $\varepsilon_t \sim N(0,1)$. The parameters of the econometric model include the cointegrating parameters and additional slope coefficients $\beta = \left(\beta_a, \beta_y, b\right)'$, where $b = (b_{a,-k}, .., b_{a,k}, b_{y,-k}, .., b_{y,k})'$, the two intercept values $\alpha_1$ and $\alpha_2$, the standard deviation of the residual $\sigma$, and the transition probabilities contained in the matrix $\boldsymbol{H}^\alpha$. Collect these parameters into a vector $\boldsymbol{\theta} = \left(\beta, \alpha_{\xi_t^\alpha}, \sigma, H^\alpha\right)'$.

Absent regime changes, *cay* is defined as:

$$cay_t^{FC} = c_t - (\alpha + \gamma_a a_t + \gamma_y y_t) \tag{3}$$

where the superscript "$FC$" stands for "fixed coefficients" because the constant $\alpha$ is fixed over time. Notice that when we impose a single regime, the Markov-switching model collapses back to the specification originally used by LL. The variable $cay_t^{FC}$ is the same as that defined in LL where it was denoted $cay_t$. For the purposes of his paper, we have added the superscript "$FC$" in order to explicitly distinguish it from the Markov-switching version. The parameters $\boldsymbol{\theta}$ of the time-series model for $cay_t^{FC}$ include the cointegrating parameters $\gamma_a$ and $\gamma_y$, the additional slope coefficients on the leads and lags in the DLS regression, and the single intercept value $\alpha$.

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Let $T$ be the sample size used in the estimation accounting for leads and lags in the regression. For the Markov-switching model, the constant $\alpha_{\xi_t^\alpha}$ depends on the regime $\xi_t^\alpha$. If the sequence $\xi^{\alpha,T} = \{\xi_1^\alpha, ..., \xi_T^\alpha\}$ of regimes in place at each point in time were observed, we could immediately compute $cay_t^{MS}$. Unfortunately, $\xi^{\alpha,T}$ is generally unobservable and needs to be inferred together with the other parameters of the model. It follows that the two values for the Markov-switching constant $\alpha_{\xi_t^\alpha}$ ($\alpha_1$ and $\alpha_2$) must be weighted by their probabilities at each point in time. For this purpose, we consider two estimates of the state probabilities distinguished as *filtered* or *smoothed* probabilities. Let $P\left(\xi_t^\alpha = i | \mathbf{Z}_t; \boldsymbol{\theta}\right) \equiv \pi_{t|t}^i$ denote the probability that $\xi_t^\alpha = i$ based on data obtained through date $t$ and knowledge of the parameters $\boldsymbol{\theta}$. We refer to these as filtered probabilities. Smoothed probabilities reflect the information about the state at time $t$ that can be extracted from the whole sample: $P\left(\xi_t^\alpha = i | \mathbf{Z}_T; \boldsymbol{\theta}\right) \equiv \pi_{t|T}^i$.[4] These measures of the regime probabilities may be used to construct two versions of a Markov-switching $cay$, based on using either smoothed or filtered probabilities. In both cases, the intercept coefficient for $cay$ is a probability weighted average of the two intercepts, $\alpha_1$ and $\alpha_2$. As a benchmark, we use the smoothed probabilities for our baseline estimate and denote it $cay_t^{MS}$. When we use filtered probabilities, we use the notation $cay_t^{MSfilt}$. Thus, $cay_t^{MS}$ is computed one of two ways:

$$cay_t^{MSfilt} \;=\; c_t - \left(\textstyle\sum_{i=1}^2 \pi_{t|t}^i \alpha_i + \beta_a a_t + \beta_y y_t\right). \tag{4}$$

$$cay_t^{MS} \;=\; c_t - \left(\textstyle\sum_{i=1}^2 \pi_{t|T}^i \alpha_i + \beta_a a_t + \beta_y y_t\right). \tag{5}$$

The econometric model (2) permits regime switches only in the intercept parameter. (The Appendix discusses alternative specifications in which $\sigma$ is also subject to regime switches.) This specification for $cay_t^{MS}$ maintains the hypothesis that a stationary linear combination of $c_t$, $a_t$, and $y_t$ exists, just as in the standard cointegration specification and consistent with the aggregate budget constraint motivation given in LL. The difference is that, here, some of the variation in the cointegrating residual is explicitly modeled via regime switches in its mean. To see this simply rewrite the above expressions for $cay_t^{FC}$ and $cay_t^{MS}$ to be inclusive of the intercept term, i.e.,

$$cay_t^{MS} + \textstyle\sum_{i=1}^2 \pi_{t|T}^i \alpha_i \;=\; \underbrace{c_t - \beta_a a_t - \beta_y y_t}_{\text{inv. wealth ratio}} \tag{6}$$

$$cay_t^{FC} + \alpha \;=\; c_t - \gamma_a a_t - \gamma_y y_t. \tag{7}$$

The right-hand-side of both equations is presumed stationary. In each case, the intercept term may be interpreted as the mean of a stationary linear combination of $c_t$, $a_t$, and $y_t$. To the extent that shifts in the intercept term $\alpha_i$ are driven by a stationary but persistent

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[4]In using the DLS regression (2) to estimate cointegrating parameters, we lose 6 leads and 6 lags. The Appendix on computing $cay^{MS}$ explains how we filter the data to obtain esitmates of the regime probabilities over the whole sample.

regime changes, $cay_t^{MS}$ will be less persistent than $cay_t^{FC}$. This is because the low frequency fluctuations of the right-hand-side of (6) are captured by the shifts in the constant, while the low frequency fluctuations of the right-hand-side of (7) must be entirely captured by $cay_t^{FC}$. The stationary linear combinations may be interpreted as a log inverse asset valuation ratio or inverse wealth ratio, akin to a log dividend-price ratio as opposed to log price-dividend ratio. Since, in population, $cay_t^{MS}$ and $cay_t^{FC}$ are mean zero random variables, the intercept terms give the mean of these inverse asset valuation ratios.[5] A high $\alpha_i$ corresponds to a low mean valuation ratio, since the residual $c_t - \beta_a a_t - \beta_y y_t$ is high whenever the value of wealth $a_t$ is low relative to the implied linear combination of $c_t$ and $y_t$ with which it is cointegrated. We refer to the log inverse asset valuation ratio on the right-hand-side of (6) interchangeably as the inverse wealth ratio, or equivalently define the (log) *wealth ratio* as $-\left[ cay_t^{MS} + \sum_{i=1}^{2} \pi_{t|T}^i \alpha_i \right]$.

## 2.1 Estimation

We use Bayesian methods with flat priors to evaluate the regression parameters in (2). We first search for the posterior mode using a maximization algorithm. The posterior of the model and the corresponding regime probabilities $\pi_{t|t}^i$ and $\pi_{t|T}^i$ are obtained by computing the likelihood using the Hamilton filter (Hamilton (1994)), and combining it with priors. Since we use flat priors, the posterior coincides with the likelihood. Our estimate of $cay_t^{MS}$ is based on the posterior mode of the parameter vector $\boldsymbol{\theta}$ and the corresponding regime probabilities. We further use a Gibbs sampling algorithm to do inference on the parameters and compute additional statistics of interest. In particular, uncertainty about the parameters, or about any desired transformation of the model parameters, can easily be characterized using the posterior distribution for the model parameters obtained with the Gibbs sampling algorithm. The full statement of the procedure and sampling algorithm is given in the Appendix. The Data Appendix provides a description of the data and sources for all series used in the paper.

# 3  Results: Breaks in the Mean of the Wealth Ratio

We estimate the Markov-switching cointegrating relation described by (2) over the sample 1952:Q1-2013:Q3 using six leads and lags. Table 1 reports the parameter estimates, while Figure 1 reports the probability of regime 1 for the Markov-switching intercept $\alpha_{\xi_t^\alpha}$ based on the posterior mode parameter estimates. The 90% credible sets are obtained making 2,000,000 draws from the posterior using the Gibbs sampling algorithm described in the Appendix. One in every one thousand draws is retained. We check convergence using the methods suggested

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[5]In a finite sample, $cay_t^{MS}$ and $cay_t^{FC}$ are not necessarily mean zero because of the leads and lags of the first differences included in the DLS regression. In population these variables are mean-zero by definition.

by Geweke (1992) and Raftery and Lewis (1992).[6] Formal model comparison tests based on the $BIC$ criterion show that $cay_t^{MS}$ describes the data far better than a model with fixed coefficients and no switches in the constant. These results are presented in the "Model Comparison" section, Appendix 7, of the Appendix.

The sample is divided into three clear subperiods characterized by two regimes for $\alpha$. Regime 1 is a high $\alpha$ regime with the posterior mode point estimate equal to $\hat{\alpha}_1 = 0.9186$. The low $\alpha$ regime 2 posterior mode estimate is $\hat{\alpha}_2 = 0.8808$. A high $\alpha$ regime for $cay$ corresponds to a low valuation ratio for the stock market, analogous to a low price-dividend ratio (Lettau and Ludvigson (2001)). Thus we shall refer to *high $\alpha$ regime 1* as the *low asset valuation regime*, and *low $\alpha$ regime 2* as the *high asset valuation regime*. Figure 1 shows that the low asset valuation regime prevails for a prolonged period of time starting from 1976:Q2 to 2001:Q2. The smoothed probability that $\alpha = \hat{\alpha}_1$ is very close to unity during this period. By contrast, the pre-1976 and post-2001 subsamples are high asset valuation regimes, where the probability that $\alpha = \alpha_1$ is virtually 0. These correspond to the subperiods 1952:Q1-1976:Q1, and 2001:Q2-2013:Q3, respectively.

Table 1 provides estimates of the difference between the high and low $\alpha$ and its distribution. The difference is positive and statistically significant, as exemplified by the third row of Table 1, which shows that a 90% credible set only contains non-zero and positive values for this difference.[7] The two regimes turn out to be very persistent as reflected in the estimates for the diagonal elements of the transition matrix $H^\alpha$, also reported in Table 1.

The mode values for the other cointegration parameters are $\beta_a = 0.26$ and $\beta_y = 0.62$. These values are comparable with those originally obtained by LL using a fixed coefficient regression ($\gamma_a = 0.31$ and $\gamma_y = 0.59$). By contrast, Table 2 reports the parameter estimates for the fixed coefficient cointegrating relation over the extended sample used in this paper, where $\gamma_a = 0.12$ and $\gamma_y = 0.78$. Therefore, in our current sample, the fixed coefficient parameter estimates differ substantially from those reported in 2001. Bearing in mind that deviations from the cointegrating relation are the result of persistent but transitory movements in $a_t$ rather than $c_t$ or $y_t$ (Lettau and Ludvigson (2004), Lettau and Ludvigson (2013)), these results suggest that the fixed-coefficient estimates of $cay_t$ attempted to "compensate" for increasingly persistent deviations in $a_t$ from its cointegrating relation with $c_t$ and $y_t$, by progressively reducing the

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[6]For Raftery and Lewis (1992) we target 90% credible sets, with a 1% accuracy to be achieved with a 95% minimum probaility. We initialize the Gibbs sampling algorithm making a draw around the posterior mode. Sims and Zha (2006) point out that in Markov-switching models it is important to first find the posterior mode and then use it as a starting point for the MCMC algorithm due to the fact that the likelihood can have multiple peaks.

[7]The Gibbs sampling algorithm is used to generate a distribution for the difference between the two means in the same manner it is used to generate a distribution for any parameter. For each draw from the joint distribution of the model parameters, we compute the difference and store it. We may then compute means and/or medians, and error bands, as for any other parameter of interest.

weight on $a_t$ and increasing the weight on $y_t$. The instability in these point estimates is largely eliminated by allowing for discrete shifts in the mean of *cay*.

To give a visual impression of the properties of these regimes, Figure 2 plots $cay_t^{MS} + \sum_{i=1}^2 \pi_{t|T}^i \alpha_i$ over time, which is the estimated Markov-switching *cay* from (5) *inclusive* of the intercept. Also plotted as horizontal lines are the values $\hat{\alpha}_1$ and $\hat{\alpha}_2$ that arise in each regime over the sample. The figure shows that this valuation variable fluctuates around two distinct means in three separate periods of the sample, a low mean in the early part of the sample, a high mean in the middle, and a low mean again in the last part of the sample.

## 3.1   Persistence of $cay^{MS}$ versus $cay^{FC}$

The most salient change in the statistical properties of the estimated $cay_t^{FC}$ series in the time since it was originally introduced is that it has become more persistent. Figure 3 plots the fixed coefficient $cay_t^{FC}$ and the Markov-switching $cay_t^{MS}$ as defined in (3) and (5), respectively. (Unlike Figure 2, these values subtract the estimated $\alpha$ or probability-weighted $\alpha$, respectively.) The two vertical bars mark the beginning and the end of the time span during which the high $\alpha$ regime was most likely to be in place, according to the smoothed probability estimates. As Figure 3 shows, $cay_t^{FC}$ exhibits persistent deviations from zero, especially during the last subperiod 2001:Q2-2013:Q3, which coincides with the second appearance in our sample of the high asset valuation regime. The estimated $cay_t^{FC}$ also exhibits persistent deviations from zero during the period starting around 1980 and ending in the early 2000s, roughly coinciding with the occurrence of the low asset valuation regime. Most of this subperiod was included in the original estimation of $cay_t^{FC}$, so it has contributed less to the growth in the persistence in the series since that time.

It is evident from Figure 3 that $cay_t^{MS}$ is quite different from $cay_t^{FC}$ in that it does not exhibit such persistent deviations from its demeaned value of zero. The reason is that the persistent deviations are instead captured by low-frequency regime changes in the constant of the cointegrating relation. To formalize this visual impression, the first column of Table 3 reports the first-order autoregressive coefficient estimate for the two versions of *cay*. The estimated autocorrelation coefficient for $cay_t^{FC}$ is 0.94. The estimated first-order autocorrelation coefficient for $cay_t^{MS}$ is 0.81, which is close to the 0.79 estimated coefficient reported in LL. Allowing for low frequency mean shifts in the cointegrating relation largely restores the estimated persistence of *cay* to its original values.

Several other tests are employed to assess the degree of persistence in $cay_t^{MS}$ as compared to $cay_t^{FC}$. First, we apply an augmented Dickey-Fuller $t$ test to the estimated cointegrating residuals. The test statistics and corresponding critical values are reported in Table 3 . According to this test, the null hypothesis of no cointegration is rejected for the $cay_t^{MS}$ in every

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case, whereas the opposite is true for $cay_t^{FC}$. Second, we examine low frequency averages or "cosine transformations" of $cay$ to gauge its persistence following Muller and Watson (2008) and Watson (2013). The cosine transformation of $cay_t^{MS}$ displays a pattern much more consistent with an $I(0)$ series than that of $cay_t^{FC}$. These results, along with the estimation details, are presented in the Appendix. Third, we estimate fractionally integrated models for $cay_t^{MS}$ and $cay_t^{FC}$ in which $(1-L)^d cay_t^{(m)} = u_t$, where $L$ is the lag operator, $u_t$ is an $I(0)$ process and $m = cay_t^{MS}, cay_t^{FC}$. If $cay_t^{(m)}$ is $I(0)$, then $d = 0$. If it has a unit root, then $d = 1$, and non-integer values of $d > 0$ are fractionally integrated series that are more persistent than $I(0)$ but less persistent than $I(1)$. Figure 4 shows the estimated log likelihoods for $(1-L)^d cay_t^{MS}$ and $(1-L)^d cay_t^{FC}$ as a function of $d$. For $cay_t^{MS}$, the likelihood peaks at $d = 0$, while for $cay_t^{FC}$, the likelihood rises with $d > 0$ and peaks near $d = 1.2$.[8]

## 3.2  Forecasts of Excess Stock Market Returns

The variable $cay$ has been used as a stock market forecasting variable because most of its variation has been driven historically by transitory fluctuations in $a$ around more stable values for $c$ and $y$. Thus when $a$ is high relative to $c$ and $y$, that signals lower values for future excess returns, rather than higher values for $c$ and/or $y$. It is therefore important to understand whether or not adjusting for the mean breaks in $cay$ improves its forecasting power. If so, it suggests that these fundamental relationships between asset wealth, consumption and labor income exist within regimes but not across regimes. If not it suggests a breakdown in the relationship entirely.

Table 4 reports the results of long-horizon forecasts of log returns on the CRSP value-weighted stock market index in excess of a three month Treasury bill rate. The table compares the forecasting power of $cay_t^{FC}$, and $cay_t^{MSfilt}$, based on filtered probabilities and $cay_t^{MS}$, based on smoothed probabilities. The top panel reports full sample forecasts. The bottom panel reports the results of forecasts based on fully recursive estimates of these measures using data only up to time $t$, denoted $cay^{FCrec}$ and $cay^{MSrec}$, respectively.[9] These variables are then used to forecast returns over the entire subsample from 1981:Q1-2013:Q3. The recursive estimates use no forward looking data to estimate any of the parameters, including the regime probabilities, regimes values, or transition probabilities. In both panels we report the coefficient estimates on the regressor, the Newey and West (1987) corrected $t$-statistic, and the adjusted $R^2$ statistic.

The top panel shows that all measures of $cay$ estimated over the full sample have statistically

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[8]Please refer to Appendix 7 for details about the estimation of the fractionally integrated model.

[9]The recursive estimates are obtained as follows. First, all parameters $\boldsymbol{\theta}$ for each model are estimated in an initial period using data available from 1952:Q1 through 1980:Q4. All parameters are then reestimated recursively on data from 1952:Q1-1981:Q1, 1952:Q1-1981:Q2, and so on, until the final recursive estimate of $cay$ is obtained based on data over the full sample 1952:Q1-2013:Q3. These variables are then used to forecast returns over the entire subsample from 1981:Q1-2013:Q3.

significant forecasting power for future excess stock market returns over horizons ranging from one to 16 quarters. But the coefficients, $t$-statistics and $R^2$ values are all larger using the Markov-switching versions $cay_t^{MSfilt}$ and $cay_t^{MS}$ than they are for $cay_t^{FC}$. The comparison is more stark if we compare recursively estimated values of $cay$ to full sample values. For example, the full sample estimate of $cay_t^{FC}$ explains 21% of the 16 quarter-ahead log excess stock market return in the subsample 1981:Q1-2013Q3, while $cay_t^{MSrec}$ explains 42%. Moreover, in this subsample, $cay_t^{FC}$ has little forecasting power for excess returns at all but the longest horizon, whereas $cay_t^{MSrec}$ has much stronger forecasting power.

Table 4 shows that $cay_t^{FCrec}$ also has stronger predictive power than $cay_t^{FC}$ over this subsample. By recursively estimating the parameters in $cay_t^{FC}$, we allow them to change *every period*. In this way, a recursively estimated fixed-coefficient model can "compete" with the Markov-switching version, which explicitly models shifts in the mean parameter. But finding that $cay_t^{FCrec}$ performs better than $cay_t^{FC}$ in forecasting returns hardly provides support for the hypothesis that the fixed-coefficient model is a better description of the data than the Markov-switching model. (Indeed, this hypothesis is explicitly rejected by a comparison based on the $BIC$ criterion of these models, as documented in the Appendix.) On the contrary, this finding may be taken as additional evidence of the instability in the fixed-coefficient parameters. If there were no such instability, $cay_t^{FCrec}$ would be identically equal to $cay_t^{FC}$.

Table 5 reports mean-square forecast errors (MSEs) from out-of-sample forecasts. The forecasting relation is estimated in an initial period using data available from 1952:Q1 through 1980:Q4. Forecasts over the next $h$ quarters are computed and forecast errors stored. The forecasting relation is then reestimated in rolling subsamples moving forward, (i.e., over the period 1952:Q1 through 1981:Q1), and forecasts and forecast errors are computed over the next $h$ periods. This process is repeated until the end of the sample. Table 5 reports MSEs for several forecasting regressions.

Table 5 shows that all versions of $cay$ also have lower lower MSEs than a simple autoregressive forecasting model or a model that uses only the (constant) sample mean of excess returns as a predictor. Among those versions that are estimated using the full sample, the two Markov-switching versions, $cay_t^{MSfilt}$, and $cay_t^{MS}$, are much better predictors than the fixed-mean version $cay_t^{FC}$, having MSEs that are almost 50% smaller for 16-quarter return forecasts. The recursively estimated versions $cay_t^{FCrec}$ and $cay_t^{MSrec}$ have about the same predictive power over most horizons, although the Markov-switching $cay$ offers a slight improvement over the fixed-mean $cay$ at the longest (16 quarter) horizon. Because these recursive versions are estimated over short subsamples, the estimates of parameters are much noisier than they are for the full-sample versions, so it is not surprising that they have higher MSEs. For this reason, it is notable that $cay_t^{MSrec}$ preforms as well (and slightly better at long horizons) as $cay_t^{FCrec}$, given that the former has many more parameters that require estimation over short subsamples

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of our quarterly dataset. Postwar samples of the size currently available are, however, much larger than the repeated subsamples used to construct the recursive estimates for this exercise. Going forward, such samples should provide less noisy estimates of $cay_t^{MS}$ parameters. Researchers using $cay_t$ as a predictor variable may wish to consider both measures as predictors of long-horizon stock market returns.[10]

# 4  What's Behind the Breaks in Asset Valuation?

For the rest of this paper, we search for empirical explanations from the macroeconomy for the breaks observed in $cay$. For example, perhaps the high valuation periods reflect an optimistic outlook for economic growth or lower macroeconomic uncertainty. We therefore carry out our investigation using macroeconomic data, studying how structural changes in the statistical properties of these data might be connected to the documented breaks in asset valuations as measured by $cay$.

To do so, we estimate a baseline MS-VAR using data on output growth, inflation, investment growth, R&D growth, and the federal funds rate, allowing the parameters of the VAR to potentially undergo structural breaks during the periods that correspond to the shifts identified in our estimates for $cay^{MS}$. Specifically, we impose the formerly estimated regime sequence for $cay$ on the VAR, but the parameters characterizing the different regimes, as well as the transition matrix, are freely estimated.[11] We denote the MS-VAR transition matrix $\boldsymbol{H}^A$ in order to distinguish it from the $cay^{MS}$ transition matrix $\boldsymbol{H}^\alpha$. Note that our objective is not to estimate independent regimes for the variables in the MS-VAR and see if they align with the previously estimated breaks in $cay$. Instead, the goal is to establish what, if anything, is different in the MS-VAR variables across the two previously estimated asset valuation regimes that could help explain the breaks in the mean of $cay$. We therefore deliberately "tie our hands" by forcing the regime sequence for the MS-VAR to correspond to breaks in $cay$. Note that under this restriction, there is no implication that the macro variables must necessarily show evidence of structural change. The point of this procedure is to ask whether the conditional moments of the macro variables in the MS-VAR show any evidence of important structural shifts under the previously estimated regime sequence, when they are freely estimated and could in principle show no shift.

All MS-VARs estimated in this section and the next are implemented using Bayesian methods with flat priors. The Appendix provides estimation details.[12]

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[10]Both series are regularly updated and available on the authors' websites.

[11]To impose the formerly estimated regime sequence, we choose the particular regime sequence $\hat{\xi}^{\alpha,T} = \{\hat{\xi}_1^\alpha, ..., \hat{\xi}_T^\alpha\}$ that is most likely to have occurred, given our estimated posterior mode parameter values for $\boldsymbol{\theta}$. See the Appendix for details.

[12]Bayesian methods are used because they offer significant computational advantages in characterizing un-

We consider the following MS-VAR model with $n$ variables and $m = 2$ regimes:

$$Z_t = c_{\xi_t} + A_{1,\xi_t} Z_{t-1} + A_{2,\xi_t} Z_{t-2} + V_{\xi_t} \varepsilon_t, \ \varepsilon_t \sim N(0, I) \tag{8}$$

where $Z_t$ is an $n \times 1$ vector of variables, $c_{\xi_t}$ is an $n \times 1$ vector of constants, $A_{l,\xi_t}$ for $l = 1, 2$ is an $n \times n$ matrix of coefficients, $V_{\xi_t} V'_{\xi_t}$ is an $n \times n$ covariance matrix for the $n \times 1$ vector of shocks $\varepsilon_t$. The process $\xi_t$ controls the regime that is in place at time $t$ and assumes two values, 1 and 2, based on the regime sequence identified in our estimates for $cay^{MS}$.

In the baseline MS-VAR, the vector $Z_t$ includes fives variables at quarterly frequency: GDP growth, inflation, investment growth, R&D growth, and the effective federal funds rate (FFR). Inflation is defined as the year-to-year differences of the logarithm of the GDP price deflator. GDP growth, investment growth, and R&D growth are defined as the year-to-year differences of the logarithm of real GDP per capita, real investment per capita, real R&D per capita, respectively. The quarterly FFR is obtained by taking the average of monthly figures of the effective federal funds rate. In a secondary MS-VAR, the vector $Z_t$ includes fives variables at quarterly frequency: GDP growth, inflation, the change in the labor compensation share of GDP ("labor share" for short), aggregate dividend growth for all firms traded on NYSE, NASDAQ and AMEX, and the effective federal funds rate. The Data Appendix provides a detailed description of our data and sources. The sample for this estimation spans the period 1955:Q3-2013:Q3.[13]

We are interested in knowing the conditional expectation and the conditional standard deviation of each variable in the MS-VAR as well as for the implied real interest rate (RIR), defined as the difference between the FFR and one-step-ahead inflation expectations. For each variable $z_t \in Z_t$, the conditional expectation and conditional standard deviation are given by $\mathbb{E}_t(z_{t+s})$ and $sd_t(z_{t+s}) = \sqrt{\mathbb{V}_t(z_{t+s})} = \sqrt{\mathbb{E}_t[z_{t+s} - \mathbb{E}_t(z_{t+s})]^2}$, where $\mathbb{E}_t(\cdot) \equiv \mathbb{E}(\cdot | \mathbb{I}_t)$ and $\mathbb{I}_t$ denotes the information available at time $t$. We assume that $\mathbb{I}_t$ includes knowledge of the regime in place at time $t$, the data up to time $t$, $Z^t$, and the VAR parameters for each regime. Both statistics are computed from the MS-VAR parameters and transition matrix $\boldsymbol{H}^A$, taking into account that future regimes are unknown and that there exists an entire posterior distribution of VAR parameters and transition matrix $\boldsymbol{H}^A$, translating into posterior distributions for $\mathbb{E}_t(z_{t+s})$ and $sd_t(z_{t+s})$. (Details on how these are calculated are presented in the Appendix.) Note that inflation expectations are computed using the MS-VAR estimates, therefore RIR is not included directly in the MS-VAR but derived ex-post based on the MS-VAR estimates. Note that $sd_t(z_{t+s})$ can be considered a measure of economic uncertainty, as implied by the MS-VAR.

Figure 5 reports the conditional expectations (from the full sample VAR) of each variable in the baseline MS-VAR plus the RIR. The figure reports the median and 68% credible sets from

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certainty about parameter transformations such as risk-premia.

[13] The beginning of the sample is three years later than the sample used to estimate *cay* because the federal funds rate data is only available starting in 1955:Q3.

the posterior distribution of the conditional expectations. The figure shows striking evidence of structural change in the expected long-run RIR that coincide with the regime sequence estimated for the mean of *cay*. The occurrence of the low asset valuation regime in the middle subsample from 1976:Q2-2001:Q2, coincides with an expectation of sharply higher values for the real federal funds rate, while the periods of high asset valuation at the beginning (1955:Q3-1976:Q1) and end (2001:Q3-2013:Q3) of our sample coincide with expectations of much lower real interest rates. The differences across subsamples are strongly statistically significant: the figure also reports the 68% posterior credible sets for the conditional expectations and shows that the sets for the two regimes never overlap. Note that, because the MS-VAR parameters are freely estimated, the estimation could have found no evidence of structural change in the expected real interest rate across these subsamples and/or that changes occur in variables other than the expected real interest rate. Figure 5 also shows that the estimated regime shifts in the expected future RIR show up prominently in the expectations for the real policy rate five to ten years ahead. This finding underscores the extent to which low frequency shifts in the mean of *cay* coincide with a persistent low or high interest rate environment, rather than transitory movements in these rates.

There is no clear pattern with inflation. Thus the breaks in the expected real interest rate five to ten years ahead appear mostly attributable to breaks in the conditional expected value of the *nominal* interest rate, which the Federal Reserve directly influences. Of course the Federal Reserve may also have considerable influence over expected inflation. But movements in expected inflation do not line up as well with the regime sequence for breaks in the mean of *cay* as do movements in the expected nominal interest rate: in the first subperiod, corresponding to the first instance of the high asset valuation regime, expected inflation was low and then high, while in the second subperiod, corresponding to the low asset valuation regime, inflation was high and then low, where it remained throughout the entire span of the third subperiod corresponding to the second instance of the high asset valuation regime. To the best of our knowledge, these findings provide among the first formal statistical evidence that low frequency shifts in asset values relative to economic fundamentals are strongly associated with persistent changes in value of a policy instrument under direct control of the central monetary authority.

Why are high asset valuations associated with low expected long-run policy rates, and vice versa? High wealth ratios could be associated with low expected long-run policy rates because the latter are expected to generate either faster long-run economic growth, or lower uncertainty about that growth. Conversely, regimes characterized by low wealth ratios and high expected policy rates could be explained by lower expectations for long-run growth and/or higher uncertainty about that growth. Figure 5, however, provides no evidence that the low frequency shifts to high asset valuation regimes are associated with higher expected economic growth, or vice versa; indeed the opposite is true. The high asset valuation subperiods at the

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beginning and end of our sample are associated with *lower* expected GDP growth five and 10 years ahead than is the low asset valuation subperiod in the middle of the sample. High asset valuation regimes are also marked by significantly lower expected R&D growth, and weaker investment growth. Figure 7 shows the analogous results for the secondary MS-VAR, which includes the change in the labor share and stock market dividend growth, in place of investment and R&D growth. High asset valuation regimes are associated with an expectation of persistent declines in the labor share, whereas the low asset valuation subperiod is characterized by the expectation of a stable labor share. Thus, the boom periods for asset values are associated with broad-based economic weakness and deteriorating payouts to workers. In an interesting exception to this pattern, there is some evidence that fundamentals for shareholders improve. Figure 7 shows that high asset valuation/low RIR regimes also marked by significantly higher expected dividend growth than the low asset valuation/high RIR regimes. These differences are statistically significant.

Figure 6 presents a different perspective on the interest rate regimes over the three asset valuation subperiods of our sample. The figure superimposes the 5- and 10-year-ahead expected real FFR implied by the baseline MS-VAR on the graph along with the actual (quarterly) real FFR, equal to the quarterly nominal FFR minus the one-step-ahead expected inflation rate implied by the VAR. The actual real quarterly FFR is more volatile than the VAR expected future funds rates. However, it is evident from this plot that the low-high-low pattern in the longer-run conditional expected values for the real FFR across the three subperiods is the result of persistent fluctuations in the quarterly federal funds rate around very different levels. Since the Federal Reserve can tightly control the real quarterly FFR, this shows that the lower frequency regimes in the 5- and 10-year-ahead conditional means are strongly attributable to highly persistent Federal Reserve interest rate policies, presumably aimed at achieving specific macroeconomic objectives. What is interesting is that, even though the shocks that motivated the persistently low interest rate policies in the early and late subperiods were likely quite different, the two subperiods are nonetheless characterized by similar outcomes for asset valuations and risk premia (as we show later). We discuss these subperiods as they relate to narratives about monetary policy below.

In order to gain further understanding of the role played by monetary policy in these interest rate regimes, we compute the long term response of the (nominal) FFR to a permanent change in inflation and output growth *conditional* on being in a certain regime. Following Primiceri (2005) and Sims and Zha (2006), we use this framework to characterize the stance of monetary policy in each regime, specifically the degree of activism in the systematic monetary policy responses to macroeconomic objectives. The goal is to compare these stances across regimes.

Consider the MS-VAR equation describing the behavior of the FFR. Equation (8) describes the reduced-form VAR, but here we focus on the implied structural-form. Zeroing in on the

terms relevant for the responses to inflation and output growth (setting all other terms to zero), we have the structural-form equation for the FFR:

$$
\begin{aligned}
FFR_t &= A_{0,FFR,\pi,\xi_t}\pi_t + A_{1,FFR,\pi,\xi_t}\pi_{t-1} + A_{2,FFR,\pi,\xi_t}\pi_{t-2} \\
&\quad + A_{0,FFR,\Delta GDP,\xi_t}\Delta GDP_t + A_{1,FFR,\Delta GDP,\xi_t}\Delta GDP_{t-1} + A_{2,FFR,\Delta GDP,\xi_t}\Delta GDP_{t-2} \\
&\quad + A_{1,FFR,\xi_t}FFR_{t-1} + A_{2,FFR,\xi_t}FFR_{t-2} + \omega_{FFR,t}
\end{aligned}
\tag{9}
$$

where the elements of the matrix $A_0$ capture the contemporaneous relations between the variables included in the VAR and $\omega_{FFR,t}$ is a structural monetary policy shock.[14] The expression (9) can be interpreted as a Taylor rule in which the FFR today depends on inflation, real activity, and past values of the FFR. This framework can be used to study the long run responses of the FFR to increases in inflation or output growth. Specifically, suppose that the inflation rate increases permanently by 1%. Then the long term response of the FFR under regime $\xi_t$ is given by:

$$
LR_{FFR,\pi,\xi_t} = \left[1 - \sum_{j=1}^{2} A_{j,FFR,\xi_t}\right]^{-1} \sum_{j=0}^{2} A_{j,FFR,\pi,\xi_t}.
$$

Similarly, if we are interested in the long run response of the FFR to a permanent 1% increase in output growth, we have:

$$
LR_{FFR,\Delta GDP,\xi_t} = \left[1 - \sum_{j=1}^{2} A_{j,FFR,\xi_t}\right]^{-1} \sum_{j=0}^{2} A_{j,FFR,\Delta GDP,\xi_t}.
$$

Table 6 reports the results for the long term responses of the FFR to inflation and output growth. The table reports the median and 68% posterior credible sets from the posterior distribution of the long term responses. The median values show that, no matter which regime, a permanent increase in inflation or GDP growth increases the FFR in the long run. But under the low valuation/high real interest rate regime, the long term response of the FFR to an increase in inflation is substantially larger than under the high valuation/low real interest rate regime. The opposite is true for the long term response of the FFR to output growth across the regimes. These differences are statistically significant. In other words, under the low valuation regime, the Federal Reserve seems to be more concerned with inflation stabilization, while under the high valuation regime it seems more concerned with output stabilization. These results provide indirect evidence that part of the reason for the breaks in the longer-term conditional mean value of the real interest rate across the previously documented subperiods involves shifts in the stance of monetary policy.

In theory, higher asset valuations could be the result of lower expected economic uncertainty (e.g., Lettau, Ludvigson, and Wachter (2008)). Figure 8 reports the conditional standard

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[14]We use a Cholesky identification scheme to pin down the contemporaneous effects of inflation on the FFR, under the assuption that the FFR can react contemporaneosuly to all other variables in the VAR, while the other variables react with a lag to movements in the FFR. This identification assumption is quite common in the structural VAR analysis.

of each variable in the MS-VAR plus RIR. The figure reports the median and 68% credible sets from the posterior distribution of the conditional standard deviations. The conditional standard deviation represents a statistical measure of uncertainty. The result in Figure 8 shows that macroeconomic uncertainty is *higher* rather than lower in subperiods of the high asset valuation regime as compared to the subperiod of the low asset valuation regime. This is true for uncertainty about GDP growth, inflation, investment growth, and R&D growth. Yet the opposite is true for the nominal and real federal funds rate. According to this evidence, infrequent shifts to high mean wealth ratios cannot be explained by lower macroeconomic uncertainty. The finding that macro and fed funds rate uncertainty vary inversely across the regimes is consistent with a more active role of the Federal Reserve in stabilizing inflation and real activity. As the Federal Reserve is expected to respond more aggressively by raising interest rates to counter higher inflation and/or a lower output gap, macroeconomic volatility is reduced, whereas the volatility of the FFR *can* increase. Still, this observation provides no support for the hypothesis that the high asset valuation regimes were the product of low economic uncertainty.

Table 7 reports, for the baseline MS-VAR, the means and standard deviations of the real interest rate, GDP growth, R&D growth, and investment growth, conditional on being in a particular regime $i$. Table 8 reports the same output for the secondary MS-VAR that investigates the change in the labor share and dividend growth. For each draw from the posterior distribution of the MS-VAR parameters, we compute means and standard deviations conditional on being in regime $i$ (see the Appendix for the precise calculation). This procedure gives an entire posterior distribution that we then summarize with the median and 68% posterior credible sets, reported in the Tables. We refer to these as *conditional steady state* values for the moments. These statistics corroborate the non-steady state evidence presented above where the possibility of a regime shift is incorporated into expectations: the two regimes present a clear difference for the mean and volatility of the real interest rate. The high asset valuation regime is characterized by sharply lower expected real policy rates and lower uncertainty about those rates, while the opposite is true for the low asset valuation regime. By contrast, Tables 7 and 8 taken together show that the high asset valuation regime is characterized by lower expected economic growth, lower expected investment growth, lower expected R&D growth, declines in the labor share, faster dividend growth, and higher uncertainty about all of these variables. In contrast to the high valuation subperiods, the low valuation subperiod is marked by increases in the labor share. Because the conditional steady states do not depend on the estimated transition matrix, they show that the main conclusions on the differences across regimes are robust to estimation error on the transition matrix.

Figure 9 gives a visual impression of the result. The figure plots the "wealth ratio" (the inverse of $cay^{MS}$ without removing the Markov-switching constant), along with the ten-year-

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ahead conditional expectation of the real federal funds rate implied by the baseline MS-VAR, on separate scales. The red dashed line in the figure shows the most likely value of the unconditional mean of the wealth ratio in each regime (given by the inverse of the regime-probability weighted average of $\alpha_1$ and $\alpha_2$). The mean shows clear regime shifts in wealth ratios that move from high to low to high over the sample, coinciding with a low then high then low expected long-run real federal funds rate. Regime shifts in the expected federal funds rate are large, ranging from about 1% in the low expected interest rate regimes to 3% in the high expected interest rate regime.

Some classic theories of rational bubbles suggest that higher policy rates can lead to *higher* asset values (e.g., Galí (2014)). But the evidence here is inconsistent with this story, since high wealth ratios are associated with low policy rates rather than high. An alternative explanation, consistent with the evidence here, is that any change in the expected short-term real interest rate will always have some effect on asset values because it changes the "fundamental" value of the asset. If prices are sticky and the Federal Reserve reduces the nominal interest rate, changes in monetary policy may reduce the rate at which investor's discount future payouts by reducing the real "risk-free" rate component of the discount rate, thereby increasing asset values. This effect would also be present in bubbles of the resale-option variant, since unlike the classic rational bubble, the resale-option bubble is proportional to fundamental value (e.g., Harrison and Kreps (1978); Scheinkman and Xiong (2003)). Regardless of whether a bubble of this form is present or not, asset valuations would be further increased if the risk premium component of the discount rate falls simultaneously with the risk-free rate because investors' willingness to tolerate risk is for some reason inversely related to the long-run expected value of the Federal Reserve's core policy instrument, consistent with a "reaching for yield" channel. We present tests of this hypothesis in the next section.

These findings capture three distinct periods of post-WWII US economic history. The first manifestation of regime 2 is in the subperiod from 1952:Q1-1976:Q1 and coincides with the run-up of inflation in the 1960s and 1970s, accommodative monetary policy, and low real interest rates. Economists have provided several possible explanations for why monetary policy failed to react aggressively to inflation during those years. However, they generally tend to agree that this was a period of high uncertainty and possibly passive monetary policy (Clarida, Gali, and Gertler (2000); Lubik and Schorfheide (2004); Sims and Zha (2006); Bianchi (2013)). The occurrence of the low asset valuation regime, in the middle subperiod from 1976:Q2-2001:Q2, precedes by three years Volcker's appointment as Chairman of the Federal Reserve and the disinflation that followed. The first attempts to bring inflation down started in the late 1970s, but Volcker succeeded only in the early 1980s, perhaps because of the political backing provided by the Reagan administration (Bianchi and Ilut (2015)). As a result, the beginning of Great Moderation is generally placed in the mid-1980s, when the economy experienced a

substantial reduction in volatility (McConnell and Perez-Quiros (2000); Stock and Watson (2002).) Macroeconomists interested in the Great Inflation tend to identify the change in the anti-inflationary stance of the Federal Reserve with the appointment of Volcker in August 1979. However, Sims and Zha (2006) estimate a structural MS-VAR and find a change in the conduct of monetary policy from less to more active toward the end of 1977, in line with the results here. Real interest rates increased significantly during the Volcker disinflation and they remained higher than in the 1970s for a prolonged period of time. In part this may have been attributable to a perceived need on the part of the Federal Reserve to rebuild credibility for low and stable inflation.

The second occurrence of the high asset valuation regime in the subperiod 2001:Q3-2013:Q3 starts with the end of the information technology (IT) boom and the beginning of the Federal Reserve's accommodative response to the recession that followed. Economists have identified the end of the Great Moderation with the 2008 recession, consistent with the estimated break patterns in Figure 8 for GDP growth uncertainty. At the same time, some have argued that monetary policy underwent a regime shift after the end of the IT boom (Campbell, Pflueger, and Viceira (2014)) and/or that interest rates were held "too low for too long" (Taylor (2007)) in response to the IT bust and the aftermath of 9/11. Asset values quickly recovered in 2002, and after a brief but dramatic decline in the financial crisis of 2007-2009, equity valuations resumed their upward march in 2009. This period of high equity valuations persists today with rates in the zero lower bound (ZLB) range coinciding with positive rates of inflation. Our estimates characterize this third subperiod as a return to a period of prolonged low real interest rates.

The three distinct *cay* regimes we estimate are remarkably close to the three distinct monetary policy regimes estimated by Campbell, Pflueger, and Viceira (2014), who use a completely different approach. Instead of identifying the break dates by using a cointegration relation in *cay*, they estimate break dates in the parameters of an estimated Taylor rule. Their first subperiod covers the period 1960:Q2-1977:Q1, the middle period is 1977:Q2-2000:Q4, and the last subperiod 2001:Q1 to the end of their sample 2011:Q4. They find that these regimes line up closely with shifts in estimated bond market betas. Although our focus is on regime shifts in an asset valuation ratio, *cay*, taken together, the results are suggestive of an important role for the Federal Reserve in driving persistent movements in equity and interest rate behavior.

## 5   Reaching for Yield?

We have found that low frequency swings in post-war asset valuation are strongly associated with low frequency shifts in the long-run expected value of interest rates, with low expected values for the real federal funds rate associated with high asset valuations, and vice versa. Moreover, while persistently low policy rates are associated with high asset valuations, this is not

because they signal strong economic growth, favorable changes in inflation, or low uncertainty. This suggests that persistent changes in monetary policy affect asset valuations because they change the rate at which investor's discount future payouts, in a manner that is independent of uncertainty about the aggregate economy. This could occur simply because the Central Bank influences the riskless real interest rate, a component of the discount rate. But the magnitude of this discount rate effect would be amplified if it went beyond the riskless rate to affect risk *premia*. If a switch from a high to low interest rate regime prompts investors to take on more risk, to "reach for yield," then the risk premium component of the discount rate would fall, further stimulating risky asset values relative to economic fundamentals. The reverse would occur in a shift from persistently low expected rates to high. We refer to this general idea as the *reaching for yield* hypothesis, or RFY for brevity.[15]

Observe that a change in discount rates driven by the risk-free rate alone influences all assets in the same way, regardless of their riskiness. By contrast, RFY implies that investors shift portfolio allocations toward riskier/higher return assets in low interest rate environments. Thus a change in discount rates accompanied by RFY will have effects that differ across assets, depending on the riskiness of the asset. As interest rates move from high to low, RFY implies a greater increase in the market value, relative to fundamentals, of higher return/higher Sharpe ratio assets than it does for lower return/lower Sharpe ratio assets. Equivalently, risk premia should fall more for riskier assets. We investigate this possibility here, using data on book equity relative to the market values of individual stock market portfolios that exhibit strong cross-sectional variation in return premia.

The book-market ratios of individual assets can also be differentially affected by a shift in interest rates because their expected future earnings are differentially affected. This has nothing to do with risk premia or discount rates, so it is important to correct for possible differences in expected earnings growth when assessing the RFY.

To do so, we carry out a log-linearization that follows Vuolteenaho (2000) and constructs earnings from book-market and return data using clean surplus accounting. Let $B_t$ denote book value and $M_t$ denote market value, and let the logarithm of the book-market ratio $\log(B_t/M_t)$ be denoted $\theta_t$. Vuolteenaho (2000) shows that the $\theta_t$ of an asset or portfolio can be decomposed as:

$$\theta_t = \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t r_{t+1+j} + \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t f_{t+1+j} - \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t e^*_{t+1+j} \tag{10}$$

where $\rho < 1$ is a parameter, and $r_{t+1+j}$, $f_{t+1+j}$, and $e^*_{t+1+j}$ stand for log excess return, log risk-free rate, and log earnings, respectively.[16] In other words, the logarithm of the book-market

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[15] In what follows we use the terms "risk" premia and return premia interchangeably to refer to the expected return on an asset in excess of the risk-free rate. We remain agnostic as to whether the observed premia are attributable to mispricing, to covariance with systematic risk factors, or both.

[16] Specifically, $e^*$ is the log of 1 plus the earnings-book ratio, adjusted for approximation error. See Vuolteenaho (2000).

ratio $\theta_t$ depends on the present discounted value (PDV) of expected excess returns (risk premia), expected risk-free rates, and expected earnings.

Given our objective to assess whether assets with different risk/return profiles respond differently to the regime changes identified above, we begin by focusing on the difference between the book-market ratios of portfolios known to have different risk/return profiles. Specifically, given two portfolios $x$ and $y$, the spread in their book-market ratios, $\theta_{x,t} - \theta_{y,t}$, is given by:

$$\underbrace{\theta_{x,t} - \theta_{y,t}}_{\text{Spread in BM ratios}} = \underbrace{\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( r_{x,t+1+j} - r_{y,t+1+j} \right)}_{\text{PDV of spread in risk premia}} - \underbrace{\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( e^*_{x,t+1+j} - e^*_{y,t+1+j} \right)}_{\text{PDV}_t \text{ of spread in expected earnings}}$$

Note that the risk-free rate has no affect on this spread, since all portfolios are affected in the same way by the risk-free rate. Instead only the risk premium differential and expected earnings differential affect the book-market spread. Since RFY pertains only to the return premium differential, we adjust the book-market spread for the spread in expected earnings to isolate the return premium differential:

$$\underbrace{\theta_{x,t} - \theta_{y,t} + \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( e^*_{x,t+1+j} - e^*_{y,t+1+j} \right)}_{\text{Spread in BM ratios adjusted for earnings}} = \underbrace{\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( r_{x,t+1+j} - r_{y,t+1+j} \right)}_{\text{PDV of the spread in risk premia}} \qquad (11)$$

The above expression shows that the spread in book-market ratios adjusted for expected future earnings is equal to the PDV of the spread in expected excess returns, or risk premia.

Denote the *adjusted* (for expected earnings) book-market ratio for portfolio $x$ in regime $i$ with a tilde as

$$\widetilde{\theta}^i_{x,t} \equiv \theta^i_{x,t} + \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t e^{*i}_{x,t+1+j}.$$

Let $x$ denote a high return premia portfolio while $y$ denotes a low return premia portfolio. Reaching for yield implies that, in a shift from a high ($i = 1$) to low ($i = 2$) interest rate regime, the adjusted book-market ratio of $x$ should fall more than that of $y$, implying $\left( \widetilde{\theta}^1_{x,t} - \widetilde{\theta}^2_{x,t} \right) - \left( \widetilde{\theta}^1_{y,t} - \widetilde{\theta}^2_{y,t} \right) > 0$, or that the *difference-in-difference* of adjusted book-market ratios should be positive across regimes:

$$\left( \widetilde{\theta}^1_{x,t} - \widetilde{\theta}^1_{y,t} \right) - \left( \widetilde{\theta}^2_{x,t} - \widetilde{\theta}^2_{y,t} \right) > 0. \qquad (12)$$

In summary, RFY implies that the *spread* in the *adjusted* book-market ratios between the high return/high risk portfolio and the low return/low risk portfolio should be greater in regime 1 than in regime 2.

To assess empirically whether the spread in adjusted book-market ratios between assets with different risk/return profiles is statistically different across the two regimes, we again estimate MS-VAR, now using portfolio data rather than macro data. The portfolio data are chosen to exhibit strong cross-sectional variation in average return premia and Sharpe ratios. The VAR specification takes the same form as equation (8), only the variables differ. Just as

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in the previous section, we impose the formerly estimated regime sequence for *cay* on the portfolio MS-VAR, but the parameters characterizing the different regimes, as well as the transition matrix, are freely estimated. The reasoning for doing so is the same as given above for the macro MS-VAR: we are interested in knowing whether the previously documented regime sequence for *cay* is characterized by evidence of RFY. This requires that we impose the previously estimated regime sequence, but since the MS-VAR parameters are freely estimated, the empirical procedure is free to find no evidence of structural change across these subsamples if indeed there is none.

We consider five different spread-portfolio VAR specifications using data on portfolios of stocks sorted by size (market capitalization) and book-market (BM) ratio, and portfolios of stocks sorted according to momentum (recent past return performance). The VARs differ according to which portfolios of a given size quintile are used to build the value book-market and return spreads. Specifically, the first portfolio VAR specification includes data on the five Fama-French risk factors $R_m - R_f$, $SMB$, $HML$, $RMW$, and $CMA$ (Fama and French (2015)), the real FFR computed as FFR-inflation, and the following four variables:

1. Value BM spread: The difference between the logarithm of the BM ratio of the small (size quintile 1) high book-market portfolio and the logarithm of the BM ratio of the small (size 1) low book-market portfolio.

2. Momentum BM spread: The difference between the logarithm of the BM ratio of the extreme winner (M10) portfolio and the logarithm of the BM ratio of the extreme loser (M1) portfolio.

3. Value return spread: The difference between the excess return of the small (size 1) high BM portfolio and the excess return of the small (size 1) low BM portfolio.

4. Momentum return spread: The difference between the excess return of the extreme winner (M10) portfolio and the excess return of the extreme loser (M1) portfolio.

The other four portfolio VARs are obtained by replacing the value BM spread and the value return spread by the corresponding series for the portfolio in a different size quintile. We denote these size quintiles S1, S2,...,S5, where S1 is the smallest quintile and S5 the largest. The five Fama/French factors are always included in the VAR because they contain predictive information for the return premia on these portfolios, and thus for the PDV of the spread in expected excess returns. The BM data are constructed starting from the 25 Fama/French portfolios sorted by size and BM, and the 10 portfolios sorted on momentum.[17] The data on BM ratios for individual portfolios are constructed from CRSP and Compustat exactly as the

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[17]http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html

Fama-French portfolio returns are constructed. We mimic the selection and breakpoints of this construction and compute the book-market ratio of each portfolio. The sample for this estimation is 1964:Q1-2013:Q4. This is shorter than the one previously used for $cay^{MS}$ and the macro VAR because reliable data for book-market ratios are not available prior to 1964:Q1.

It is well known that equity assets formed by sorting stocks into portfolios on the basis of book-market and size, and on the basis of recent past return or earnings performance, exhibit sizable cross-sectional variation in return premia and conventional measures of the risk/return trade-off. High BM portfolios earn much higher average returns than low BM portfolios, especially in the small size categories. Along the momentum dimension, recent past winner stocks earn much higher returns than recent past losers. Table 9 reports sample statistics in our data on the annualized Sharpe ratios and means for the long-short value and momentum strategies. Specifically, the table reports these statistics for a portfolio that is long in the extreme value portfolio (highest BM ratio) and short in the extreme growth portfolio (lowest BM ratio) of a given size category, and for a portfolio that is long in the extreme winner portfolio (M10) and short in the extreme loser portfolio (M1). The table also reports the same statistics for the individual portfolio returns in excess of the risk-free rate, where the risk-free rate for this purpose is the one used by Fama-French, denoted $R_f$, and equal to the one-month Treasury bill rate. It is evident that the value strategies that go long in the high BM portfolio and short the low BM portfolio have high Sharpe ratios and return premia, especially those in the three smallest size categories. The Sharpe ratio for the smallest value long-short strategy is 0.62 with a mean return of 10%, while that on the medium size quintile has a Sharpe ratio of 0.42 and mean return of 7%. The momentum strategy has an annualized Sharpe ratio of 0.64 and mean return of over 15%. Moreover, for the individual value, growth, winner, loser portfolios, value portfolios have much higher risk-premia than growth portfolios in the same size category, while the winner portfolio has a much higher risk premium than the loser portfolio. Indeed, the loser portfolio has a negative average return premium in the full sample, suggesting that it is not a risky asset and may even provide insurance. These statistics confirm the well known heterogeneity in the risk/return profiles of these portfolios.

Our objective is exploit this heterogeneity to isolate the effects of the previously estimated regime changes in the mean of $cay$ on the adjusted BM ratios of different portfolios. To do so, we begin by computing the *regime average* values of the adjusted BM *spreads* between the high and low return premia portfolios, $\overline{\theta}^i_{xy} \equiv \overline{\theta}^i_x - \overline{\theta}^i_y$, for each regime $i$. The regime average value of $\overline{\theta}^i_{xy}$ is defined to be the expected value of $\widetilde{\theta}_{xy,t}$, conditional on being in regime $i$ today and on the variables of the VAR being equal to their conditional steady state mean values for regime $i$. (The Appendix gives formal expressions for the regime average, and explains how they are computed from the MS-VAR parameters.) For each draw of the VAR parameters from the posterior

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distribution of these parameters, we compute the median and 68% credible sets for $\overline{\widetilde{\theta}}^i_{xy}$, which are reported in Table 10. The high $(x)$ and low $(y)$ return premia portfolios along the BM dimension are always the extreme value (highest BM) and the extreme growth portfolio (lowest BM), respectively, in each size category. Likewise, the high and low return premia portfolios along the momentum dimension are always the extreme winner (M10) and extreme loser portfolio (M1). The third row reports the analogous values for the regime average of the difference-in-difference of adjusted book-market ratios between the high and low return premia portfolios across the two regimes, i.e., the difference between the spreads $\left(\widetilde{\theta}^1_{x,t} - \widetilde{\theta}^1_{y,t}\right) - \left(\widetilde{\theta}^2_{x,t} - \widetilde{\theta}^2_{y,t}\right)$, as implied by the VAR estimates. To interpret the table, keep in mind that regime 1 is the low asset valuation/high interest rate regime, while regime 2 is the high asset valuation/low interest rate regime.

For all five size quintiles and every VAR, Table 10 shows that the adjusted BM spreads between the high and low return premia portfolios are positive in both regimes. This is not surprising because portfolios that have higher risk premia should have lower market values, holding fixed expected earnings and book value. Importantly, however, the table shows that these spreads are greater in regime 1 (low asset valuations/high interest rates) than in regime 2 (high asset valuations/low interest rates). Thus the difference-in-difference across regimes is always positive. This implies that the adjusted book-market ratios of high return/high risk portfolios fall more in a shift from high to low interest rate regime than do those of low return/low risk portfolios. Put differently, the return premia of evidently riskier/higher return assets decline more in environments with persistently high aggregate wealth ratios and low Federal Reserve policy rates than do less risky/lower return assets.

The third row also reports the 68% posterior credible intervals in parentheses for the difference-in-difference. The break in the BM spreads is proportionally smaller than the break in the momentum spreads. However, in all cases it is positive and, with the exception of the adjusted BM spread between the value and growth stocks in the fourth largest size quintile (size 4), this difference is strongly statistically significant. These results are supportive of a channel that implies an increased appetite for risk-taking in low interest rate environments.

The magnitude of the RFY channel for some portfolios is striking, and the case of momentum in particular deserves emphasis. The results indicate that, in every VAR estimated, the spread in adjusted BM ratios between the winner and loser portfolios is nearly one and a half times as high in the high interest rate regime than in the low interest rate regime, suggesting quantitatively large shifts toward greater risk-taking in the low interest rate subperiods. It is notable that momentum investing, with portfolios that exhibit the largest shift in these spreads that we find, is known to be among the most volatile equity investment strategies studied, one that is subject to infrequent but extreme crashes especially prevalent in times of crisis or panic (Daniel and Moskowitz (2013)).

These findings may be equivalently stated in terms of risk premia (see 11): long-short portfolios that exploit value and momentum spreads have lower risk premia in low interest rate regimes, and higher risk premia in high interest rate regimes. Specifically, the PDV of all future risk premia spreads is estimated to be considerably lower in low interest regimes than in high interest rate regimes.

The findings in Table 10 report the regime average values of the PDV of risk premia, conditional on one regime or another. We can also estimate the PDV of risk premia as it evolves over the sample, rather than an average value conditional on a regime. These values are estimated as the VAR forecasts, or conditional expected values, of the return premia $\mathbb{E}_t(\cdot) \equiv \mathbb{E}(\cdot | \mathbb{I}_t)$, where $\mathbb{I}_t$ again includes knowledge of the regime in place at time $t$, the data up to time $t$, $Z^t$, and the VAR parameters for each regime. Given the posterior distribution of the VAR parameters, these forecasts have a posterior distribution and we report as solid (blue) lines the median values of these forecasts in Figure 10. The regime averages are given by the dashed (red) lines. (The Appendix explains how this is computed from the MS-VAR estimates.) The first row of Figure 10 reports these median forecasts for the value-growth spreads in each size category and the winner-loser spread. (The results for the latter are arbitrarily reported using the estimates of the (S5) VAR, since the results are very similar across the five VAR specifications). Although the risk premia are volatile, there are some clear patterns. With the exception of the S4 value-growth portfolio, risk premia fluctuate around a lower point in the high asset valuation/low interest rate regime than they do in the low asset valuation/high interest rate regime. For the four long-short strategies other than this one, the estimated risk premia reach lows or near-lows in the post-millennial period, during the second occurrence of the high asset valuation/low interest rate regime, after shooting up briefly in the aftermath of the financial crisis of 2007-2008. The risk premia then return to low levels in the post-crisis ZLB period.

We also investigate how the risk premia on the individual value and momentum portfolios, as opposed to the long-short strategies (or spreads), have changed over the sample. To do so, we form an estimate of the first term on the right-hand-side of (10), namely the PDV of all future risk premia, for *each portfolio*. These estimates are formed from six MS-VAR specifications. The first five VARs pertain to portfolios sorted according to size and book-market, and differ by size quintile. The sixth VAR uses the same data but for the momentum portfolios. Specifically, each VAR includes data on the five Fama-French risk factors $R_m - R_f$, $SMB$, $HML$, $RMW$, and $CMA$ (Fama and French (2015)), the real FFR computed as FFR-inflation, and the following four variables:

1. The logarithm of the BM ratio of the high book-market portfolio (value) in a given size quintile (small, S2, S3, S4, large) or the logarithm of the BM ratio of the extreme winner

(M10) portfolio.

2. The logarithm of the BM ratio of the low book-market portfolio (growth) in a given size quintile (small, S2, S3, S4, Large) or the logarithm of the BM ratio of the extreme loser (M1) portfolio.

3. The excess return of the BM ratio of the high book-market portfolio (value) in a given size quintile (small, S2, S3, S4, large) or the excess return of the BM ratio of the extreme winner (M10) portfolio.

4. The excess return of the BM ratio of the low book-market portfolio (growth) in a given size quintile (small, S2, S3, S4, Large) or the excess return of the BM ratio of the extreme loser (M1) portfolio.

The specification of these six VARs differs from that of the five VARs used above to estimate adjusted BM spreads across portfolios, where the momentum variables were included in each size/book-market VAR. There are two reasons we use a different VAR specification for estimating the risk premia on the spreads than for estimating those for the individual portfolios. First, although the VAR specified above for the individual portfolios could in principal be used to estimate risk premia for the long-short spreads, in practice risk premia for the spreads are estimated far more precisely using a VAR that includes data on the spreads, rather than data on the individual portfolios separately. Second, including the momentum variables in the book-market VARs for the individual portfolios (as was done for the spreads) would have required two additional variables in each VAR compared to the case for the spreads, a number that overwhelmed the numerical calculation and caused many draws to be rejected due to non-stationarity. Stationarity is required to form accurate estimates of the conditional expectations embedded in risk premia and the PDV of expected returns. The above specification with six VARs exhibited no problems with stationarity.

The results on evolution of risk premia for the individual portfolios are displayed in the second and third rows of Figure 10. The solid (blue) line is the estimated PDV of risk premia over time, while the dashed (red) line is the regime average value for the PDV of risk premia. For nine of the twelve portfolios, the PDV of risk premia are lower in the high asset valuation/low interest rate regime than in the low asset valuation/high interest rate regime. For several portfolios these premia reach lows or near-lows in the post-2000 period and the post-crisis period, when in the latter case interest rates entered the ZLB range. Almost all portfolios exhibit an increase in risk premia during the years corresponding to the financial crisis, but in every case risk premia decline subsequent to the crisis. It is worth noting that the one big exception to this is the loser portfolio, which as noted has an average return over the risk-free rate that is *negative*, suggesting not only that it is not a risky asset, and may even provide

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insurance. If so, we would not expect the behavior of this portfolio to be even qualitatively consistent with those of the risky portfolios, since the predictions of either literature discussed in the introduction for how the risk premium on an asset changes with the interest rate are predicated on the presumption that the asset in question carries a positive risk premium on average.

We may use the estimated posterior distribution of the regime average value of risk premia to compute the exact probability that risk premia fall in low interest rate regimes. These probabilities are computed as the percentage of draws from the posterior distribution of regime averages for which risk premia are lower in regime 2 than in regime 1. The results are reported in Table 11. The first and second rows refer to the (regime average) risk premia for the long-short spread portfolios, while the third and fourth rows refer to the risk premia on individual portfolios. For five out of the six spread portfolios, the probability that premia go down in high asset valuation/low interest rate regimes is at least 85%. The exception is value-growth size 4 spread portfolio. But three of the value-growth portfolios assign probabilities in excess of 90%, and the winner-loser portfolios assign a probability of effectively 100%. For the individual portfolios, the estimation likewise assigns a large probability of a decline in risk premia in most cases. For 8 out of 12 portfolios this probability is at least 85%. Moreover, taking the results for all portfolios together, the estimated probability appears related to the riskiness of the portfolio in the expected way. The long-short (spread) portfolios are arguably the riskiest since they require leverage. For these portfolios the probability that risk premia fall in the high asset valuation/low interest rate regime is above 90% for the three value-growth strategies that have the highest average risk premia (those for S1, S2, and S3), and is often close to 100% for winner-loser strategies. Similarly, the high risk premia value and winner individual portfolios exhibit greater probabilities of a decline in premia than do the lower risk premia growth and loser portfolios. Finally, in line with the evidence presented above, these calculations show that the probability of a decline in risk premia for the loser portfolio is close to zero (5%), consistent with the evidence that this portfolio does not command a positive risk premium on average and so would not be expected to exhibit a decline in its premium in regimes where the appetite for risk-taking is on the rise.

In summary, the results from estimating portfolio VARs indicate that low interest rate regimes are associated with lower risk premia for the majority of portfolios that carry a positive average risk premium, while the riskier portfolios exhibit larger declines in premia over the course of a low interest rate regime than do less risky portfolios. The findings are supportive of reaching for yield theories. They present a challenge, however, for theories that explain persistently low interest rate environments with shocks that shift in the composition of wealth toward more risk averse or more pessimistic investors (e.g., Barro and Mollerus (2014); Caballero and Farhi (2014); Hall (2016)). In contrast to RFY, these theories imply that low interest rates

coincide with higher rather than lower risk premia. The findings here indicate that, not only are risk premia lower conditional on being in a low interest rate regime but, for most portfolios the estimated historical variation in these premia reaches lows or near-lows in the post-millennial period and again at the onset of the ZLB period, after a brief but sharp spike upward during the financial crisis.

# 6 Other Valuation Ratios

This section briefly comments on the behavior of other stock market valuation ratios over our sample. Many such ratios, e.g., price-dividend ratios for the aggregate stock market, or price-payout ratios, appear to differ from *cay* in that they exhibit trends. For example, price-dividend ratios drift up over the post-war period. The methods employed here are not well suited to explaining non-stationary trends, although we view this phenomenon as an interesting area for future research. For the purposes of this paper, we could address this crudely by simply removing a trend before analyzing the data. Instead of simply removing the lowest frequency components, however, we use a band-pass filter to remove both the highest and lowest frequency components of each series, examining frequencies that correspond to cycles between 10 and 50 years, or "medium-term" components. These frequencies roughly coincide with the persistence of our previously estimated regimes for $cay^{MS}$ and so form a natural basis for comparison with those results.

Figure 11 plots the medium-term components of five different stock market valuation ratios, overlaying the $cay^{MS}$ regime subperiods on the figure, with the low *cay*-valuation subperiod indicated in gray shading and the high valuation subperiods at the beginning and end of our sample indicated in white shading. The five valuation ratios plotted are a CRSP value-weighted stock price-dividend ratio on a portfolio that does not reinvest dividends, the Flow of Funds (FOF) price-payout ratio, the FOF price-dividend ratio, Shiller's price-earnings ratio[18], and the value-weighted price-dividend ratio for all firms in NYSE, NASDAQ, and AMEX from a COMPUSTAT/CRSP merge. The figure shows that the medium-term components of these five other valuation ratios exhibit similar high-low-high valuation patterns over the same subperiods that characterize the high-low-high $cay_t^{MS}$ valuation regimes. Moreover, the average values of these series in each regime differ noticeably, especially for the price-dividend and price-payout ratios. The results presented here are consistent with Bianchi, Ilut, and Schneider (2017), who also find evidence of low frequency fluctuations in the price-dividend ratio in an estimated business cycle model with endogenous financial asset supply and ambiguity-averse investors.

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[18]http://www.multpl.com/shiller-pe/

# 7  Conclusion

This paper presents evidence of infrequent shifts, or "breaks," in the mean of the consumption-wealth variable $cay_t$, an asset market valuation ratio driven by fluctuations in stock market wealth relative to economic fundamentals. These infrequent mean shifts generate low frequency fluctuations in asset values relative to fundamentals as measured by $cay$. A Markov-switching $cay_t$, denoted $cay_t^{MS}$, is estimated and shown to be less persistent and have superior forecasting power for excess stock market returns compared to the conventional estimate. Evidence from a Markov-Switching VAR shows that these low frequency swings in post-war asset valuation are strongly associated with low frequency swings in the long-run expected value of the Federal Reserve's primary policy interest rate, with low expected values for the real federal funds rate associated with high asset valuations, and vice versa. The findings suggest that the expectation of persistently low policy rates may be partly responsible for the high asset valuations of the last several years, and vice versa for the low asset valuation regime in the middle part of our post-war sample.

At the same time, we find no evidence that the estimated structural shifts to high asset valuation regimes and persistently low policy rates are associated with rational optimism about the future in the form of expectations for stronger long-run economic growth or lower uncertainty about that growth. Indeed, high valuation regimes, including the post-millennial period, are marked by expected economic weakness in GDP growth, investment growth, and R&D growth, along with sharp declines in the labor compensation share of GDP. The one exception to this evidence of declining expected prosperity in high valuation regimes is the stock market itself, where the fundamentals for shareholders appear to improve significantly in the form of higher dividend growth. From the perspective of most macroeconomic theories, the apparent divergence between the fundamentals that govern broad-based economic prosperity and those for the stock market is puzzling.

Finally, we present evidence using cross-sections of expected portfolio returns that the high valuation/low interest rate regimes are marked by significantly lower equity market risk premia, consistent with the hypothesis that investors' willingness to tolerate risk in equity markets rises when the long-run expected value of the real federal funds rate is low. The magnitude of this effect is especially pronounced for some of the most volatile equity investment strategies subject to infrequent but extreme crashes, such as those based on leveraged momentum investing. The theoretical literature on "reaching for yield" channels is still in its infancy, while this paper is an empirical study. More theoretical work is needed to understand how and why this channel may arise, and to elicit additional implications for asset markets and macroeconomic quantities.

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# Tables and Figures

|  | Mode | Mean | 5% | 95% |
|---|---|---|---|---|
| $\alpha_1$ | 0.9186 | 0.9153 | 0.8853 | 0.9460 |
| $\alpha_2$ | 0.8808 | 0.8767 | 0.8467 | 0.9077 |
| $\alpha_1 - \alpha_2$ | 0.0378 | 0.0385 | 0.0358 | 0.0413 |
| $\beta_a$ | 0.2606 | 0.2679 | 0.2505 | 0.2852 |
| $\beta_y$ | 0.6156 | 0.6071 | 0.5873 | 0.6270 |
| $\sigma$ | 0.0080 | 0.0087 | 0.0080 | 0.0094 |
| $H_{11}^{\alpha}$ | 0.9900 | 0.9901 | 0.9705 | 0.9995 |
| $H_{22}^{\alpha}$ | 0.9925 | 0.9923 | 0.9771 | 0.9996 |

**Table 1:** Posterior modes, means, and 90% error bands of the parameters of the Markov-switching cointegrating relation. Flat priors are used on all parameters of the model. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

**Parameter Estimates:** $cay^{FC}$

| $\alpha$ | $\gamma_a$ | $\gamma_y$ |
|---|---|---|
| 0.8706 | 0.1246 | 0.7815 |
| (0.0345) | (0.0150) | (0.0168) |

**Table 2:** Parameter estimates for the fixed coefficient cointegrating relation. Standard errors are in parantheses. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

**Cointegration Tests**

| | | Dickey–Fuller t-statistic | | | | Critical values | |
|---|---|---|---|---|---|---|---|
| | Persistence cay | Lag = 1 | Lag = 2 | Lag = 3 | Lag = 4 | 5% | 10% |
| MS | 0.8131 | -4.7609 | -4.4168 | -4.4586 | -4.7618 | -3.80 | -3.52 |
| FC | 0.9377 | -2.2911 | -2.1556 | -1.8894 | -1.6583 | -3.80 | -3.52 |

**Table 3:** The first column reports the first-order autoregressive coefficient obtained regressing $cay_t$ on its own lagged value and a constant. The next four columns report augmented Dickey-Fuller $t$-statistics $(\hat{\rho} - 1)/\hat{\sigma}_{\hat{\rho}}$, where $\hat{\rho}$ is the estimated value for the autoregressive coeffcient used to test the null hypothesis of no cointegration. This test is applied to estimates of the cointegrating residual, $cay_t$. We include up to four lags of the first difference of $cay_t$. The critical values for the test when applied to cointegrating residual are reported in the last two columns and are taken from Phillips and Ouliaris (1990). The results for $cay_t^{MS}$ do not account for sampling error in the estimated Markov-switching mean. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

## Long Horizon Forecasting Regressions: Stock Returns

$h$-period regression: $\sum_{i=1}^{h}(r_{t+i} - r_{f,t+i}) = k + \gamma \; z_t + \epsilon_{t,t+h}$

Horizon $h$ (in quarters)

| $z_t =$ | 1 | 4 | 8 | 12 | 16 |
|---|---|---|---|---|---|
| | | | Full sample | | |
| $cay^{FC}$ | **0.60** | **2.26** | **4.16** | **5.68** | **7.42** |
| | (2.00) | (2.21) | (2.47) | (2.73) | (3.71) |
| | [0.01] | [0.05] | [0.10] | [0.14] | [0.20] |
| $cay^{MSfilt}$ | **1.54** | **6.38** | **11.60** | **13.56** | **13.61** |
| | (4.07) | (5.22) | (6.53) | (6.03) | (6.18) |
| | [0.04] | [0.18] | [0.35] | [0.37] | [0.34] |
| $cay^{MS}$ | **1.49** | **6.83** | **11.88** | **13.79** | **13.78** |
| | (3.86) | (6.08) | (6.63) | (6.11) | (6.25) |
| | [0.04] | [0.21] | [0.36] | [0.38] | [0.34] |
| | | Sub-sample 1981Q1-2013Q3, recursive | | | |
| $cay^{FC}$ | 0.17 | 1.00 | 2.48 | 3.96 | 6.39 |
| | (0.48) | (0.83) | (1.04) | (1.18) | (1.82) |
| | [-0.01] | [0.00] | [0.03] | [0.06] | [0.11] |
| $cay^{FCrec}$ | 0.30 | 1.67 | **4.04** | **6.16** | **8.10** |
| | (0.97) | (1.65) | (2.29) | (2.79) | (4.17) |
| | [ 0.00] | [0.04] | [0.16] | [0.27] | [0.41] |
| $cay^{MSrec}$ | 0.41 | 2.13 | **6.01** | **8.65** | **10.33** |
| | (1.10) | (1.92) | (2.73) | (3.51) | (5.17) |
| | [ 0.00] | [0.04] | [0.21] | [0.31] | [0.37] |

**Table 4:** This tables reports the results from regressions of of $h$-period-ahead CRSP-VW returns in excess of a 3-month Treasury-bill rate, $r_{f,t}$, on the variable listed in the first column. $cay^{FC}$ is the fixed-coefficient consumption-wealth ratio; $cay^{MSfilt}$ denotes the Markov-switching version of $cay$ using filtered probabilities and $cay^{MS}$ denotes the benchmark Markov-switching $cay$ using smoothed probabilities. The bottom panel reports results from regressions using recursively estimated versions of $cay$, in which all parameters are estimated using data up to time $t$ rather than using the full sample. The models are first estimated on data from 1952Q1-1970Q1. We then recursively add observations and reestimate the $cay$ variables over expanding sub-samples using data only up to the end of that subsample, continuing in this way until the end of the sample, 2013:Q3. Results are reported for the subsample since 1980. $cay^{FCrec}$ denotes the fixed coefficient $cay$ estimated recursively, while $cay^{MSrec}$ denotes the Markov-switching $cay$ estimated recursively using smoothed probabilities. For each regression, the table reports OLS estimates of the regressors, Newey-West (1987) corrected $t$-statistics (in parentheses), and adjusted $R^2$ statistics in square brackets. Significant coefficients based on a $t$-test at the 5% significance level are highlighted in bold face. The full sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

**Out-Of-Sample Forecasts**

$h$-period regression: $\sum_{i=1}^{h}(r_{t+i} - r_{f,t+i}) = k + \gamma \; z_t + \epsilon_{t,t+h}$

Horizon $h$ (in quarters)

| $z_t =$ | 1 | 4 | 8 | 12 | 16 |
|---|---|---|---|---|---|
| | | | Mean-squared errors | | |
| const | 0.75 | 3.08 | 5.48 | 7.92 | 9.73 |
| $r - r_f$ | 0.71 | 2.99 | 5.32 | 7.67 | 9.36 |
| $cay^{FC}$ | 0.71 | 2.90 | 4.67 | 6.74 | 7.36 |
| $cay^{MSfilt}$ | 0.70 | 2.47 | 2.64 | 3.01 | 3.72 |
| $cay^{MS}$ | 0.70 | 2.35 | 2.53 | 2.92 | 3.68 |
| $cay^{FCrec}$ | 0.72 | 2.87 | 4.38 | 5.72 | 6.61 |
| $cay^{MSrec}$ | 0.71 | 2.86 | 4.49 | 5.75 | 6.14 |

**Table 5:** This tables reports the mean-squared forecast errors from out-of-sample $h$-period-ahead forecasts of CRSP-VW returns in excess of a 3-month Treasury-bill rate using 60-quarter rolling subsamples. The single predictor variable in each regression is listed in the first column. The forecasting regression is first estimated on data from 1952Q1-1980Q1, and forecasts are made over the next $h$ periods. We then repeat this forecasting regression using data from the next 60 quarters of the sample, continuing in this way until the end of the sample, 2013:Q3. Mean-square-errors are reported for the subsample since 1980. $cay^{FC}$ is the fixed-coefficient consumption-wealth ratio, $cay^{MSfilt}$ and $cay^{MS}$ are the Markov-switching $cay$ variables using filtered and smoothed probabilities, respectively, $cay^{FCrec}$ is the recursively estimated $cay$ with fixed coefficients, and $cay^{MSrec}$ is the recursively estimated Markov-switching $cay$. The recursive estimates use data only up to time $t$. The full sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

**Long term responses of the FFR**

|  | Low Val. regime | High Val. regime | Difference |
|---|---|---|---|
| Inflation | 1.9994<br>(1.8050,2.2733) | 1.1573<br>(1.1223,1.1878) | 0.8458<br>(0.6455,1.1226) |
| Output growth | 1.0364<br>(0.5994,1.4451) | 1.6546<br>(1.4653,1.8383) | −0.6408<br>(−1.1055,−0.1494) |

**Table 6:** Long-term responses to a permanent 1% increase in inflation or GDP growth. This table reports the median and (in parentheses) the 68% posterior credible sets of the long term response of the FFR to inflation and GDP growth conditional on being in a certain regime. The last column reports the distribution for the difference in the long term responses between the low and high asset valuation regimes. The sample spans the period 1955:Q3-2013:Q3.

**Summary statistics for macroeconomic variables**

| | Real Interest Rate | | GDP growth | |
|---|---|---|---|---|
| | Low Val. | High Val. | Low Val. | High Val. |
| Conditional Mean | 3.6786 <br> (3.6422,3.7146) | 0.5767 <br> (0.5240,0.6551) | 2.3059 <br> (2.2521,2.3645) | 1.4325 <br> (1.3780,1.5070) |
| Conditional St. Dev. | 2.1339 <br> (2.0147,2.2716) | 1.6175 <br> (1.5390,1.7138) | 2.0653 <br> (1.9518,2.2170) | 2.6418 <br> (2.4873,2.8313) |
| | R&D growth | | Investment growth | |
| | Low Val. | High Val. | Low Val. | High Val. |
| Conditional Mean | 5.5673 <br> (5.5006,5.6363) | 3.0308 <br> (2.9336,3.1276) | 2.3107 <br> (2.1342,2.5069) | 0.9955 <br> (0.7280,1.3785) |
| Conditional St. Dev. | 3.6599 <br> (3.4127,3.9542) | 4.5715 <br> (4.3236,4.8426) | 7.8355 <br> (7.3873,8.4243) | 10.3755 <br> (9.7330,11.2078) |

**Table 7:** Conditional Steady States. This table reports the median and (in parentheses) 68% posterior credible sets of the conditional mean and standard deviation for the real interest rate, GDP growth, R&D growth, and investment growth based on the VAR estimates conditional on staying in each regime.The sample spans the period 1955:Q3-2013:Q3.

**Summary statistics for dividend growth and change in labor share**

| | Dividend growth | | Change in labor share | |
|---|---|---|---|---|
| | Low Val. | High Val. | Low Val. | High Val. |
| Conditional Mean | 1.5944 (1.2009,1.9428) | 2.8355 (2.2177,3.4550) | 0.0335 (0.0236,0.0437) | −0.1562 (−0.1599,−0.1527) |
| Conditional St. Dev. | 6.1340 (5.5882,7.0407) | 11.1681 (10.3621,12.5204) | 0.4944 (0.4514,0.5504) | 0.5381 (0.5116,0.5702) |
| | Real Interest Rate | | GDP growth | |
| | Low Val. | High Val. | Low Val. | High Val. |
| Conditional Mean | 3.4906 (3.3796,3.5864) | 0.5417 (0.4951,0.5915) | 2.4061 (2.3419,2.4715) | 1.3812 (1.3536,1.4063) |
| Conditional St. Dev. | 2.0918 (1.9586,2.2457) | 1.5957 (1.4995,1.7175) | 2.0614 (1.9414,2.2112) | 2.5955 (2.4418,2.7675) |

**Table 8:** Conditional Steady States. This table reports the median and (in parentheses) 68% posterior credible sets of the conditional mean and standard deviation for dividend growth, change in the labor share, the real interest rate, and GDP growth based on the VAR estimates conditional on staying in each regime. The sample spans the period 1955:Q3-2013:Q3.

Annualized Sharpe Ratios and Mean Returns

| Portfolio | SR | Mean | Portfolio | SR | Mean | Portfolio | SR | Mean |
|---|---|---|---|---|---|---|---|---|
| V-G (S1) | 0.6225 | 0.1047 | V-RF (S1) | 0.5271 | 0.1346 | G-RF (S1) | 0.0909 | 0.0299 |
| V-G (S2) | 0.3807 | 0.0637 | V-RF (S2) | 0.5072 | 0.1202 | G-RF (S2) | 0.1958 | 0.0565 |
| V-G (S3) | 0.4025 | 0.0671 | V-RF (S3) | 0.5786 | 0.1260 | G-RF (S3) | 0.2255 | 0.0589 |
| V-G (S4) | 0.1681 | 0.0271 | V-RF (S4) | 0.4312 | 0.0974 | G-RF (S4) | 0.2973 | 0.0703 |
| V-G (S5) | 0.1899 | 0.0285 | V-RF (S5) | 0.3978 | 0.0785 | G-RF (S5) | 0.2733 | 0.0500 |
| W-L | 0.6446 | 0.1588 | W-RF | 0.5689 | 0.1315 | L-RF | −0.0859 | −0.0273 |

**Table 9:** The table reports annualized Sharpe ratios, "SR," and mean returns, "Mean," for different portfolios. The Sharpe ratio is defined to be the unconditional mean return divided by the standard deviation of the portfolio return. The long-short portfolios "V-G" are the value-growth portfolios in a given size quintile, S1=smallest, S5=largest. long-short portfolios "W-L" are the winner-loser portfolio. For each size category, the return of the V-G portfolio portfolio return is the difference between the return on the extreme value (highest BM ratio) and the return of the extreme growth portfolio (lowest BM ratio). The return of the W-L portfolio return is the difference in returns between the extreme winner (M10) and the extreme loser (M1). The rows denoted "V-RF", "G-RF", "W-RF" and "L-RF" report the same statistics for the value, growth, winner and loser portfolios, respectively, in excess of the risk-free rate. All returns are computed at quarterly frequencies but the Sharpe ratios and mean returns are reported in annualized units. The sample spans the period 1964:Q1-2013:Q3.

## Breaks in Book-Market Ratio Spreads

| | Size 1 (Small) | | Size 2 | | Size 3 | |
|---|---|---|---|---|---|---|
| | Val-Gr | W-L | Val-Gr | W-L | Val-Gr | W-L |
| Regime 1 | 2.5402 (2.4791,2.6127) | 4.4258 (3.8221,4.8132) | 1.5740 (1.4708,1.6530) | 4.4525 (3.8418,4.8380) | 1.6086 (1.5560,1.6491) | 4.3910 (3.8142,4.7666) |
| Regime 2 | 2.3591 (2.2143,2.4869) | 2.9713 (2.2002,3.7477) | 1.1795 (0.9913,1.3387) | 2.9560 (2.2761,3.7596) | 1.4296 (1.3055,1.5265) | 2.9494 (2.1859,3.7089) |
| Diff-in-Diff | 0.1832 (0.0489,0.3441) | 1.3134 (0.8419,1.9107) | 0.3817 (0.2537,0.5411) | 1.2738 (0.8419,1.9323) | 0.1849 (0.0809,0.2976) | 1.2922 (0.8253,1.9347) |
| | Size 4 | | Size 5 (Large) | | | |
| | Val-Gr | W-L | Val-Gr | W-L | | |
| Regime 1 | 0.6981 (0.6520,0.7559) | 4.43157 (3.8702,4.799) | 0.7617 (0.6450,0.8566) | 4.4110 (3.7568,4.8146) | | |
| Regime 2 | 0.6768 (0.5581,0.8359) | 3.0214 (2.2810,3.7489) | 0.6217 (0.4501,0.7692) | 2.8464 (2.0979,3.6757) | | |
| Diff-in-Diff | 0.0174 ($-$0.1188,0.1412) | 1.2572 (0.8220,1.8571) | 0.1281 (0.0141,0.2822) | 1.3442 (0.8616,2.0611) | | |

**Table 10:** The first two rows report the conditional steady states for the spread in adjusted (for expected earnings) book-market ratios between the high and low return premia porfolios in each regime. The columns labeled "Val-Gr" report the spreads for portfolios sorted along the book-market dimension, in a given size category (extreme value minus extreme growth). The columns labeled "W-L" report the spreads for portfolios sorted along the recent past return performance dimension (extreme winner minus extreme loser). The row labeled "Diff-in-Diff" reports the difference between these spreads across the two wealth ratio/interest rate regimes. The numbers in each cell are the median values of the statistic from the posterior distribution while in parentheses we report 68% posterior credible sets. The sample spans the period 1964:Q1-2013:Q3.

**Probability of a Decline in Premia**

| | Size 1 (Small) | | Size 2 | | Size 3 | |
|---|---|---|---|---|---|---|
| Spread | V-G | W-L | V-G | W-L | V-G | W-L |
| % draws | 91.25 | 100 | 99.85 | 100 | 95.20 | 100 |

| | Size 4 | | Size 5 (Large) | | | |
|---|---|---|---|---|---|---|
| Spread | V-G | W-L | V-G | W-L | | |
| % draws | 54.85 | 99.95 | 86.70 | 100 | | |

| Portfolio | V (S1) | V (S2) | V (S3) | V (S4) | V (S5) | Winner |
|---|---|---|---|---|---|---|
| % draws | 99.90 | 99.85 | 99.00 | 94.50 | 95.00 | 77.60 |

| Portfolio | G (S1) | G (S2) | G (S3) | G (S4) | G (S5) | Loser |
|---|---|---|---|---|---|---|
| % draws | 94.20 | 6.65 | 49.80 | 88.60 | 88.90 | 0.05 |

**Table 11:** The table reports the probability of a reduction in the estimated risk premia when moving from the low market valuation regime 1 to the high market valuation regime 2. These probabilities are computed as the percentage of draws from the posterior distribution for which the return premia in regime 1 than in regime 2. The first and second rows refer to the spread portfolios, whereas the third and fourth rows refer to the individual portfolios.
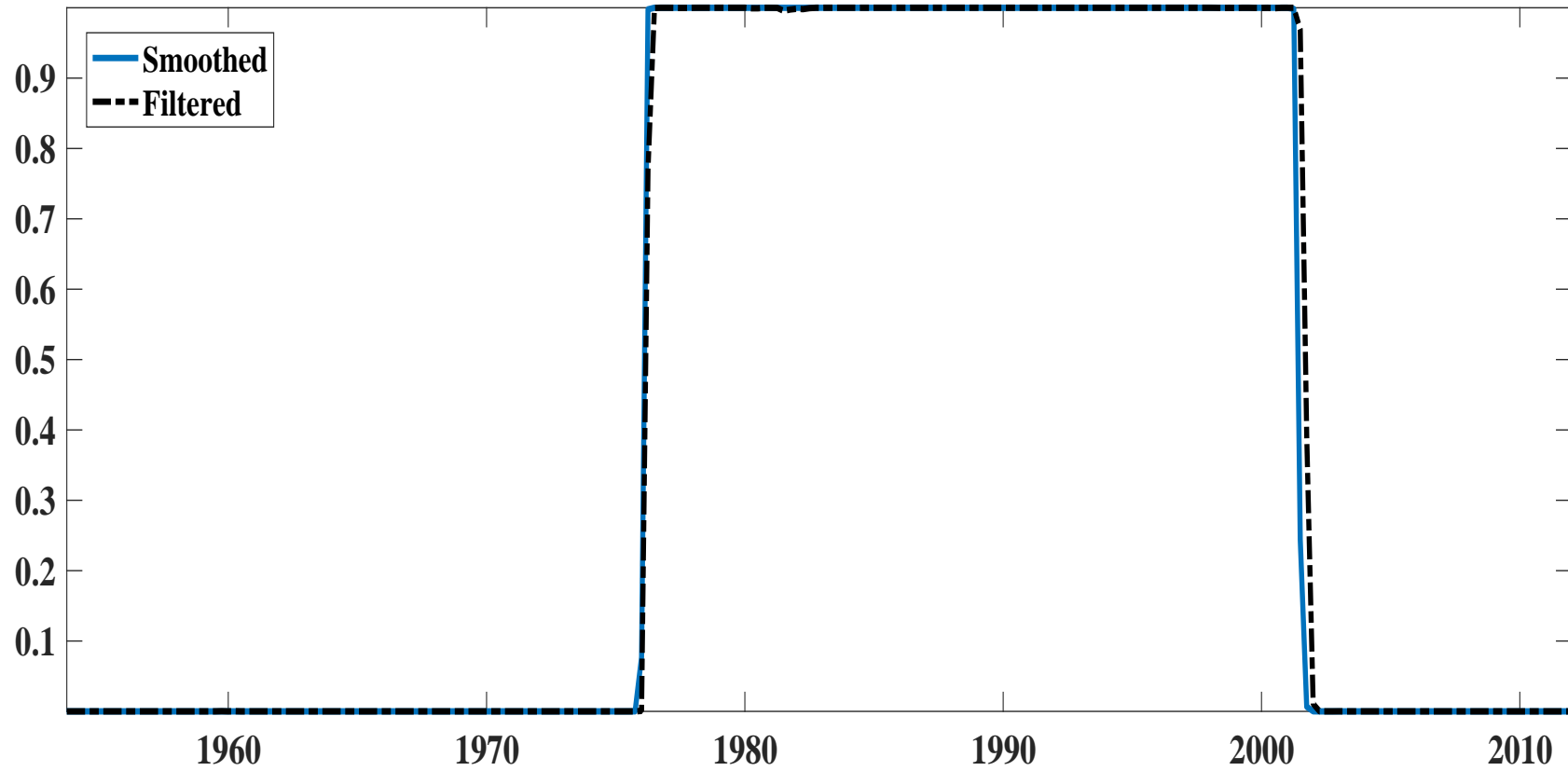
**Figure 1:** Smoothed probability of a low asset valuation regime for the Markov-switching cointegrating relation. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

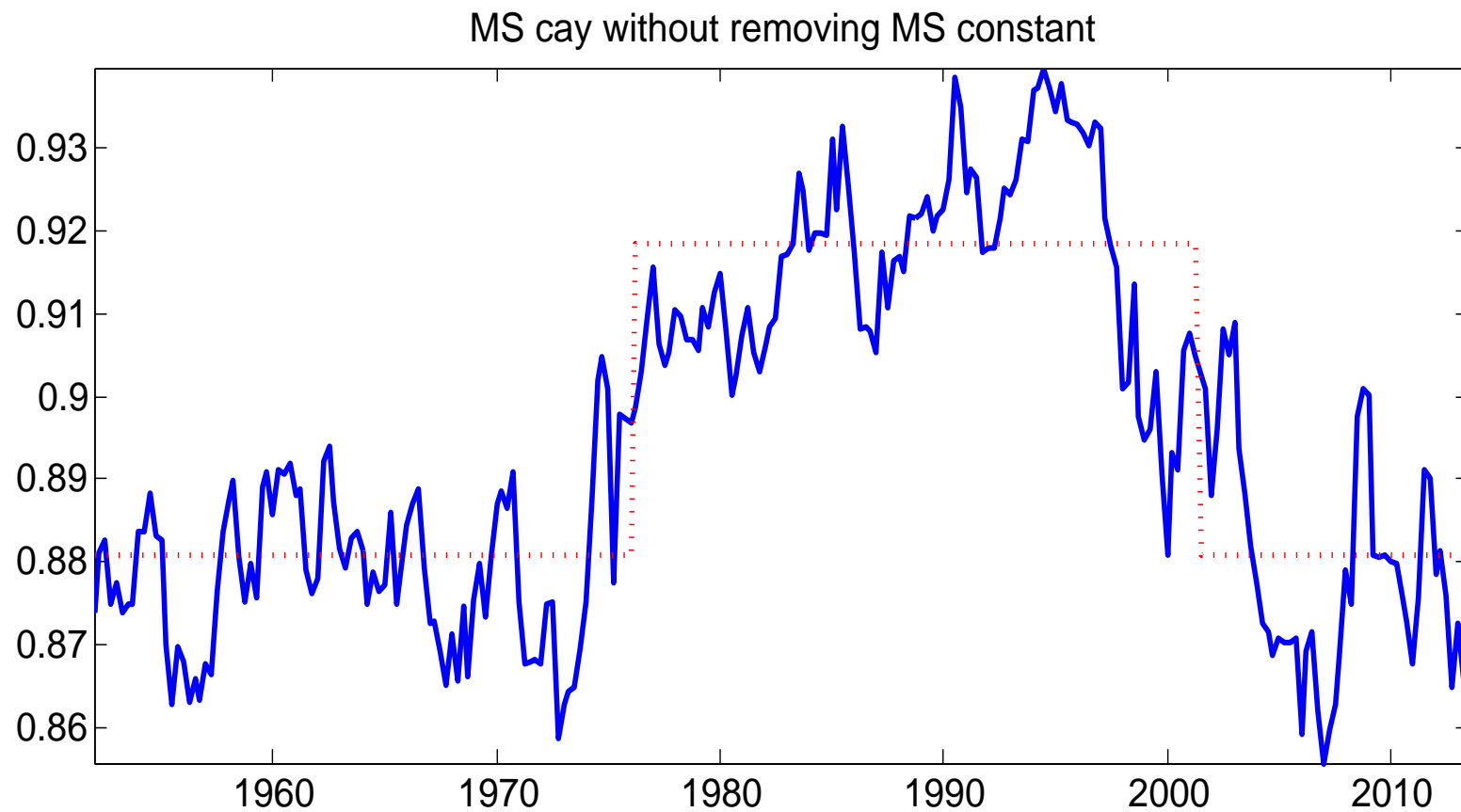**Figure 2:** The Markov-switching estimated $cay^{MS}$ is plotted with*out* removing the constant. The red dashed lines are the values of $\alpha_1$ and $\alpha_2$, which correspond to the most likely mean values in each regime. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.
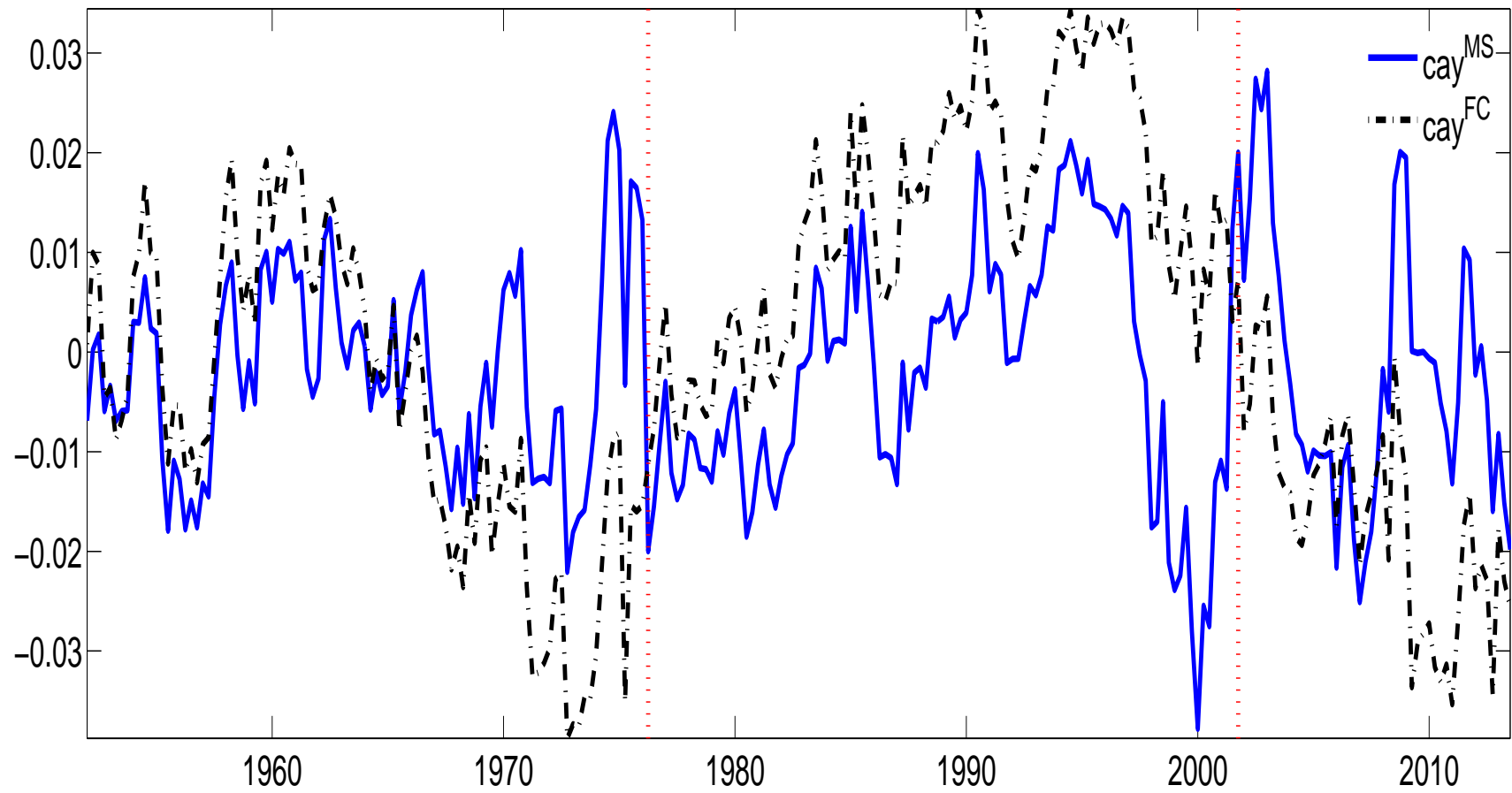
**Figure 3:** Markov-switching and fixed coefficients *cay*. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.
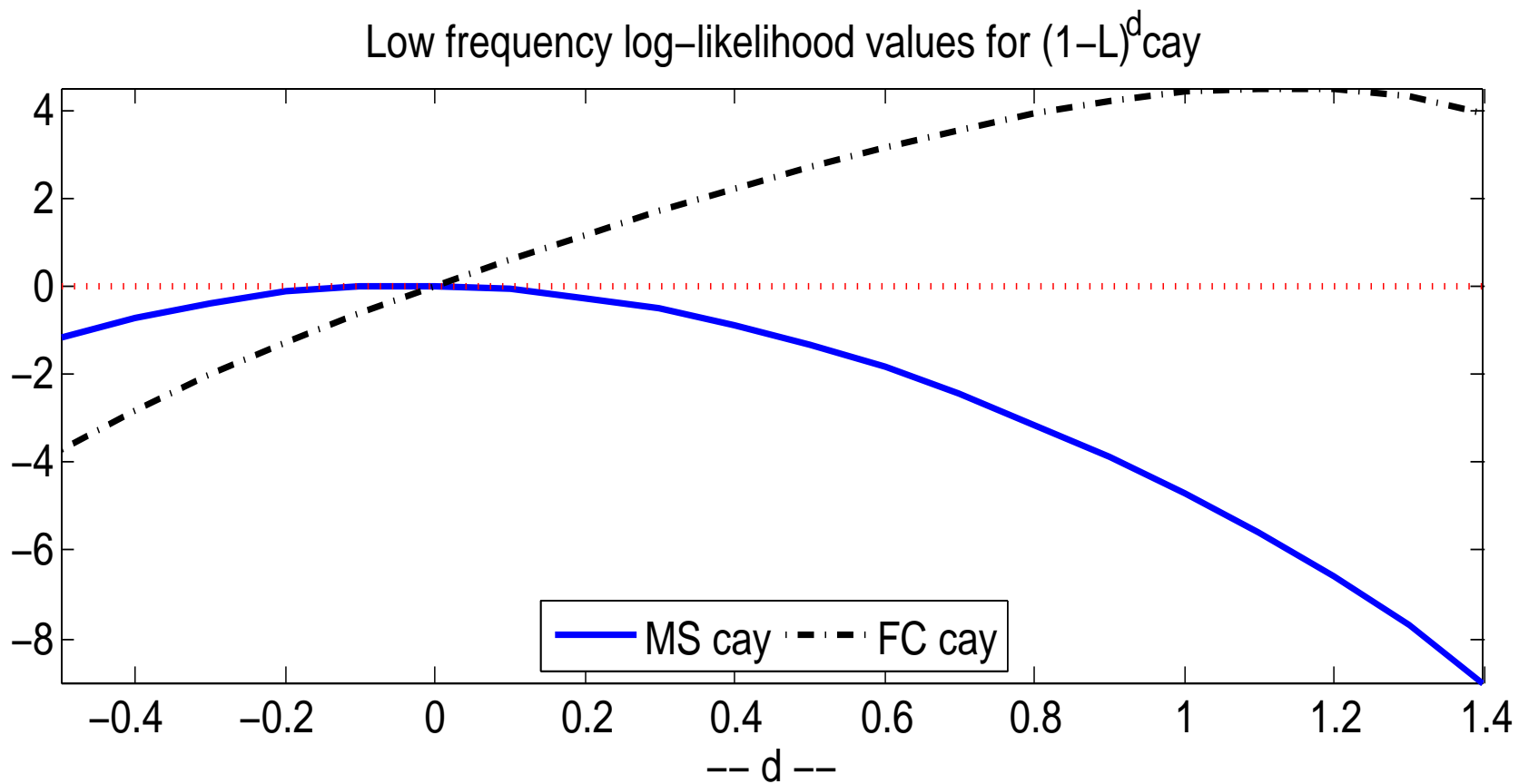
**Figure 4:** Low frequency log likelihood values for $(1 - L)^d cay_t$. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.
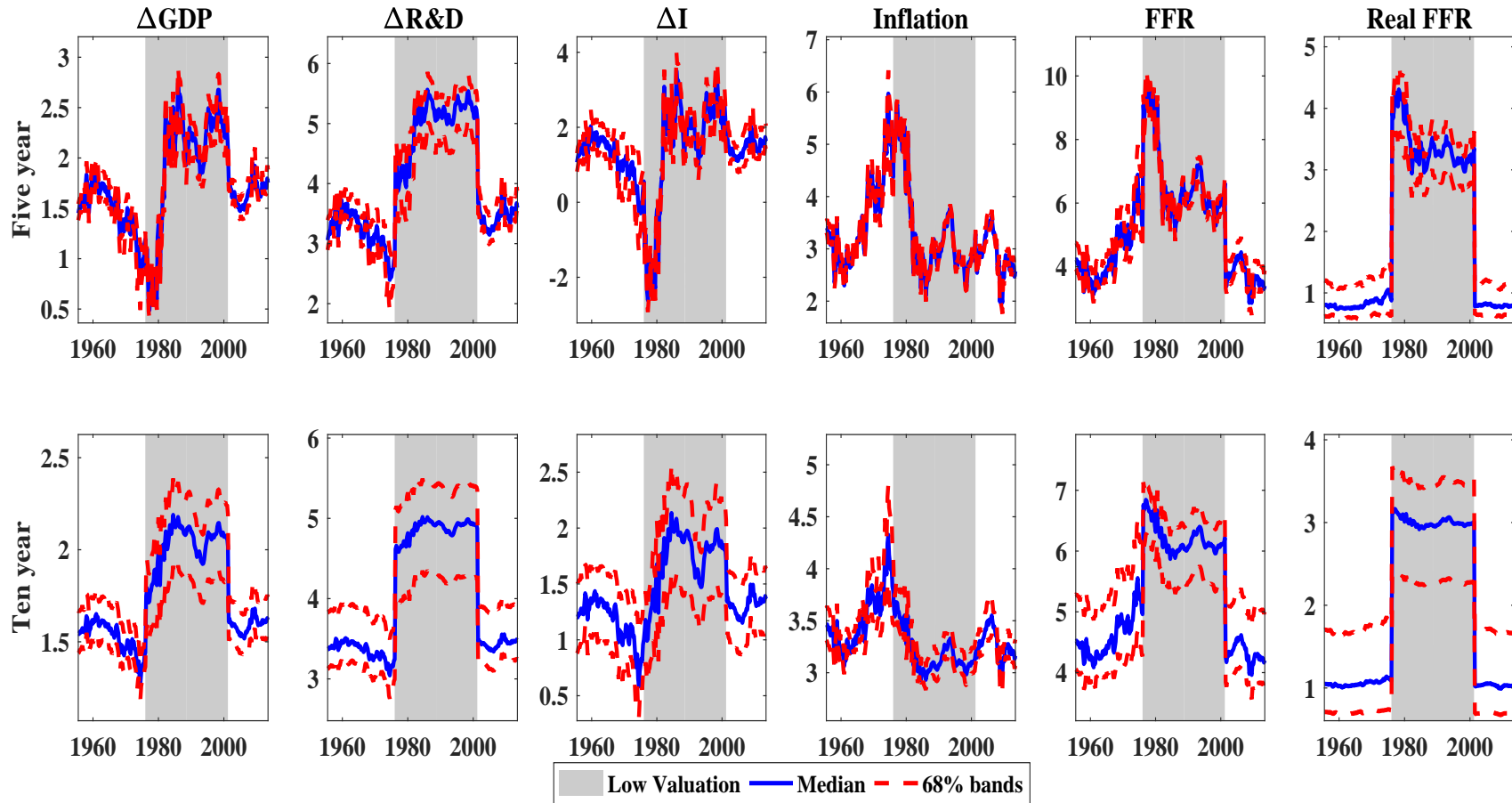
**Figure 5: Conditional expectations from baseline MS-VAR.** The figure reports the conditional expectations based on the baseline MS-VAR at different horizons taking into account the possibility of regime changes. The sample spans 1955:Q3 to 2013:Q3.
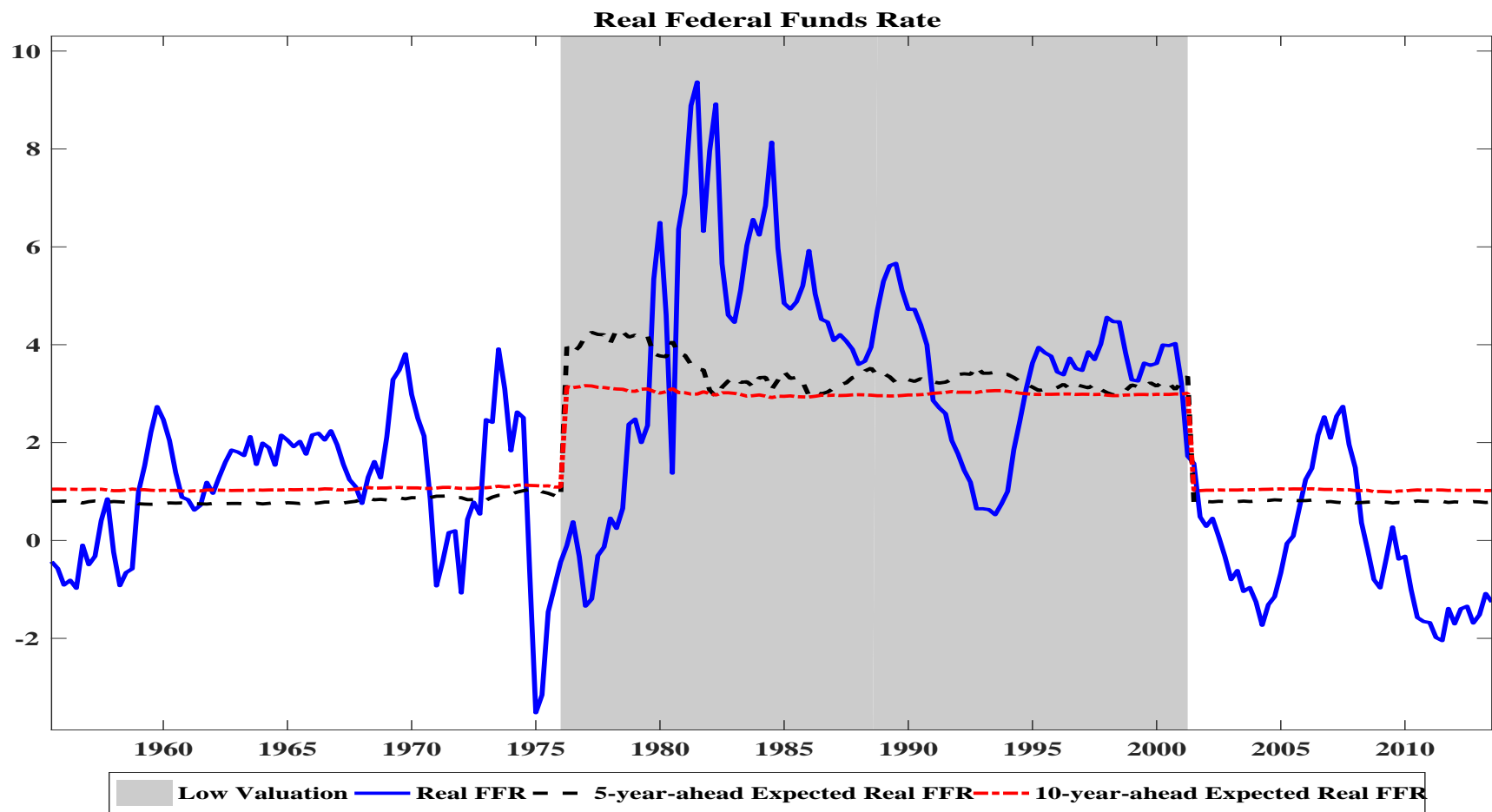
**Figure 6: Real Federal Funds Rate.** The figure reports the evolution of the Real FFR over time together with the 5-year-ahead and 10-year-ahead expectations as implied by the baseline MS-VAR. Expectations are computed taking into account the possibility of regime changes. The sample spans 1955:Q3 to 2013:Q3.
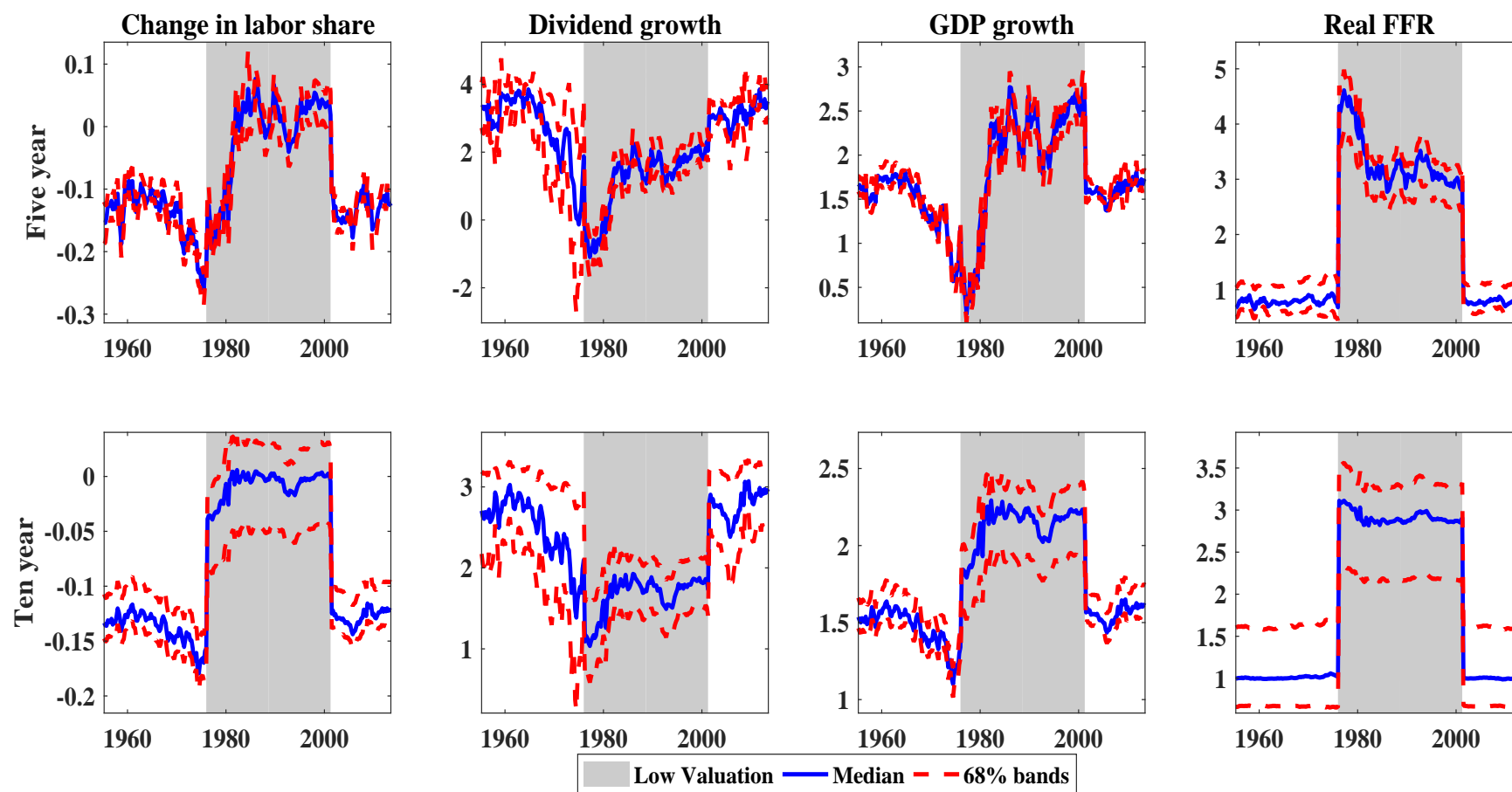
**Figure 7: Dividend growth and labor share.** The figure reports the conditional expectations for dividend growth, change in the labor share, GDP growth, and the real FFR. The expectations are based on a MS-VAR taking into account the possibility of regime changes. The sample spans 1955:Q3-2013:Q3.
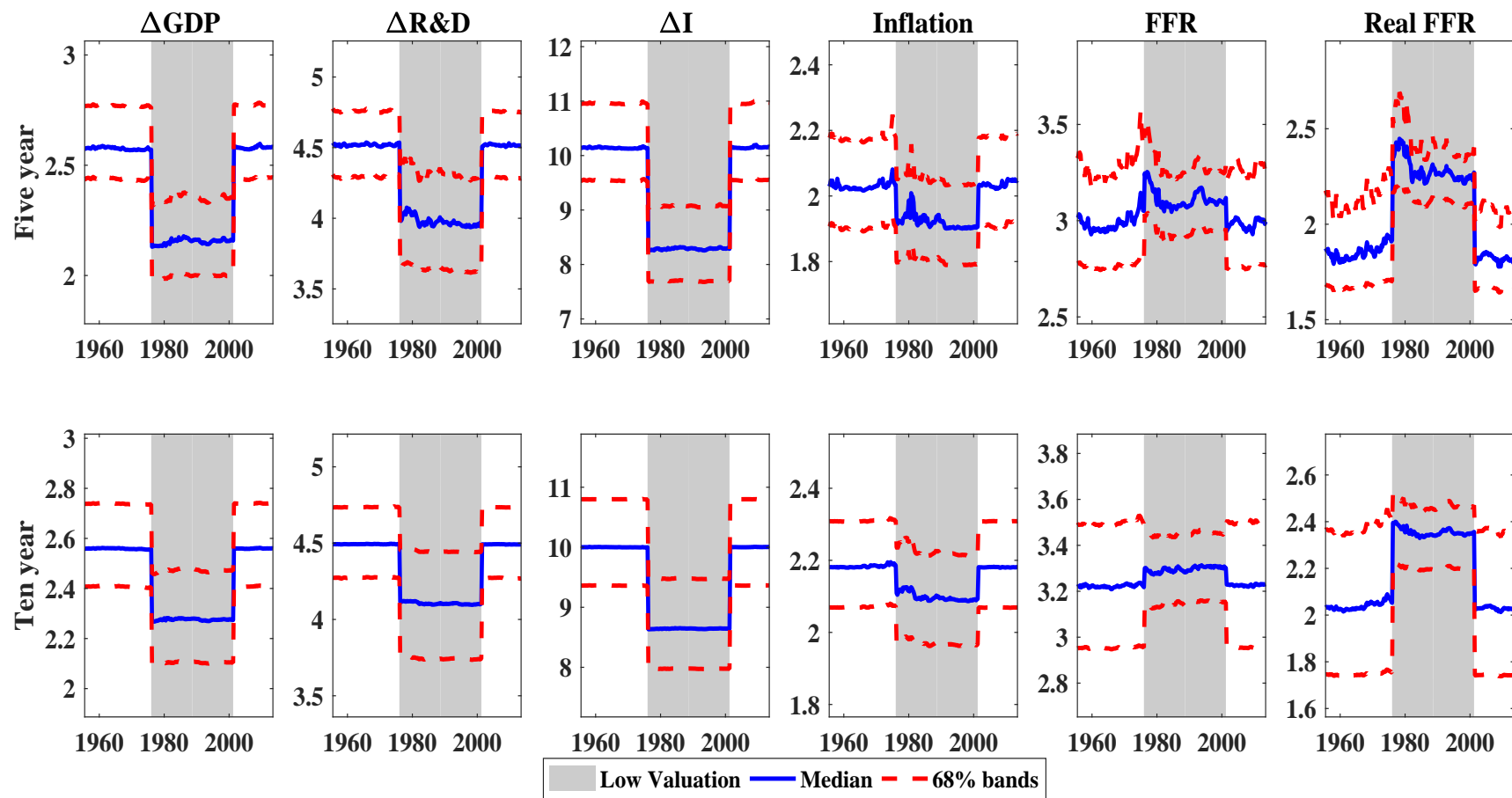
**Figure 8: Uncertainty based on MS-VAR.** The figure reports the conditional standard deviations at different horizons based on the MS-VAR taking into account the possibility of regime changes. The sample spans 1955:Q3-2013:Q3.
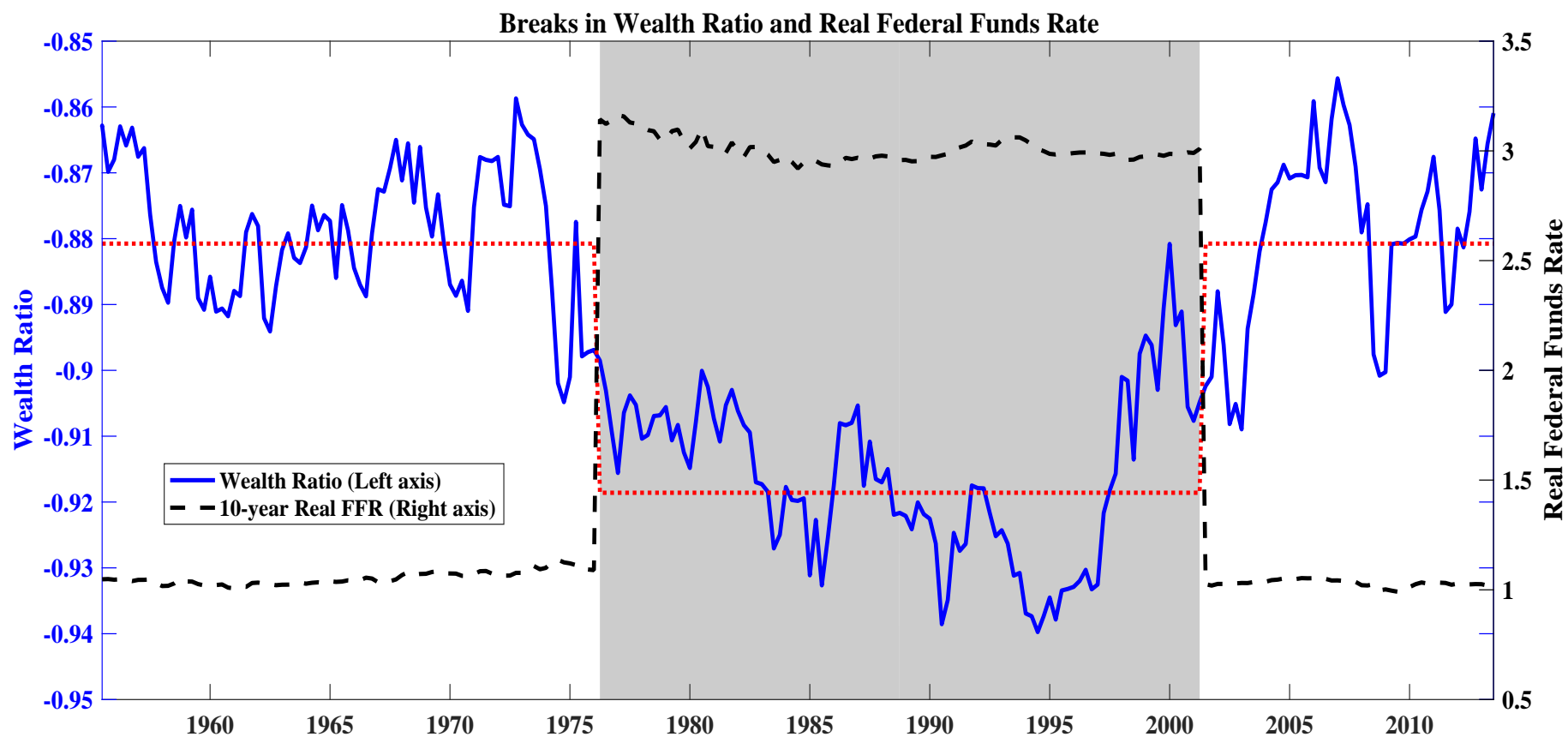
**Figure 9:** Wealth Ratio and federal funds rate. The wealth ratio (solid blue line, left axis) is plotted together with the conditional expectation of the ten-year-ahead real federal funds rate from the baseline MS-VAR (black dashed line, right axis). The wealth ratio is defined as the log inverse of $cay^{MS}$ without removing the Markov-switching constant. The red dashed line represents the log inverse of the regime-probability weighted average of the constants $\alpha_1$ and $\alpha_2$. The sample is quarterly and spans the period 1955:Q4 to 2013:Q3.
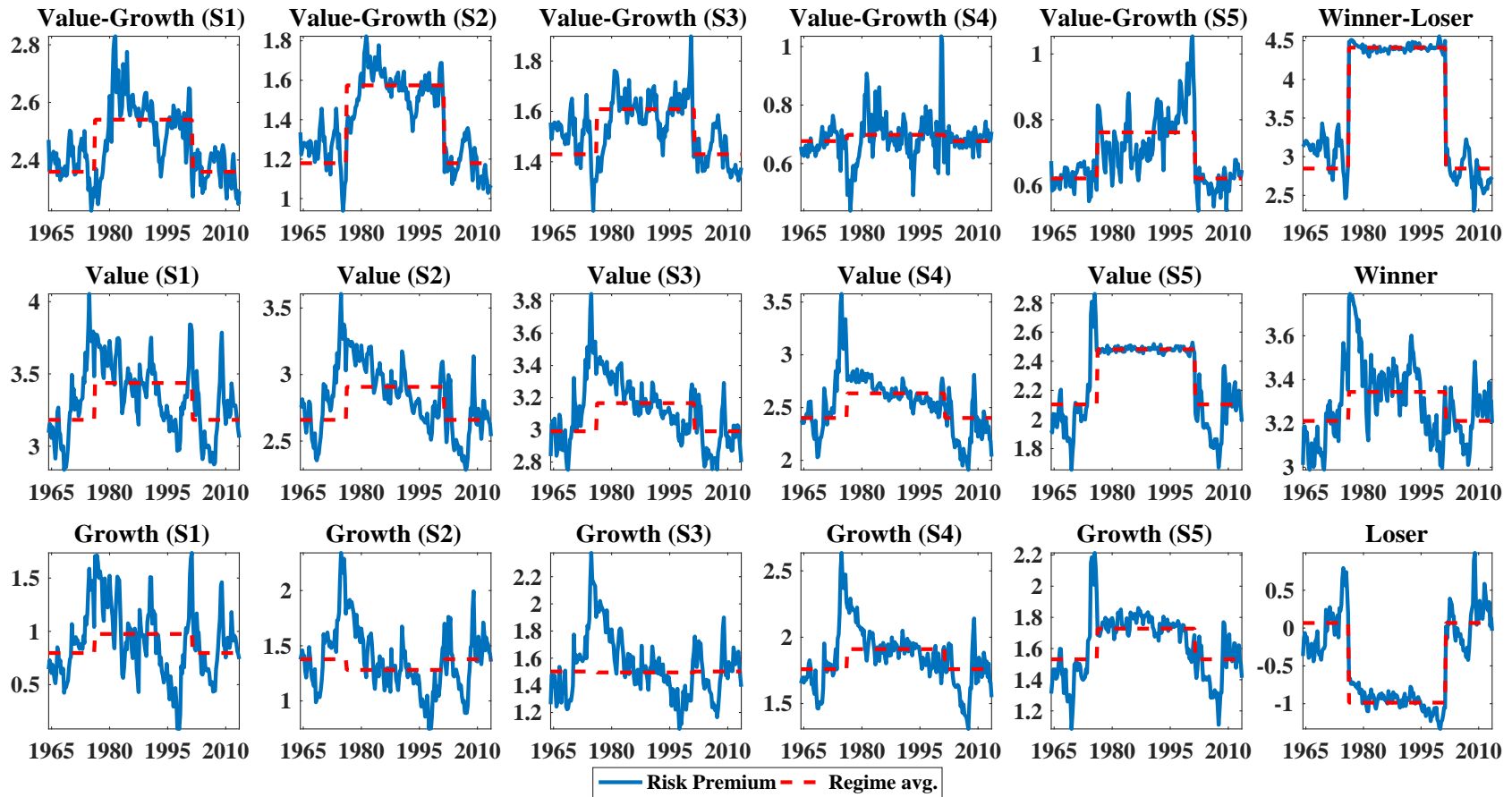
**Figure 10: Evolution of Risk Premia.** The figure reports the evolution of the the PDV of risk premia for different portfolios. The blue solid line reports the evolution of the risk premia over time, while the red dashed line corresponds to the conditional steady state of the PDV based on the regime in place. Both are computed by taking into account the possibility of regime changes. The sample spans the period 1964:Q1-2013:Q3.
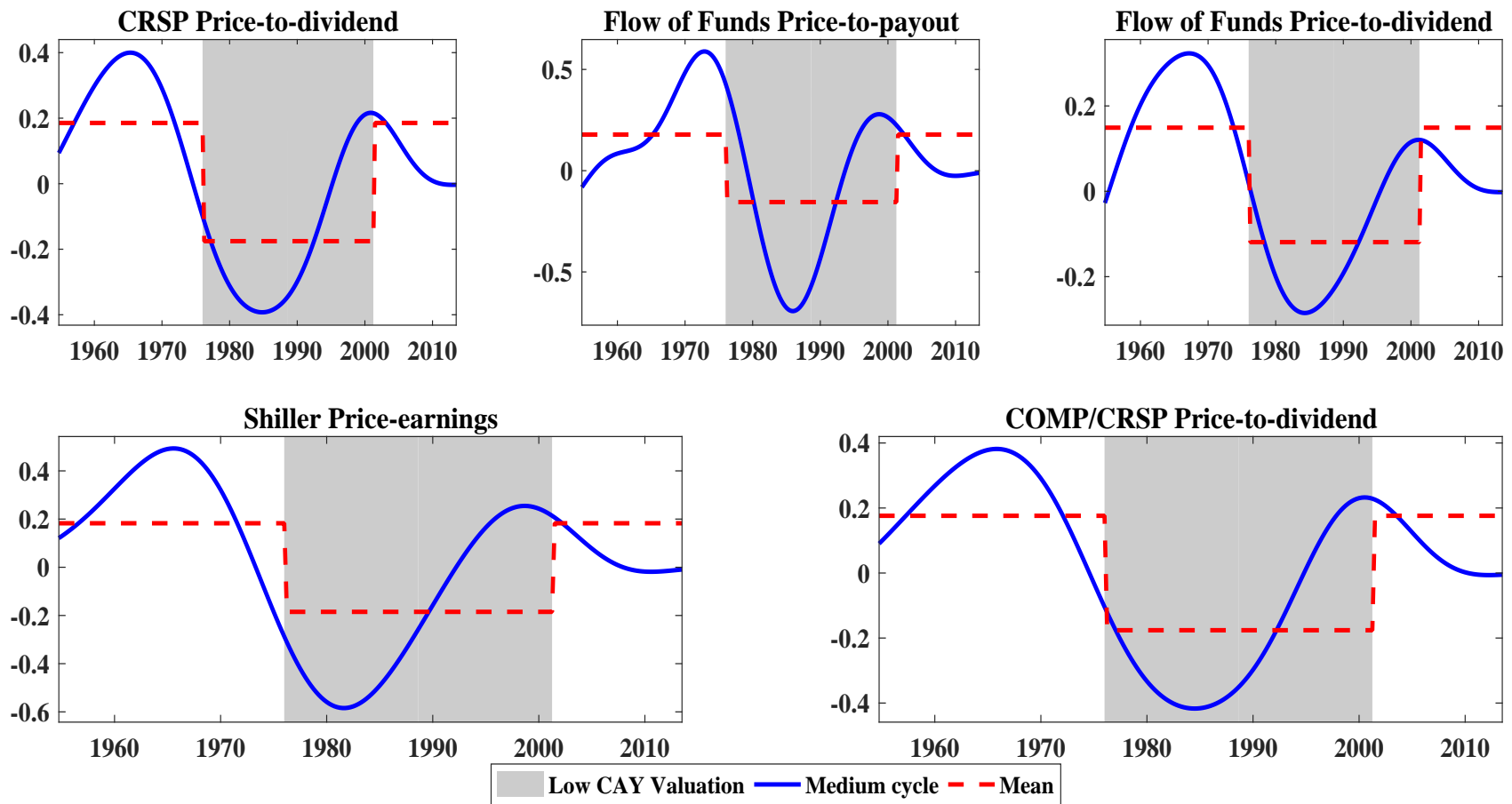
**Figure 11:** This figure presents the evolution of alternative valuation ratios at medium-term frequencies corresponding to cycles between 10 and 50 years. The sample spans 1955:Q3-2013:Q3.

# Appendix for Online Publication

## Data Appendix

This appendix describes the data used in this study.

CONSUMPTION

Consumption is measured as either total personal consumption expenditure or expenditure on nondurables and services, excluding shoes and clothing. The quarterly data are seasonally adjusted at annual rates, in billions of chain-weighted 2005 dollars. The components are chain-weighted together, and this series is scaled up so that the sample mean matches the sample mean of total personal consumption expenditures. Our source is the U.S. Department of Commerce, Bureau of Economic Analysis.

LABOR INCOME

Labor income is defined as wages and salaries + transfer payments + employer contributions for employee pensions and insurance - employee contributions for social insurance - taxes. Taxes are defined as [ wages and salaries/(wages and salaries + proprietors' income with IVA and CCADJ + rental income + personal dividends + personal interest income)] times personal current taxes, where IVA is inventory valuation and CCADJ is capital consumption adjustments. The quarterly data are in current dollars. Our source is the Bureau of Economic Analysis.

POPULATION

A measure of population is created by dividing real total disposable income by real per capita disposable income. Our source is the Bureau of Economic Analysis.

WEALTH

Total wealth is household net worth in billions of current dollars, measured at the end of the period. A break-down of net worth into its major components is given in the table below. Stock market wealth includes direct household holdings, mutual fund holdings, holdings of private and public pension plans, personal trusts, and insurance companies. Nonstock wealth includes tangible/real estate wealth, nonstock financial assets (all deposits, open market paper, U.S. Treasuries and Agency securities, municipal securities, corporate and foreign bonds and mortgages), and also includes ownership of privately traded companies in noncorporate equity, and other. Subtracted off are liabilities, including mortgage loans and loans made under home equity lines of credit and secured by junior liens, installment consumer debt and other. Wealth is measured at the end of the period. A timing convention for wealth is needed because the level of consumption is a flow during the quarter rather than a point-in-time estimate as is wealth (consumption data are time-averaged). If we think of a given quarter's consumption data as measuring spending at the beginning of the quarter, then wealth for the quarter should

be measured at the beginning of the period. If we think of the consumption data as measuring spending at the end of the quarter, then wealth for the quarter should be measured at the end of the period. None of our main findings discussed below (estimates of the cointegrating parameters, error-correction specification, or permanent-transitory decomposition) are sensitive to this timing convention. Given our finding that most of the variation in wealth is not associated with consumption, this timing convention is conservative in that the use of end-of-period wealth produces a higher contemporaneous correlation between consumption growth and wealth growth. Our source is the Board of Governors of the Federal Reserve System. A complete description of these data may be found at http://www.federalreserve.gov/releases/Z1/Current/.

CRSP PRICE-DIVIDEND RATIO

The stock price is measured using the Center for Research on Securities Pricing (CRSP) value-weighted stock market index covering stocks on the NASDAQ, AMEX, and NYSE. The data are monthly. The stock market price is the price of a portfolio that does not reinvest dividends. The CRSP dataset consists of $vwretx(t) = (P_t/P_{t-1}) - 1$, the return on a portfolio that doesn't pay dividends, and $vwretd_t = (P_t + D_t)/P_t - 1$, the return on a portfolio that does pay dividends. The stock price index we use is the price $P_t^x$ of a portfolio that does not reinvest dividends, which can be computed iteratively as

$$P_{t+1}^x = P_t^x \left(1 + vwretx_{t+1}\right),$$

where $P_0^x = 1$. Dividends on this portfolio that does not reinvest are computed as

$$D_t = P_{t-1}^x \left(vwretd_t - vwretx_t\right).$$

The above give monthly returns, dividends and prices. The annual log return is the sum of the 12 monthly log returns over the year. We create annual log dividend growth rates by summing the log differences over the 12 months in the year: $d_{t+12} - d_t = d_{t+12} - d_{t+11} + d_{t+11} - d_{t+10} + \cdots + d_{t+1} - d_t$. The annual log price-dividend ratio is created by summing dividends in levels over the year to obtain an annual dividend in levels, $D_t^A$, where $t$ denotes a year hear. The annual observation on $P_t^x$ is taken to be the last monthly price observation of the year, $P_t^{Ax}$. The annual log price-dividend ratio is $\ln\left(P_t^{Ax}/D_t^A\right)$. Note that this value for dividend growth is only used to compute the CRSP price-dividend ratio on this hypothetical portfolio. When we investigate the behavior of stock market dividend growth in the MS-VAR, we use actual dividend data from all firms on NYSE, NASDAQ, and AMEX. See the data description for MS-VARs below.

FLOW OF FUNDS EQUITY PAYOUT, DIVIDENDS, PRICE

Flow of Funds payout is measured as "Net dividends plus net repurchases" and is computed using the Flow of Funds Table F.103 (nonfinancial corporate business sector) by subtracting Net Equity Issuance (FA103164103) from Net Dividends (FA106121075). We define net repurchases

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to be repurchases net of share issuance, so net repurchases is the negative of net equity issuance. Net dividends consists of payments in cash or other assets, excluding the corporation's own stock, made by corporations located in the United States and abroad to stockholders who are U.S. residents. The payments are netted against dividends received by U.S. corporations, thereby providing a measure of the dividends paid by U.S. corporations to other sectors. The price used for FOF price-dividend and price-payout ratios is "Equity," the flow of funds measure of equities (LM103164103).

PRICE DEFLATOR FOR CONSUMPTION AND ASSET WEALTH

The nominal after-tax labor income and wealth data are deflated by the personal consumption expenditure chain-type deflator (2005=100), seasonally adjusted. In principle, one would like a measure of the price deflator for total flow consumption here. Since this variable is unobservable, we use the total expenditure deflator as a proxy. Our source is the Bureau of Economic Analysis.

DATA FOR MS-VARs

In the baseline MS-VAR, we use five observables: real R&D per capita growth, real investment per capita growth, real GDP per capita growth, annualized quarterly inflation, the federal funds rate. Our data sources are the NIPA tables constructed by the Bureau of Economic Analysis and the St. Louis Fed. Real GDP per capita is obtained by dividing nominal GDP (NIPA 1.1.5, line 1) by the GDP deflator (NIPA 1.1.4, line 1) and population. Consumption is defined as the sum of personal consumption expenditures on non-durable goods (NIPA 1.1.5, line 5) and services (NIPA 1.1.5, line 6). The series for nominal investment in physical capital is the sum of gross private domestic investment (NIPA 1.1.5, line 7) and personal consumption expenditure in durables (NIPA 1.1.5, line 4) *minus* intellectual property products (NIPA 1.1.5, line 12). Both series are then divided by the GDP deflator and population. Nominal R&D investment coincides with the series of intellectual property products (NIPA 1.1.5, line 12). The series is then divided by GDP deflator and population. Inflation is measured as the annual log-difference in the GDP deflator. The effective FFR is downloaded from the St. Louis Fed website, while all the other series are extracted from the NIPA tables constructed by the Bureau of Economic Analysis. The sample spans 1954:Q3 to 2013:Q3.

In the secondary MS-VAR that includes dividend growth and the change in the labor share, we compute dividends for all firms on NYSE, NASDAQ and AMEX using a merge of the Compustat Annual Industrial Database and the CRSP Monthly Stock Database, following the approach presented in Larrain and Yogo (2008). The labor share is defined as wages and salaries + employer contributions for employee pensions and insurance - employee contributions for social insurance. The quarterly data are in current dollars. Our source for these variables is the Bureau of Economic Analysis. Dividend growth is defined as the year-to-year change in log-dividends. The change in the labor share is computed by taking the year-to-year difference

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in the annual mean of the labor share. These two variables are combined in the VAR with real GDP per capita growth, annualized quarterly inflation, and the effective federal funds rate.

## Gibbs Sampling Algorithm

This appendix describes the Bayesian methods used to characterize uncertainty in the parameters of the regression (2). To simplify notation, we denote the vector containing all variables whose coefficients are allowed to vary over time $x_{M,t}$, while $x_{F,t}$ is used to denote the vector containing all the variables whose coefficients are kept constant. We then obtain:

$$c_t = \alpha_{\xi_t^\alpha} x_{M,t} + \beta x_{F,t} + \sigma \varepsilon_t$$

where, in our case, $\beta = \left[\beta_a, \beta_y, b_{a,-k}, ..., b_{a,+k}, b_{y,-k}, ..., b_{y,+k}\right]$ and the vector $x_{M,t}$ is unidimensional and always equal to 1.

Suppose the Gibbs sampling algorithm has reached the $r-th$ iteration. We then have draws for $\beta_r$, $\alpha_{\xi_t^\alpha,r}$, $\sigma_r$, $H_r^\alpha$, and $\xi_r^{\alpha,T}$, where $\xi_r^{\alpha,T} = \{\xi_{1,r}^\alpha, \xi_{2,r}^\alpha,...,\xi_{T,r}^\alpha\}$ denotes a draw for the whole regime sequence. The sampling algorithm is described as follows.

1. **Sampling $\beta_{r+1}$**: Given $\alpha_{\xi_t^\alpha,r}$, $\sigma_r$, and $\xi_r^{\alpha,T}$ we transform the data:

$$\widetilde{c}_t = \frac{c_t - \alpha_{\xi_t^\alpha,r} x_{M,t}}{\sigma_r} = \beta \frac{x_{F,t}}{\sigma_r} + \varepsilon_t = \beta \widetilde{x}_t + \varepsilon_t.$$

The above is a regression with fixed coefficients $\beta$ and standardized residual shocks. Standard Bayesian methods may be used to draw the coefficients of the regression. We assume a Normal conjugate prior $\beta \sim N\left(B_{\beta,0}, V_{\beta,0}\right)$, so that the conditional (on $\alpha_{\xi_t^\alpha,r}$, $\sigma_r$, and $\xi_r^{\alpha,T}$) posterior distribution is given by

$$\beta_{r+1} \sim N\left(B_{\beta,T}, V_{\beta,T}\right)$$

with $V_{\beta,T} = \left(V_{\beta,0}^{-1} + \widetilde{X}_F'\widetilde{X}_F\right)^{-1}$ and $B_{\beta,T} = V_{\beta,T}\left[V_{\beta,0}^{-1}B_{\beta,0} + \widetilde{X}_F'\widetilde{C}\right]$, where $\widetilde{C}$ and $\widetilde{X}_F$ collect all the observations for the transformed data and $B_{\beta,0}$ and $V_{\beta,0}^{-1}$ control the priors for the fixed coefficients of the regression. With flat priors, $B_{\beta,0} = 0$ and $V_{\beta,0}^{-1} = 0$ and $B_{\beta,T}$ and $V_{\beta,T}$ coincide with the maximum likelihood estimates, conditional on the other parameters.

2. **Sampling $\alpha_{i,r+1}$ for $i = 1, 2$**: Given $\beta_{r+1}$, $\sigma_r$, and $\xi_r^{\alpha,T}$ we transform the data:

$$\widetilde{c}_t = \frac{c_t - \beta_{r+1} x_{F,t}}{\sigma_r} = \alpha_{\xi_t^\alpha} \frac{x_{M,t}}{\sigma_r} + \varepsilon_t = \alpha_{\xi_t^\alpha} \widetilde{x}_{M,t} + \varepsilon_t.$$

The above regression has standardized shocks and Markov-switching coefficients in the transformed data. Using $\xi_r^{\alpha,T}$ we can group all the observations that pertain to the same

regime $i$. Given the prior $\alpha_i \sim N\left(B_{\alpha_i,0}, V_{\alpha_i,0}\right)$ for $i = 1, 2$ we use standard Bayesian methods to draw $\alpha_i$ from the conditional (on $\beta_{r+1}$, $\sigma_r$, and $\xi_r^{\alpha,T}$) posterior distribution:

$$\alpha_{i,r+1} \sim N\left(B_{\alpha_i,T}, V_{\alpha_i,T}\right) \text{ for } i = 1, 2$$

where $V_{\alpha_i,T} = \left(V_{\alpha_i,0}^{-1} + \widetilde{X}'_{M,i}\widetilde{X}_{M,i}\right)^{-1}$ and $B_{\alpha_i,T} = V_{\alpha_i,T}\left[V_{\alpha_i,0}^{-1}B_{\alpha_i,0} + \widetilde{X}'_{M,i}\widetilde{C}_i\right]$ where $\widetilde{C}_i$ and $\widetilde{X}_{M,i}$ collect all the observations for the transformed data for which regime $i$ is in place. The parameters $B_{\alpha_i,0}$ and $V_{\alpha_i,0}^{-1}$ control the priors for the MS coefficients of the regression: $\alpha_i \sim N\left(B_{\alpha_i,0}, V_{\alpha_i,0}\right)$ for $i = 1, 2$. With flat priors, we have $B_{\alpha_i,0} = 0$ and $V_{\alpha_i,0}^{-1} = 0$ and $B_{\alpha_i,T}$ and $V_{\alpha_i,T}$ coincide with the maximum likelihood estimates, conditional on the other parameters.

3. **Sampling** $\sigma_{r+1}$: Given $\beta_{r+1}$, $\alpha_{\xi_t^\alpha,r+1}$, and $\xi_r^{\alpha,T}$ we can compute the residuals of the regression:

$$\widetilde{c}_t = c_t - \beta_{r+1}x_{F,t} - \alpha_{\xi_t^\alpha}x_{M,t} = \sigma\varepsilon_t.$$

With the prior that $\sigma$ has an inverse gamma distribution, $\sigma \sim IG\left(Q_0, v_0\right)$, we use Bayesian methods to draw $\sigma_{r+1}$ from the conditional (on $\beta_{r+1}$, $\alpha_{\xi_t^\alpha,r+1}$, and $\xi_r^{\alpha,T}$) posterior inverse gamma distribution:

$$\sigma_{r+1} \sim IG\left(Q_T, v_T\right), \ v_T = T + v_0, \ Q_T = Q_0 + E'E$$

where $E$ is a vector containing the residuals, $T$ is the sample size, and $Q_0$ and $v_0$ control the priors for the standard deviation of the innovations: $\sigma \sim IG\left(Q_0, v_0\right)$. With flat priors, we have $Q_0 = 0$ and $v_0 = 0$. The mean of a random variable with distribution $\sigma \sim IG\left(Q_T, v_T\right)$ is $Q_T/v_T$. With flat priors we have $Q_0 = 0$ and $v_0 = 0$, and the mean of $\sigma$ is therefore $\left(E'E\right)/T$, which coincides with the standard maximum likelihood (MLE) estimate of $\sigma$, conditional on the other parameters.

4. **Sampling** $\xi_{r+1}^{\alpha,T}$: Given $\beta_{r+1}$, $\alpha_{\xi_t^\alpha,r+1}$, and $H_r^\alpha$ we can obtain filtered probabilities for the regimes, as described in Hamilton (1994). Following Kim and Nelson (1999) we then use a Multi-Move Gibbs sampling to draw a regime sequence $\xi_{r+1}^{\alpha,T}$.

5. **Sampling** $H_{r+1}^\alpha$: Given the draws for the MS state variables $\xi_{r+1}^{\alpha,T}$, the posterior for the transition probabilities does not depend on other parameters of the model and follows a Dirichlet distribution if we assume a prior Dirichlet distribution.[19] For each column of $H_{r+1}^\alpha$ the posterior distribution is given by:

$$H_{r+1}^\alpha(:,i) \sim D(a_{ii}^\alpha + \eta_{ii,r+1}^\alpha, a_{ij}^\alpha + \eta_{ij,r+1}^\alpha)$$

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[19] The Dirichlet distribution is a generalization of the beta distribution that allows one to potentially consider more than 2 regimes. See e.g., Sims and Zha (2006).

where $\eta_{ij,r+1}^{\alpha}$ denotes the number of transitions from state $i^{\alpha}$ to state $j^{\alpha}$ based on $\xi_{r+1}^{\alpha,T}$, while $a_{ii}^{\alpha}$ and $a_{ij}^{\alpha}$ the corresponding priors. With flat priors, we have $a_{ii}^{\alpha} = 0$ and $a_{ij}^{\alpha} = 0$.

6. If $r + 1 < R$, where $R$ is the desired number of draws, go to step 1, otherwise stop.

These steps are repeated until convergence to the posterior distribution is reached. We check convergence by using the Raftery-Lewis Diagnostics for each parameter in the chain. See section below. We use the draws obtained with the Gibbs sampling algorithm to characterize parameter uncertainty in Table 1.

# Computing $cay^{MS}$

Our estimate of $cay_t^{MS}$ is based on the posterior mode of the parameter vector $\boldsymbol{\theta}$ and the corresponding regime probabilities. The filtered probabilities reflect the probability of a regime conditional on the data up to time $t$, $\pi_{t|t} = p(\xi_t^{\alpha}|Y^t; \boldsymbol{\theta})$, for $t = 1, ..., T$, and are part of the output obtained computing the likelihood function associated with the parameter draw $\boldsymbol{\theta} = \{\beta, \alpha_{\xi_t^{\alpha}}, \sigma, \boldsymbol{H}^{\alpha}\}$. They can be obtained using the following recursive algorithm:

$$\pi_{t|t} = \frac{\pi_{t|t-1} \odot \eta_t}{\mathbf{1}' \left(\pi_{t|t-1} \odot \eta_t\right)}$$
$$\pi_{t+1|t} = \boldsymbol{H}^{\alpha}\pi_{t|t}$$

where $\eta_t$ is a vector whose $j$-th element contains the conditional density $p(c_t|\xi_t^{\alpha} = j, x_{M,t}, x_{F,t}; \boldsymbol{\theta})$, i.e.,

$$p(c_t|\xi_t^{\alpha} = j, x_{M,t}, x_{F,t}; \boldsymbol{\theta}) = \frac{1}{\sqrt{2\pi\sigma^2}}\exp\left(-\frac{\{c_t - (\alpha_j x_{M,t} + \beta x_{F,t})\}^2}{2\sigma^2}\right),$$

the symbol $\odot$ denotes element by element multiplication, and $\mathbf{1}$ is a vector with all elements equal to 1. To initialize the recursive calculation we need an assumption on the distribution of $\xi_0^{\alpha}$. We assume that the two regimes have equal probabilities: $p(\xi_0^{\alpha} = 1) = .5 = p(\xi_0^{\alpha} = 2)$.

The smoothed probabilities reflect all the information that can be extracted from the whole data sample, $\pi_{t|T} = p(\xi_t^{\alpha}|Y^T; \boldsymbol{\theta})$. The final term, $\pi_{T|T}$ is returned with the final step of the filtering algorithm. Then, a recursive algorithm can be implemented to derive the other probabilities:

$$\pi_{t|T} = \pi_{t|t} \odot \left[\boldsymbol{H}^{\alpha\prime}\left(\pi_{t+1|T}\left(\div\right)\pi_{t+1|t}\right)\right]$$

where $(\div)$ denotes element by element division.

In using the DLS regression (2) to estimate cointegrating parameters, we lose 6 leads and 6 lags. For estimates of $cay_t^{FC}$, we take the estimated coefficients and we apply them to the whole sample. To extend our estimates of $cay_t^{MS}$ over the full sample, we can likewise apply the parameter estimates to the whole sample but we need an estimate of the regime probabilities

in the first 6 and last 6 observations of the full sample. For this we run the Hamilton filter from period from $-5$ to $T+6$ as follows. When starting at -5, we assume no lagged values are available and the DLS regression omits all lags, but all leads are included. When at $t = -4$ we assume only one lag is available and the DLS regression includes only one lag, and so on, until we reach $t = 0$ when all lags are included. Proceeding forward when $t = T + 1$ is reached we assume all lags are available and all leads except one are available, when $t = T + 2$, we assume all lags and all leads but two are available, etc. Smoothed probabilities are then computed with standard methods as they only involve the filtered probabilities and the transition matrix $\boldsymbol{H}^{\alpha}$.

## Model Comparison

This section reports the results of tests that compare competing models for $cay_t^{MS}$. We analyze two alternative models and compare them to our two benchmark models for $cay$: $cay^{MS}$ in which only the constant is allowed to switch over time, and $cay^{FC}$ in which there is no Markov-switchin in any of the parameters. In the first alternative, we allow for Markov-switching in heteroskedasticity as well as Markov-switching in the constant. In the second alternative, we only allow for Markov-switching in only heteroskedasticity. We use the Bayesian information criterion ($BIC$) to compare these different models. This is computed as:

$$BIC = -2(maxli) + k \log(T)$$

where $maxli$ is the maximized $log$-likelihood, $k$ is the number of parameters, and $T$ the sample size. A model with a smaller $BIC$ criterion is preferred to one with a larger criterion. The $BIC$ criterion rewards models with higher likelihoods but also penalizes models that have more parameters.

Table A.1 reports the estimates for the key $cay$ parameters and the $BIC$ for each model. The $BIC$ is lowest by the model is the one that allows for switching in both heteroskedasticity and the constant (MS $\alpha$ and MS $\sigma$). But our benchmark model with switches only in the constant (MS $\alpha$ only) is preferred to the model that with switches only in heteroskedasticity (MS $\sigma$ only), and to the fixed coefficient model (FC). These results support the hypothesis of shifts in the constant. And although the model with switching in both the constant and the volatility is preferred, the estimates for the cointegrating vector and the timing of regimes are essentially unchanged when introducing heteroskedasticity in our benchmark model. For this reason, we choose the simpler model with only shifts in the constant as our benchmark model.

## Cosine Transformations

Figure A.1 is based on weighted averages that summarize low-frequency variability in a series. Specifically, following Muller and Watson (2008) and Watson (2013), the figure plots the "cosine

transformations" of each version of *cay*

$$f_j = \sum_{t=1}^{T} \cos\left(j(t - 0.5)\pi T^{-1}\right) cay_t \quad \text{for } j = 1, ..., k.$$

As Muller and Watson (2008) show, the set of sample averages $\{f_j\}_{j=1}^{k}$, capture the variability in *cay* for periods greater than $2T/k$, where $T$ is the sample size. Thus, with $T = 247$ quarters, the $k = 12$ points plotted in Figure A.1 summarize the variability in *cay* for periods greater than $2*247/12 = 41.1667$ quarters, or approximately 10 years. Smaller values of $j$ correspond to lower frequencies, so values of $f_j$ plotted for small $j$ (e.g., $j = 1, 2, 3$) give the variability in $cay_t$ at low frequencies, while values of $f_j$ plotted for higher $j$ (e.g., $j = 10, 11, 12$) give the variability in $cay_t$ at higher frequencies. A series that is integrated of order zero, $I(0)$, corresponding to covariance stationary, displays roughly the same variability (same value of $f_j$) at all frequencies $j$. By contrast, a series that is more persistent than $I(0)$ displays higher variability at low frequencies, resulting in higher values of $f_j$ for low $j$ than for high $j$. Figure A.1 shows that the cosine transformation of $cay_t^{MS}$ displays a pattern much more consistent with an $I(0)$ series than that of $cay_t^{FC}$, which shows a clear spike at $j = 3$, corresponding to a period of roughly 41 years.
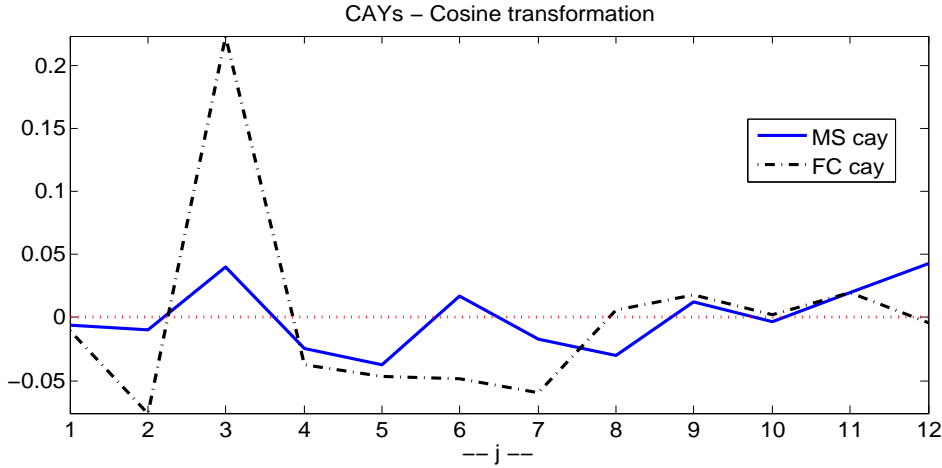


**Figure A.1:** Low frequency averages of *cay*. The figure plots the set of averages $\{f_j\}_{j=1}^{k}$, which capture the variability in *cay* for periods greater than $2T/k$, where $T$ is the sample size. Thus, with T = 247 quarters, the $k = 12$ points plotted summarize the variability in *cay* for periods greater than $2*247/12 = 41.1667$ quarters, approximately 10 years. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

| Model | $\alpha_1$ | $\alpha_2$ | $\beta_a$ | $\beta_y$ | $\sigma_1$ | $\sigma_2$ | $BIC$ |
|---|---|---|---|---|---|---|---|
| MS $\alpha$ and MS $\sigma$ | 0.9186 | 0.8810 | 0.2599 | 0.6162 | 0.0016 | 0.0105 | $-1472.0$ |
| MS $\alpha$ only | 0.9186 | 0.8808 | 0.2606 | 0.6156 | 0.0080 | | $-1390.3$ |
| MS $\sigma$ only | | 0.8056 | 0.1275 | 0.7845 | 0.0029 | 0.0204 | $-1230.2$ |
| FC | | 0.8706 | 0.1246 | 0.7815 | 0.0158 | | $-1202.2$ |

**Table A.1:** The table reports the estimates for the cointegration parameters, the estimates for the volatilities, and the Bayesian Information Criterion (BIC) for four different models. The BIC is used to compare the fit of different models taking into account the number of parameters used in the estimates. MS $\alpha$ and MS $\sigma$: The model allows for changes in the constant and heteroskedasticity. MS $\alpha$ only: Benchmark model with only changes in the constant. MS $\sigma$ only: The model allows for heteroskedasticity, but not changes in the constant. FC: Standard fixed coefficient regression.

## Estimation of Fractionally Integrated Models

In order to evaluate the likelihood for the fractionally integrated model we closely follow Muller and Watson (2013). We in fact use a series of Matlab codes that are available on Mark Watson's webpage. The first step consists of computing the cosine transformation of cay:

$$f_j = \iota_{jT} T^{-1} \sum_{t=1}^{T} \sqrt{2} \cos\left(j(t-0.5)\pi T^{-1}\right) cay_t \quad \text{for } j = 1, ..., k.$$

where $\iota_{jT} = (2T/(j\pi)) \sin(j\pi/(2T))$. As explained in Muller and Watson (2013), this transformation is useful to isolate variation in the sample at different frequencies. Specifically, $f_j$ captures variation at frequency $j\pi/T$. Mueller and Watson (2008, 2013) explain that working with a subset of the cosine transformations implies truncating the information set. They provide two reasons for why this is a convenient approach. First, given that each variable is a weighted average of the original data, a central limit allows to work with a limiting Gaussian distribution. Second, such a choice implies robustness of the results: Low-frequency information is used to study the low-frequency properties of the model. Given that we are mostly interested in the low frequency properties of $cay$, we can work using a limited number of (low) frequencies. We therefore choose $k = 12$.

We can then collect all the cosine transformations in a vector $X_{T,1:k}$ and compute an invariant transformation $X_{T,1:k}^s = X_{T,1:k}/\sqrt{X_{T,1:k}' X_{T,1:k}}$ (notice that this implies that the results that will follow are independent of scale factors). As explained in Muller and Watson (2013), the limiting density for the invariant transformations is given by:

$$p_{X^s}(x^s) = \frac{1}{2} \Gamma(k/2) \pi^{-k/2} |\Sigma_X|^{-1/2} \left(x^{s\prime} \Sigma_X^{-1} x^s\right)^{-q/2} \tag{A1}$$

where $X^s = X_{1:k}/\sqrt{X'_{1:k}X_{1:k}}$, $\Sigma_X = E\left(X^s X^{s\prime}\right)$, and $\Gamma$ is the gamma function.

We then assume a fractionally integrated model for $cay_t$: $(1-L)^d cay_t = u_t$, where $L$ is the lag operator and $u_t$ is an $I(0)$ process and $d$ is a parameter that is allowed to be fractional. The fractional model implies a binomial series expansion in the lag operator:

$$
\begin{aligned}
(1-L)^d cay_t &= \left[\sum_{k=0}^{\infty} \binom{d}{k}(-L)^k\right] cay_t \\
&= \left[\sum_{k=0}^{\infty} \frac{\prod_{a=0}^{k-1}(d-a)(-L)^k}{k!}\right] cay_t \\
&= \left[1 - dL + \frac{d(d-1)}{2!}L^2 - ...\right] cay_t
\end{aligned}
$$

Note that when $d=1$, the fractional integrated model implies that $cay_t$ has a unit root, $cay_t = cay_{t-1} + u_t$, while for $d=0$, $cay_t = u_t$, i.e. $cay_t$ is an $I(0)$ process.

We compute the covariance matrix $\Sigma_X(d)$ associated with different values of $d$ in the fractionally integrated model. The matrix $\Sigma_X(d)$ is obtained in two steps. First, we compute the matrix of autocovariances $\Sigma(d)$ associated with a fractionally integrated model. The $(i, i+h)$ element of this matrix is given by the autocovariance $\gamma(h)$:

$$
\Sigma(d)_{(i,i+h)} = \gamma(h) = \frac{\Gamma(1-2d)}{\Gamma(1-d)\Gamma(d)} \frac{\Gamma(h+d)}{\Gamma(1+h-d)}
$$

Second, we transform the autocovariance matrix $\Sigma(d)$ in order to obtain the covariance matrix for the cosine transformations: $\Sigma_X(d) = \Psi'\Sigma(d)\Psi$ where $\Psi$ is a $(T \times k)$ matrix collecting all the weights used for the cosine transformation:

$$
\Psi_{(t,j)} = \iota_{jT}T^{-1}\sum_{t=1}^{T}\sqrt{2}\cos\left(j(t-0.5)\pi T^{-1}\right)
$$

Finally, we evaluate (A1) to obtain the likelihood for the different values of $d$ given that $\Sigma_X(d)$ is now a function of the parameter $d$ of the fractionally integrated model.

## Additional Statistical Results

The tables below pertain to convergence of the Gibbs sampling algorithm.

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| Variable | Total(N) | I-stat | Variable | Total(N) | I-stat | Variable | Total(N) | I-stat |
|---|---|---|---|---|---|---|---|---|
| $\alpha_1$ | 17413 | 9.541 | $\Delta a_{t+1}$ | 1799 | 0.986 | $\Delta y_{t-4}$ | 1850 | 1.014 |
| $\alpha_2$ | 16949 | 9.287 | $\Delta y_{t+1}$ | 1812 | 0.993 | $\Delta a_{t+4}$ | 1793 | 0.982 |
| $\beta_a$ | 1918 | 1.051 | $\Delta a_{t-2}$ | 1830 | 1.003 | $\Delta y_{t+4}$ | 1820 | 0.997 |
| $\beta_y$ | 1843 | 1.01 | $\Delta y_{t-2}$ | 1801 | 0.987 | $\Delta a_{t-5}$ | 1797 | 0.985 |
| $\sigma$ | 1797 | 0.985 | $\Delta a_{t+2}$ | 1886 | 1.033 | $\Delta y_{t-5}$ | 1850 | 1.014 |
| $H_{11}^\alpha$ | 1826 | 1.001 | $\Delta y_{t+2}$ | 1767 | 0.968 | $\Delta a_{t+5}$ | 1826 | 1.001 |
| $H_{22}^\alpha$ | 1820 | 0.997 | $\Delta a_{t-3}$ | 1858 | 1.018 | $\Delta y_{t+5}$ | 1850 | 1.014 |
| $\Delta a_t$ | 1823 | 0.999 | $\Delta y_{t-3}$ | 1808 | 0.991 | $\Delta a_{t-6}$ | 1850 | 1.014 |
| $\Delta y_t$ | 1850 | 1.014 | $\Delta a_{t+3}$ | 1847 | 1.012 | $\Delta y_{t-6}$ | 1826 | 1.001 |
| $\Delta a_{t-1}$ | 1839 | 1.008 | $\Delta y_{t+3}$ | 1820 | 0.997 | $\Delta a_{t+6}$ | 1839 | 1.008 |
| $\Delta y_{t-1}$ | 1866 | 1.022 | $\Delta a_{t-4}$ | 1830 | 1.003 | $\Delta y_{t+6}$ | 1866 | 1.022 |

**Table A.2:** Raftery-Lewis Diagnostics for each parameter in the chain. The minimum number of draws under the assumption of i.i.d. draws would be 1825. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.

## MS-VAR Estimation

In this appendix we provide details on the estimation of the MS-VAR 8. As stated above, we take the regime sequence as given based on our estimates for the breaks in $cay^{MS}$. Specifically, we choose the particular regime sequence $\hat{\xi}^{\alpha,T} = \{\hat{\xi}_1^\alpha, ..., \hat{\xi}_T^\alpha\}$ that is most likely to have occurred, given our estimated posterior mode parameter values for $\boldsymbol{\theta}$ . This sequence is computed as follows. First, we run Hamilton's filter to get the vector of filtered probabilities $\pi_{t|t}$, $t = 1, 2, ..., T$. The Hamilton filter can be expressed iteratively as

$$\pi_{t|t} = \frac{\pi_{t|t-1} \odot \eta_t}{\mathbf{1}' \left(\pi_{t|t-1} \odot \eta_t\right)}$$
$$\pi_{t+1|t} = \boldsymbol{H}^\alpha \pi_{t|t}$$

where $\eta_t$ is a vector whose $j$-th element contains the conditional density $p(c_t|\xi_t^\alpha = j, x_{M,t}, x_{F,t}; \boldsymbol{\theta})$, the symbol $\odot$ denotes element by element multiplication, and $\mathbf{1}$ is a vector with all elements equal to 1. The final term, $\pi_{T|T}$ is returned with the final step of the filtering algorithm. Then, a recursive algorithm can be implemented to derive the other smoothed probabilities:

$$\pi_{t|T} = \pi_{t|t} \odot \left[\boldsymbol{H}^{\alpha\prime} \left(\pi_{t+1|T} \left(\div\right) \pi_{t+1|t}\right)\right]$$

where $(\div)$ denotes element by element division. To choose the regime sequence most likely to have occurred given our parameter estimates, consider the recursion in the next to last period

$t = T - 1$:

$$\pi_{T-1|T} = \pi_{T-1|T-1} \odot \left[ \boldsymbol{H}^{\alpha\prime} \left( \pi_{T|T} \left( \div \right) \pi_{T|T-1} \right) \right].$$

We first take $\pi_{T|T}$ from the Hamilton filter and choose the regime that is associated with the largest probability, i.e., if $\pi_{T|T} = (.9, .1)$, where the first element corresponds to the probability of regime 1, we select $\hat{\xi}_T^\alpha = 1$, indicating that we are in regime 1 in period $T$. We now update $\pi_{T|T} = (1, 0)$ and plug into the right-hand-side above along with the estimated filtered probabilities for $\pi_{T-1|T-1}$, $\pi_{T|T-1}$ and estimated transition matrix $\boldsymbol{H}^\alpha$ to get $\pi_{T-1|T}$ on the left-hand-side. Now we repeat the same procedure by choosing the regime for $T - 1$ that has the largest probability at $T - 1$, e.g., if $\pi_{T-1|T} = (.2, .8)$ we select $\hat{\xi}_{T-1}^\alpha = 2$, indicating that we are in regime 2 in period $T - 1$, we then update to $\pi_{T-1|T} = (0, 1)$, which is used again on the right-hand-side now

$$\pi_{T-2|T} = \pi_{T-2|T-2} \odot \left[ \boldsymbol{H}^{\alpha\prime} \left( \pi_{T-1|T} \left( \div \right) \pi_{T-1|T-2} \right) \right].$$

We proceed in this manner until we have a regime sequence $\hat{\xi}^{\alpha,T}$ for the entire sample $t = 1, 2, ..., T$. Two aspects of this procedure are worth noting. First, it fails if the updated probabilities are exactly $(.5, .5)$. Mathematically this is virtually zero. Second, note that this procedure allows us to choose the most likely regime sequence by using the recursive formula above to update the filtered probabilities sequentially from $T$ to time $t = 1$. This allows us to take into account the time dependence in the regime sequence as dictated by the transition probabilities.

Taking regime sequence as given in this way, we need only estimate the transition matrix and the parameters of the MS-VAR across the two regimes. The model is estimated by using Bayesian methods with flat priors on all parameters. As a first step, we group all the observations that belong to the same regime. Conditional on a regime, we have a fixed coefficients VAR. We can then follow standard procedures to make draws for the VAR parameters as follows.

Rewrite the VAR as

$$\underset{T \times n}{Y} = \underset{(T \times k)(k \times n)}{X A_{\xi_t}} + \underset{T \times n}{\varepsilon}, \ \xi_t = 1, 2$$

$$\varepsilon_t \sim N\left(0, \Sigma_{\xi_t}\right)$$

where $Y = [Z_1, ..., Z_T]'$, the $t$-th row of $X$ is $X_t = \left[1, Z'_{t-1}, Z'_{t-2}\right]$, $A_{\xi_t} = \left[c_{\xi_t}, A_{1,\xi_t}, A_{2,\xi_t}\right]'$, the $t$-th row of $\varepsilon$ is $\varepsilon_t$, and where $\Sigma_{\xi_t} = V_{\xi_t} V'_{\xi_t}$. If we specify a Normal-Wishart prior for $A_{\xi_t}$ and $V_{\xi_t}$:

$$\Sigma_{\xi_t}^{-1} \sim W\left(S_0^{-1}/v_0, v_0\right)$$

$$vec\left(A_{\xi_t}|\Sigma_{\xi_t}\right) \sim N\left(vec\left(B_0\right), \Sigma_{\xi_t} \otimes N_0^{-1}\right)$$

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where $E\left(\Sigma_{\xi_t}^{-1}\right) = S_0^{-1}$, the posterior distribution is still in the Normal-Wishart family and is given by

$$\begin{aligned}
\Sigma_{\xi_t}^{-1} &\sim W\left(S_T^{-1}/v_T, v_T\right) \\
vec\left(A_{\xi_t}|\Sigma_{\xi_t}\right) &\sim N\left(vec\left(B_T\right), \Sigma_{\xi_t} \otimes N_T^{-1}\right)
\end{aligned}$$

Using the estimated regime sequence $\xi^{\alpha,T}$ we can group all the observations that pertain to the same regime $i$. Therefore the parameters of the posterior are computed as

$$\begin{aligned}
v_T &= T_i + v_0, \quad N_T = X_i'X_i + N_0 \\
B_T &= N_T^{-1}\left(N_0 B_0 + X_i'X_i\widehat{B}_{MLE}\right) \\
S_T &= \frac{v_0}{v_T}S_0 + \frac{T_i}{v_T}\widehat{\Sigma}_{MLE} + \frac{1}{v_T}\left(\widehat{B}_{MLE} - \widehat{B}_0\right)' N_0 N_T^{-1} X_i'X_i\left(\widehat{B}_{MLE} - \widehat{B}_0\right) \\
\widehat{B}_{MLE} &= (X_i'X_i)^{-1}(X_i'Y_i), \quad \widehat{\Sigma}_{MLE} = \frac{1}{T_i}\left(Y_i - X_i\widehat{B}_{MLE}\right)'\left(Y_i - X_i\widehat{B}_{MLE}\right),
\end{aligned}$$

where $T_i, Y_i, X_i$ denote the number and sample of observations in regime $i$. We choose flat priors $(v_0 = 0, N_0 = 0)$ so the expressions above coincide with the MLE estimates using observations in regime $i$:

$$v_T = T_i, \ N_T = X_i'X_i, \ B_T = \widehat{B}_{MLE}, \ S_T = \widehat{\Sigma}_{MLE}.$$

Armed with these parameters in each regime, we can make draws from the posterior distributions for $\Sigma_{\xi_t}^{-1}$ and $A_{\xi_t}$ in regime $i$ to characterize parameter uncertainty about these parameters.

Given that we condition the MS-VAR estimates on the most likely regime sequence $\hat{\xi}^{\alpha,T}$ for $cay^{MS}$, it is still of interest to estimate the elements of the transition probability matrix for the MS-VAR parameters, $\boldsymbol{H}^A$, conditional on this regime sequence. Note that $\boldsymbol{H}^A$ can be different from $\boldsymbol{H}^\alpha$ because the former is based on a particular regime sequence $\hat{\xi}^{\alpha,T}$, while the latter reflects the entire posterior distribution for $\xi^{\alpha,T}$. The estimated transition matrix $\boldsymbol{H}^A$ can in turn be used to compute expectations taking into account the possibility of regime change (see the next subsection). Because we impose the regime sequence to be the same as that estimated for $cay^{MS}$, the posterior of $\boldsymbol{H}^A$ only depends on $\xi^{\alpha,T} = \xi$ and does not depend on other parameters of the model. The posterior has a Dirichlet distribution if we assume a prior Dirichlet distribution.[20] For each column of $\boldsymbol{H}^A$ the posterior distribution is given by:

$$\boldsymbol{H}^A(:,i) \sim D(a_{ii} + \eta_{ii,r+1}, a_{ij} + \eta_{ij,r+1})$$

where $\eta_{ij,r+1}$ denotes the number of transitions from regime $i$ to regime $j$ based on $\xi^{\alpha,T}$, while $a_{ii}$ and $a_{ij}$ the corresponding priors. With flat priors, we have $a_{ii} = 0$ and $a_{ij} = 0$. Armed with this posterior distribution, we can characterize uncertainty about $\boldsymbol{H}^A$.

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[20]The Dirichlet distribution is a generalization of the beta distribution that allows one to potentially consider more than 2 regimes. See e.g., Sims and Zha (2006).

## Conditional Expectations and Economic Uncertainty

In this appendix we explain how expectations and economic uncertainty are computed for variables in the MS-VAR. More details can be found in Bianchi (2016). Consider the following first-order MS-VAR:

$$Z_t = c_{\xi_t} + A_{\xi_t} Z_{t-1} + V_{\xi_t} \varepsilon_t, \ \varepsilon_t \sim N(0, I) \tag{A2}$$

and suppose that we are interested in $\mathbb{E}_0(Z_t) = \mathbb{E}(Z_t | \mathbb{I}_0)$ with $\mathbb{I}_0$ being the information set available at time 0. Note that the first-order VAR is not restrictive because any VAR with $l > 1$ lags can be rewritten as above by using the first-order companion form, and the methods below applied to the companion form.

Let $n$ be the number of variables in the VAR of the previous Appendix section. Let $m$ be the number of Markov-switching states. Define the $mn \times 1$ column vector $q_t$ as:

$$\underset{mn \times 1}{q_t} = \left[ q_t^{1\prime}, ..., q_t^{m\prime} \right]'$$

where the individual $n \times 1$ vectors $q_t^i = \mathbb{E}_0 \left( Z_t 1_{\xi_t = i} \right) \equiv \mathbb{E} \left( Z_t 1_{\xi_t = i} | \mathbb{I}_0 \right)$ and $1_{\xi_t = i}$ is an indicator variable that is one when regime $i$ is in place and zero otherwise. Note that:

$$q_t^i = \mathbb{E}_0 \left( Z_t 1_{\xi_t = i} \right) = \mathbb{E}_0 \left( Z_t | \xi_t = i \right) \pi_t^i$$

where $\pi_t^i = P_0 \left( \xi_t = i \right) = P \left( \xi_t = i | \mathbb{I}_0 \right)$. Therefore we can express $\mu_t = \mathbb{E}_0 (Z_t)$ as:

$$\mu_t = \mathbb{E}_0(Z_t) = \sum_{i=1}^m q_t^i = w q_t$$

where the matrix $\underset{n \times mn}{w} = [I_n, ..., I_n]$ is obtained placing side by side $m$ $n$-dimensional identity matrices. Then the following proposition holds:

**Proposition 1** *Consider a Markov-switching model whose law of motion can be described by (A2) and define $q_t^i = \mathbb{E}_0 \left( Z_t 1_{\xi_t = i} \right)$ for $i = 1...m$. Then $q_t^j = c_j \pi_t^j + \sum_{i=1}^m A_j q_{t-1}^i h_{ji}$.*

It is then straightforward to compute expectations conditional on the information available at a particular point in time. Suppose we are interested in $\mu_{t+s|t} \equiv \mathbb{E}_t(Z_{t+s})$, i.e. the expected value for the vector $Z_{t+s}$ conditional on the information set available at time $t$. If we define:

$$q_{t+s|t} = \left[ q_{t+s|t}^{1\prime}, ..., q_{t+s|t}^{m\prime} \right]'$$

where $q_{t+s|t}^i = \mathbb{E}_t \left( Z_{t+s} 1_{\xi_{t+s} = i} \right) = \mathbb{E}_t \left( Z_{t+s} | \xi_{t+s} = i \right) \pi_{t+s|t}^i$, where $\pi_{t+s|t}^i \equiv P \left( \xi_{t+s} = i | \mathbb{I}_t \right)$, we have

$$\mu_{t+s|t} = \mathbb{E}_t(Z_{t+s}) = w q_{t+s|t}, \tag{A3}$$

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where for $s \geq 1$, $q_{t+s|t}$ evolves as:

$$q_{t+s|t} = C\pi_{t+s|t} + \Omega q_{t+s-1|t} \tag{A4}$$

$$\pi_{t+s|t} = H\pi_{t+s-1|t} \tag{A5}$$

with $\pi_{t+s|t} = \left[\pi_{t+s|t}^1, ..., \pi_{t+s|t}^m\right]'$, $\Omega = bdiag\left(A_1, ..., A_m\right)\left(H \otimes I_n\right)$, and $\underset{mn \times m}{C} = bdiag\left(c_1, ..., c_m\right)$, where e.g., $c_1$ is the $n \times 1$ vector of constants in regime 1, $\otimes$ represents the Kronecker product and $bdiag$ is a matrix operator that takes a sequence of matrices and use them to construct a block diagonal matrix.

Similar formulas hold for the second moments. Before proceeding, let us define the vectorization operator $\varphi(X)$ that takes the matrix $X$ as an input and returns a column vector stacking the columns of the matrix $X$ on top of one another. We will also make use of the following result: $\varphi(X_1 X_2 X_3) = (X_3' \otimes X_1)\varphi(X_2)$.

Define the vector $n^2 m \times 1$ column vector $Q_t$ as:

$$Q_t = \left[Q_t^{1\prime}, ..., Q_t^{m\prime}\right]'$$

where the $n^2 \times 1$ vector $Q_t^i$ is given by $Q_t^i = \varphi\left[\mathbb{E}_0\left(Z_t Z_t' 1_{\xi_t=i}\right)\right]$. This implies that we can compute the vectorized matrix of second moments $M_t = \varphi\left[\mathbb{E}_0\left(Z_t Z_t'\right)\right]$ as:

$$M_t = \varphi\left[\mathbb{E}_0\left(Z_t Z_t'\right)\right] = \sum_{i=1}^m Q_t^i = WQ_t$$

where the matrix $W = [I_{n^2}, ..., I_{n^2}]$ is obtained placing side by side $m$ $n^2$-dimensional identity matrices. We can then state the following proposition:

**Proposition 2** *Consider a Markov-switching model whose law of motion can be described by (A2) and define $Q_t^i = \varphi\left[\mathbb{E}_0\left(Z_t Z_t' 1_{\xi_t=i}\right)\right]$, $q_t^i = \mathbb{E}_0\left[Z_t 1_{\xi_t=i}\right]$, and $\pi_t^i = P_0\left(\xi_t = i\right)$, for $i = 1...m$. Then $Q_t^j = \left[\widehat{cc}_j + \widehat{VV}_j \varphi\left[I_k\right]\right]\pi_t^j + \sum_{i=1}^m \left[\widehat{AA}_j Q_{t-1}^i + \widehat{DAC}_j q_{t-1}^i\right] h_{ji}$, where $\widehat{cc}_j = (c_j \otimes c_j)$, $\widehat{VV}_j = (V_j \otimes V_j)$, $\widehat{AA}_j = (A_j \otimes A_j)$, and $\widehat{DAC}_j = (A_j \otimes c_j) + (c_j \otimes A_j)$.*

It is then straightforward to compute the evolution of second moments conditional on the information available at a particular point in time. Suppose we are interested in $\mathbb{E}_t\left(Z_{t+s} Z_{t+s}'\right)$, i.e. the second moment of the vector $Z_{t+s}$ conditional on the information available at time $t$. If we define:

$$Q_{t+s|t} = \left[Q_{t+s|t}^{1\prime}, ..., Q_{t+s|t}^{m\prime}\right]'$$

where $Q_{t+s|t}^i = \varphi\left(\mathbb{E}_t\left(Z_{t+s} Z_{t+s}' 1_{\xi_{t+s}=i}\right)\right) = \varphi\left(\mathbb{E}_t\left(Z_{t+s} Z_{t+s}'|\xi_{t+s} = i\right)\right)\pi_{t+s|t}^i$, we obtain $\varphi\left(\mathbb{E}_t\left(Z_{t+s} Z_{t+s}'\right)\right) = WQ_{t+s|t}$. Using matrix algebra we obtain:

$$Q_{t+s|t} = \Xi Q_{t+s-1|t} + \widehat{DAC} q_{t+s-1|t} + \widehat{Vc}\pi_{t+s|t} \tag{A6}$$

$$q_{t+s|t} = C\pi_{t+s|t} + \Omega q_{t+s-1|t}, \quad \pi_{t+s|t} = H\pi_{t+s-1|t}. \tag{A7}$$

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where

$$\Xi = bdiag(\widehat{AA}_1,...,\widehat{AA}_m)(H \otimes I_{n^2}), \ \widehat{Vc} = \left[\widehat{VV} + \widehat{cc}\right], \ \widehat{cc} = bdiag(\widehat{cc}_1,...,\widehat{cc}_m),$$

$$\widehat{VV} = bdiag(\widehat{VV}_1\varphi\left[I_k\right],...,\widehat{VV}_m\varphi\left[I_k\right]), \ \widehat{DAC} = bdiag(\widehat{DAC}_1,...,\widehat{DAC}_m)(H \otimes I_n).$$

With the first and second moments at hand, it is then possible to compute the variance $s$ periods ahead conditional on the information available at time $t$:

$$\varphi\left[\mathbb{V}_t\left(Z_{t+s}\right)\right] = M_{t+s|t} - \varphi\left[\mu_{t+s|t}\mu'_{t+s|t}\right],\tag{A8}$$

where $M_{t+s|t} = \varphi\left(\mathbb{E}_t\left(Z_{t+s}Z'_{t+s}\right)\right) = \sum_{i=1}^{m} Q^i_{t+s|t} = WQ_{t+s|t}$.

To report estimates of (A3) and (A8) we proceed as follows. Note that $\mu_{t+s|t} = \mathbb{E}_t\left(Z_{t+s}\right) = wq_{t+s|t}$ and $M_{t+s|t}$ depend only on $q_{t+s|t}$ and $Q_{t+s|t}$. Furthermore we can express (A4)-(A5) and (A6)-(A7) in a compact form as

$$\widetilde{Q}_{t+s|t} = \widetilde{\Xi}^s\widetilde{Q}_{t|t} \text{ where } \widetilde{\Xi} = \left[\begin{array}{c|cc} \Xi & \widehat{DAC} & \widehat{Vc}H \\ \hline & \Omega & CH \\ & & H \end{array}\right],\tag{A9}$$

where $\widetilde{Q}_{t+s|t} = \left[Q'_{t+s|t}, q'_{t+s|t}, \pi'_{t+s|t}\right]'$. Armed with starting values $\widetilde{Q}_{t|t} = \left[Q'_{t|t}, q'_{t|t}, \pi'_{t|t}\right]'$ we can then compute (A3) and (A8) using (A9). To obtain $\pi'_{t|t}$ recall that we assume that $\mathbb{I}_t$ includes knowledge of the regime in place at time $t$, the data up to time $t$, $Z^t$, and the VAR parameters for each regime. Given that we assume knowledge of the current regime, $\pi^i_{t|t} \equiv P\left(\xi_t = i|\mathbb{I}_t\right)$ can only assume two values, 0 or 1. As a result $\pi'_{t|t}$ will be $(1,0)$ or $(0,1)$. As a result, and given $Z_t \in \mathbb{I}_t$, $q'_{t|t} = \left[q^{1'}_{t|t}, q^{2'}_{t|t}\right]'$ with $q^i_{t|t} \equiv \mathbb{E}_t\left(Z_t|\xi_t = i\right)\pi^i_{t|t}$, will be $\left[Z'_t \cdot 1, Z'_t \cdot 0\right]'$ or $\left[Z'_t \cdot 0, Z'_t \cdot 1\right]'$. Analogously, $Q'_{t|t} = \left[Q^{1'}_{t|t}, Q^{2'}_{t|t}\right]'$ with $Q^i_{t|t} \equiv \varphi\left(\mathbb{E}_t\left(Z_tZ'_t|\xi_t = i\right)\right)\pi^i_{t|t}$ will be $\left[\varphi\left(Z_tZ'_t \cdot 1\right)', \varphi\left(Z_tZ'_t \cdot 0\right)'\right]'$ or $\left[\varphi\left(Z_tZ'_t \cdot 0\right)', \varphi\left(Z_tZ'_t \cdot 1\right)'\right]'$.

## 7.1 Conditional Steady State

Consider a MS-VAR:

$$Z_t = c_{\xi_t} + A_{\xi_t}Z_{t-1} + V_{\xi_t}\varepsilon_t$$

where $Z_t$ is a column vector containing $n$ variables observable at time $t$ and $\xi_t = 1,...,m$, with $m$ the number of regimes, evolves following the transition matrix $H$. If the MS-VAR has more than one lag, the companion form can be used to recast the model as illustrated above.

The conditional steady state for the mean corresponds to the expected value for the vector $Z_t$ conditional on being in a particular regime. This is computed by imposing that a certain regime is in place forever:

$$\mathbb{E}_i\left(Z_t\right) = \overline{\mu}_i = \left(I_n - A_i\right)^{-1} c_i \tag{A10}$$

where $I_n$ is an identity matrix with the appropriate size. Note that unless the VAR coefficients imply very slow moving dynamics, after a switch from regime $j$ to regime $i$, the variables of the VAR will converge (in expectation) to $\mathbb{E}_i\left(Z_t\right)$ over a finite horizon. If there are no further switches, we can then expect the variables to fluctuate around $\mathbb{E}_i\left(Z_t\right)$. Therefore, the conditional steady states for the mean can also be thought as the values to which the variables converge if regime $i$ is in place for a long enough period of time.

The conditional steady state for the standard deviation corresponds to the standard deviation for the vector $Z_t$ conditional on being in a particular regime. The conditional standard deviations for the elements in $Z_t$ are computed by taking the square root of the main diagonal elements of the covariance matrix $\mathbb{V}_i\left(Z_t\right)$ obtained imposing that a certain regime is in place forever:

$$\varphi\left(\mathbb{V}_i\left(Z_t\right)\right) = \left(I_{n^2} - A_i \otimes A_i\right)^{-1} \varphi\left(V_{\xi_t} V_{\xi_t}'\right) \tag{A11}$$

where $I_{n^2}$ is an identity matrix with the appropriate size, $\otimes$ denotes the Kronecker product, and the vectorization operator $\varphi\left(X\right)$ takes a matrix $X$ as an input and returns a column vector stacking the columns of the matrix $X$ on top of one another.

## 7.2  Book-to-Market Ratio

We use the methods and assumptions of the previous subsection to obtain the present value decomposition of the book to market ratio. Consider an MS-VAR:

$$Z_t = c_{\xi_t} + A_{\xi_t} Z_{t-1} + V_{\xi_t} \varepsilon_t$$

where $Z_t$ is a column vector containing $n$ variables observable at time $t$ and $\xi_t = 1, ..., m$, with $m$ the number of regimes, evolves following the transition matrix $H$. If the MS-VAR has more than one lag, the companion form can be used to recast the model as illustrated above.

Define the column vectors $q_t$ and $\pi_t$:

$$q_t = \left[q_t^{1\prime}, ..., q_t^{m\prime}\right]', q_t^i = \mathbb{E}_0\left(Z_t 1_{\xi_t=i}\right), \pi_t = \left[\pi_t^1, ..., \pi_t^m\right]',$$

where $\pi_t^i = P_0\left(\xi_t = i\right)$ and $1_{\xi_t=i}$ is an indicator variable that is equal to 1 when regime $i$ is in place and zero otherwise. The law of motion for $\widetilde{q}_t = [q_t', \pi_t']'$ is then given by

$$\underbrace{\begin{bmatrix} q_t \\ \pi_t \end{bmatrix}}_{\widetilde{q}_t} = \underbrace{\begin{bmatrix} \Omega & CH \\ & H \end{bmatrix}}_{\widetilde{\Omega}} \begin{bmatrix} q_{t-1} \\ \pi_{t-1} \end{bmatrix}$$

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where $\pi_t = [\pi_{1,t}, ..., \pi_{m,t}]'$, $\Omega = bdiag(A_1, ..., A_m) H$, and $C = bdiag(c_1, ..., c_m)$. Recall that:

$$\mathbb{E}_0(Z_t) = \sum_{i=1}^{m} q_t^i = wq_t, \ w = \left[ \underbrace{I_n, ..., I_n}_{m} \right]$$

To compute the present value decomposition of the book-to-market ratio, define:

$$q_{t+s|t}^i = \mathbb{E}_t \left( Z_{t+s} 1_{\xi_{t+s}=i} \right) = \mathbb{E} \left( Z_{t+s} 1_{\xi_{t+s}=i} | \mathbb{I}_t \right)$$
$$1_x' = [0, ...1, ...0, 0, 0]', \ mn = m * n$$

where $\mathbb{I}_t$ contains all the information that agents have at time $t$, including the probability of being in one of the $m$ regimes. Note that $q_{t|t}^i = Z_t \pi_t^i$.

Now consider the formula from Vuolteenaho (1999):

$$\theta_t = \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t r_{t+1+j} + \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t f_{t+1+j} - \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t e_{t+1+j}^*$$

Given that our goal is to assess if assets with different risk profiles are affected differently by the breaks in the long-term interest rates, we are going to focus on the *difference* between the book-to-market ratios. Specifically, given two portfolios $x$ and $y$, we are interested in how the difference in their book-to-market ratios, $\theta_{x,t} - \theta_{y,t}$, varies across the two regimes:

$$\underbrace{\theta_{x,t} - \theta_{y,t}}_{\text{Spread in BM ratios}} = \underbrace{\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( r_{x,t+1+j} - r_{y,t+1+j} \right)}_{\text{PDV of the difference in expected excess returns}} - \underbrace{\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( e_{x,t+1+j}^* - e_{y,t+1+j}^* \right)}_{\text{PDV of the difference in expected earnings}}$$

If then we want to correct the spread in BM ratios by taking into account expected earnings, we have:

$$\underbrace{\theta_{x,t} - \theta_{y,t} + \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( e_{x,t+1+j}^* - e_{y,t+1+j}^* \right)}_{\text{Spread in BM ratios corrected for earnings}} = \underbrace{\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( r_{x,t+1+j} - r_{y,t+1+j} \right)}_{\text{PDV of the expected spread in excess returns}} \quad \text{(A12)}$$

Suppose that we have estimated a MS-VAR that includes the spread in excess returns, $r_{xy,t} \equiv r_{x,t} - r_{y,t}$. Then the right hand side of (A12) can be computed as:

$$\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( r_{xy,t+1+j} \right) = \sum_{j=0}^{\infty} \rho^j 1_{r_{xy}}' wq_{t+1+j|t}$$
$$= 1_{r_{xy}}' w (I - \rho\Omega)^{-1} \left[ \Omega q_{t|t} + C (I - \rho H)^{-1} H\pi_{t|t} \right].$$

Therefore, we have:

$$\widetilde{\theta}_{xy,t} \equiv \underbrace{\widetilde{\theta}_{x,t} - \widetilde{\theta}_{y,t} + \sum_{j=0}^{\infty} \rho^j \mathbb{E}_t \left( e_{x,t+1+j}^* - e_{y,t+1+j}^* \right)}_{\text{Spread in BM ratios corrected for earnings}} = 1_{r_{xy}}' w (I - \rho\Omega)^{-1} \left[ \Omega q_{t|t} + C (I - \rho H)^{-1} H\pi_{t|t} \right]$$

$$\text{(A13)}$$

where we have used $\widetilde{\theta}_{xy,t}$ to define the spread in BM ratios corrected for earnings.

Similar formulas are used to compute risk premia for the individual portfolios. The premium for a portfolio $z$ coincides with the present discounted value of its excess returns:

$$\underbrace{premia_{z,t}}_{\text{Premia}} \equiv \underbrace{\sum_{j=0}^{\infty} \rho^j \mathbb{E}_t\left(r_{z,t+1+j}\right)}_{\text{PDV of excess returns}} = 1'_{r_z} w\left(I - \rho\Omega\right)^{-1}\left[\Omega q_{t|t} + C\left(I - \rho H\right)^{-1} H\pi_{t|t}\right], \quad \text{(A14)}$$

where $1'_{r_z}$ is a vector used to extract the PDV of excess returns from a vector containing the PDV of all variables included in the VAR.

**Regime Average** We also compute the *regime average* value of $\widetilde{\theta}_{xy,t}$. The regime average is defined as:

$$\overline{\widetilde{\theta}}_{xy}^{i} \equiv 1'_{r_{xy}} w\left(I - \rho\Omega\right)^{-1}\left[\Omega\overline{q}_i + C\left(I - \rho H\right)^{-1} H\overline{\pi}_i\right]$$

where $\overline{\pi}_i = 1_i$ and $\overline{q}_i \equiv [0, ..., \overline{\mu}_i, ..., 0]$ is a column vector that contains the conditional steady state of for the mean value of $Z_t$ conditional on being in regime $i$, i.e., $\mathbb{E}_i\left(Z_t\right) = \overline{\mu}_i = \left(I_n - A_i\right)^{-1} c_i$, and zero otherwise. Recall that the conditional steady state, $\overline{\mu}_i$, is a vector that contains the expected value of $Z_t$ conditional on being in regime $i$. Therefore, the vector captures the values to which the variables of the VAR converge if regime $i$ is in place for a prolonged period of time. Note that $\overline{\widetilde{\theta}}_{xy}^{i}$ is computed by conditioning on the economy being initially at $Z_t = \overline{\mu}_i$ and in regime $i$, but taking into account that there might be regime changes in the future. Therefore, we can also think about $\overline{\widetilde{\theta}}_{xy}^{i}$ as the expected value of the PDV of excess returns, $\widetilde{\theta}_{xy,t}$, conditioning on being in regime $i$ today and on the variables of the VAR being equal to the conditional steady state mean values for regime $i$. Formally:

$$\overline{\widetilde{\theta}}_{xy}^{i} = \mathbb{E}\left(\widetilde{\theta}_{xy,t}|\xi_t = i, Z_t = \overline{\mu}_i\right). \quad \text{(A15)}$$

These are the values around which we expect $\widetilde{\theta}_{xy,t}$ to fluctuate as the VAR variables $Z_t$ fluctuate around $\overline{\mu}_i$.

Similarly, we can compute the *regime average* value of risk premia for an individual portfolio $z$, $premia_{z,t}$:

$$\overline{premia}_{z}^{i} \equiv 1'_{r_z} w\left(I - \rho\Omega\right)^{-1}\left[\Omega\overline{q}_i + C\left(I - \rho H\right)^{-1} H\overline{\pi}_i\right]. \quad \text{(A16)}$$

Formulas (A13), (A14), (A15), and (A16) are used in the paper to produce Figure 10 and Tables 10 and 11. For each draw of the VAR parameters from the posterior distribution, we compute the evolution of the PDV of the spread in excess returns over time between high and low return premia portfolios, $\widetilde{\theta}_{xy,t}$ and individual portfolio $premia_{z,t}$, by using (A13) and (A14). Thus, we obtain a posterior distribution for $\widetilde{\theta}_{xy,t}$ and $premia_{z,t}$. The medians of these posterior distributions are reported as the blue solid lines in Figure 10. Similarly, for each draw of the VAR coefficients, we compute $\overline{\widetilde{\theta}}_{xy}^{i}$ and the difference in the PDV between the two regimes $\overline{\widetilde{\theta}}_{xy}^{1} - \overline{\widetilde{\theta}}_{xy}^{2}$. Thus, we obtain a posterior distribution for $\overline{\widetilde{\theta}}_{xy}^{i}$ and for the difference $\overline{\widetilde{\theta}}_{xy}^{1} - \overline{\widetilde{\theta}}_{xy}^{2}$. The

medians of the distribution of $\overline{\tilde{\theta}}_{xy}^i$ and $\overline{premia}_z^i$ for $i = 1, 2$, are reported in Figure 10 (red dashed line). Table 10 reports the median and the 68% posterior credible sets both for the distribution of $\overline{\tilde{\theta}}_{xy}^i$, for $i = 1, 2$, and for the difference in these across regimes, $\overline{\tilde{\theta}}_{xy}^1 - \overline{\tilde{\theta}}_{xy}^2$. Finally, Table 11 reports the percentage of draws for which $\overline{\tilde{\theta}}_{xy}^1 - \overline{\tilde{\theta}}_{xy}^2 > 0$ and $\overline{premia}_z^1 - \overline{premia}_z^2 > 0$ as the probability that risk premia are lower in the high asset valuation/low interest rate regime than they are in the low asset valuation/high interest rate regime.

| Variable | NSE | RNE | Variable | NSE | RNE | Variable | NSE | RNE |
|---|---|---|---|---|---|---|---|---|
| $\alpha_1$ | 0.000131 | 1 | $\Delta a_{t+1}$ | 0.000263 | 1 | $\Delta y_{t-4}$ | 0.000526 | 1 |
| $\alpha_2$ | 0.000131 | 1 | $\Delta y_{t+1}$ | 0.00053 | 1 | $\Delta a_{t+4}$ | 0.000256 | 1 |
| $\beta_a$ | 0.000074 | 1 | $\Delta a_{t-2}$ | 0.000261 | 1 | $\Delta y_{t+4}$ | 0.000521 | 1 |
| $\beta_y$ | 0.000085 | 1 | $\Delta y_{t-2}$ | 0.000572 | 1 | $\Delta a_{t-5}$ | 0.000264 | 1 |
| $\sigma$ | 0 | 1 | $\Delta a_{t+2}$ | 0.000258 | 1 | $\Delta y_{t-5}$ | 0.000524 | 1 |
| $H_{11}^{\alpha}$ | 0.000069 | 1 | $\Delta y_{t+2}$ | 0.000547 | 1 | $\Delta a_{t+5}$ | 0.000252 | 1 |
| $H_{22}^{\alpha}$ | 0.000053 | 1 | $\Delta a_{t-3}$ | 0.000278 | 1 | $\Delta y_{t+5}$ | 0.000534 | 1 |
| $\Delta a_t$ | 0.000263 | 1 | $\Delta y_{t-3}$ | 0.000632 | 1 | $\Delta a_{t-6}$ | 0.000275 | 1 |
| $\Delta y_t$ | 0.000529 | 1 | $\Delta a_{t+3}$ | 0.000255 | 1 | $\Delta y_{t-6}$ | 0.000518 | 1 |
| $\Delta a_{t-1}$ | 0.000252 | 1 | $\Delta y_{t+3}$ | 0.000537 | 1 | $\Delta a_{t+6}$ | 0.000238 | 1 |
| $\Delta y_{t-1}$ | 0.000521 | 1 | $\Delta a_{t-4}$ | 0.000259 | 1 | $\Delta y_{t+6}$ | 0.000525 | 1 |

**Table A.3:** The table reports the numerical standard error (NSE) and the relative numerical efficiency (RNE) computed based on Geweke (1992). Values for NSE close to zero and values for RSE close to 1 are indicative of convergence. The sample is quarterly and spans the period 1952:Q1 to 2013:Q3.