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A Single-Factor Consumption-Based Asset Pricing Model

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Abstract

We propose a single-factor asset pricing model based on an indicator function of consumption growth being less than its endogenous certainty equivalent. This certainty equivalent is derived from generalized disappointment-aversion preferences, and it is located approximately 1 standard deviation below the conditional mean of consumption growth. Our single-factor model can explain the cross section of expected returns for size, value, reversal, profitability, and investment portfolios at least as well as the Fama–French multifactor models. Our results show strong empirical support for asymmetric preferences and question the effectiveness of the smooth utility framework, which is traditionally used in consumption-based asset pricing.

I. Introduction

Despite their universal importance, asset pricing models are usually characterized by a puzzling contradiction. On the one hand, most of the empirical factor models that can successfully explain risk premia do not provide strong guidance with respect to the deep economic mechanism that drives expected returns.

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On the other hand, many models that are motivated by economic theory have limited empirical success.

For instance, according to the standard consumption-based capital asset pricing model (CCAPM) of Breeden (1979), expected returns across assets should be explained by their exposure to aggregate consumption risk. Specifically, Breeden shows that Merton's (1973) intertemporal capital asset pricing model (ICAPM) pricing equation can be collapsed into a single-beta representation, where the expected excess return on any security is proportional to its covariance with aggregate consumption growth alone. Nevertheless, a number of studies have questioned the ability of consumption risk to explain equity premia and the cross section of expected returns (e.g., Mehra and Prescott (1985), Liu, Whited, and Zhang (2009)). Further, recent efforts to introduce alternative measures of consumption risk (e.g., Parker and Julliard (2005), Jagannathan and Wang (2007), and Savov (2011)) have been undermined by the implausibly high degree of risk aversion that these models still imply for the representative agent.

To address the poor performance of single-factor models like the capital asset pricing model (CAPM) or the CCAPM, empirical asset pricing models tend to employ an increasing number of return-generated factors (see the recent contributions of Fama and French (2015) and Hou, Xue, and Zhang (2016)). However, in most cases, these multifactor models do not provide convincing arguments regarding the economic underpinnings of the proposed pricing factors, let alone data-mining concerns (see Harvey, Liu, and Zhu (2015)).

Motivated by the empirical shortcomings of the traditional consumption model and the upward trend in the number of pricing factors, we propose a single-factor, consumption-based model featuring disappointment aversion, but second-order risk neutrality, to explain the cross section of expected stock returns. The only pricing factor in our model is an indicator function of consumption growth being less than its certainty equivalent. This certainty equivalent is derived from generalized disappointment-aversion preferences (Gul (1991), Routledge and Zin (2010)), and it is located approximately 1 standard deviation below the conditional mean of consumption growth. Using this single-factor model, we show that downside consumption risk alone, as proxied by the disappointment indicator, can rationalize both the level and the cross-sectional dispersion of expected returns.

The starting point for our single-factor model is disappointment aversion, which was first introduced by Gul (1991) and subsequently generalized by Routledge and Zin (2010). Under this axiomatic framework, investor utility over stochastic consumption exhibits three features: i) it is defined over deviations from a reference point, ii) it is steeper for losses than for gains, and iii) the reference point for gains and losses is based on the certainty equivalent of consumption.

These characteristics imply that disappointment-aversion preferences are described by utility functions with a kink, exhibiting first-order risk aversion (Segal and Spivak (1990)), contrary to standard preferences specifications that employ smooth utility functions characterized by second-order risk aversion, such as the constant relative risk aversion (CRRA) or constant absolute risk aversion (CARA) utility functions. Moreover, in the disappointment-aversion framework, the location of the kink is endogenously determined rather than being exogenously

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imposed in an ad hoc fashion, as is the case with prospect theory specifications (see Barberis, Huang, and Santos (2001)).

Despite its theoretical tractability, the implementation of the disappointment model is quite challenging. Specifically, in the generalized disappointment model of Routledge and Zin (2010), preferences are nonseparable across time, and hence the resulting stochastic discount factor (SDF) is a function of consumption growth as well as lifetime utility, which is unobservable. The presence of unobservable lifetime utility in the disappointment-aversion SDF impedes its empirical estimation. Following Delikouras (2017), we circumvent this problem by assuming that consumption growth is predictable and homoscedastic. Thus, we are able to derive explicit solutions for lifetime utility and the disappointment-aversion SDF in terms of observable consumption growth alone.

In contrast to the highly nonlinear model proposed by Delikouras (2017), we propose a model where the representative investor is disappointment averse but second-order risk neutral. In other words, the investor's preferences are characterized by a piecewise linear function with a kink. As a result, our model is quite simple and consists of a single asset pricing factor: the indicator of consumption growth being less than its certainty equivalent. The implication of our model is that risky assets yield premia as compensation for being exposed to downside consumption risk only; consumption growth risk per se is not priced.

For our main tests, we examine the fit of the proposed asset pricing model using annual returns of the size, value, reversal, profitability, and investment portfolios. These portfolios are the basis for the most frequently used pricing factors, such as the high-minus-low (HML), small-minus-big (SMB), conservative-minus-aggressive (CMA), and robust-minus-weak (RMW) factors of Fama and French (1993), (2015). The results are striking. The disappointment indicator can fit the cross sections of expected returns at least as accurately as the Fama–French 3-factor model (FF3), and its performance is comparable to the Fama–French 5-factor specification (FF5). In particular, the cross-sectional R^2s for our single-factor model are 90% for the size/book-to-market portfolios, 88% for the size/profitability portfolios, 74% for the size/investment portfolios, and 87% for the long-term reversal portfolios. The corresponding R^2s for the FF3 are 85%, 80%, 73%, and 93%, whereas for the FF5, the R^2s are 82%, 88%, 81%, and 97%, respectively. A fortiori, our single-factor model can also explain the premia of the five Fama and French (2015) factors.

Our tests also provide insights into the plausibility of the preference parameters of consumption-based models. Specifically, the key parameter in the proposed model is the disappointment-aversion coefficient, which measures the asymmetry in investor preferences over gains and losses. Our estimates of this coefficient are approximately 4 at the portfolio level, implying that investors penalize losses during disappointment events 5 times more than losses during normal times. These values are very close to the one employed by Ang, Bekaert, and Liu (2005) to explain the historical equity premium, and they yield very realistic risk-taking behavior in Rabin (2000) games (see Ang et al. (2005)).

Moreover, our estimates for the disappointment-aversion coefficient remain fairly stable across subperiods, test assets, and return frequencies. To the contrary, the standard CCAPM or alternative definitions of consumption risk, such as the

one introduced by Jagannathan and Wang (2007), yield quite high estimates of second-order risk aversion, rendering the representative investor implausibly risk averse over modest or large-stake gambles (see Rabin (2000)). This limitation is exacerbated with higher-frequency consumption data.

The results are similar when we alternatively consider a joint cross section of equity, corporate bond, Treasury bond, and equity index option portfolios. At the annual frequency, the cross-sectional R^2 of our single-factor model is 72%. More importantly, our model yields relatively low pricing errors for these alternative asset classes, both individually and jointly, with very similar and plausible disappointment-aversion coefficient estimates. To the contrary, the Fama–French (1993), (2015) models can only fit well each of these cross sections with factor coefficient estimates whose signs and magnitudes are strikingly inconsistent.

Next, we assess the importance of the generalized disappointment threshold. To this end, we compare the empirical fit of our single-factor model, which is derived from the generalized disappointment-aversion specification of Routledge and Zin (2010), to that of a restricted version based on the original disappointment framework of Gul (1991). We find that in the sample of equity portfolios, the two models exhibit similar cross-sectional performance. In contrast, when we consider the sample of corporate bond, Treasury bond, and equity index option portfolios, the empirical fit of the generalized disappointment model is better than that of its restricted version.

We conclude our analysis by showing that the proposed model does not suffer from weak identification, and its empirical success is neither spuriously driven by the use of an indicator factor nor is particularly sensitive to the marginal characterization of some consumption growth observations as disappointing or not.

Overall, this study contributes to various strands of the asset pricing literature. We contribute to the literature on disappointment aversion, which has considerably grown following the works of Ang et al. (2005) and Routledge and Zin (2010).[1] Routledge and Zin (2010) and Bonomo, Garcia, Meddahi, and Tédongap (2011) calibrate consumption-based models with disappointment aversion to explain the equity premium but do not provide results for the cross section of expected returns.

Ostrovnaya, Routledge, and Zin (2006) and Farago and Tédongap (2018) conduct cross-sectional tests based on disappointment aversion, but they substitute out consumption, and they propose multifactor pricing models based on stock market returns. In contrast, our single-factor model is expressed in terms of observable consumption growth alone, and we show that the disappointment consumption indicator is sufficient to explain the cross section of expected returns. This is particularly important because, unlike the aforementioned models, our specification provides a direct link between the macroeconomy and asset prices.

Our work is also related to that of Delikouras (2017), who incorporates disappointment aversion in Epstein–Zin preferences and derives a highly nonlinear SDF. To the contrary, assuming a disappointment-averse but second-order

[1]Choi, Fisman, Gale, and Kariv (2007) and Dahlquist, Farago, and Tédongap (2017) examine portfolio choices, whereas Gill and Prowse (2012) focus on effort provision. Schreindorfer (2014) uses disappointment aversion to price put options.

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risk-neutral representative investor, we introduce a much simpler SDF, which is linear in the price of disappointment risk. This innovation facilitates the estimation of our model, allowing us to conduct standard asset pricing tests using large cross sections and to contribute to the ongoing debate in the literature regarding the theoretical underpinnings of multifactor models.

Our results also contribute to the consumption-based asset pricing literature. Specifically, we show that a nonlinear transformation of the consumption growth process, which derives from generalized disappointment-aversion preferences, can address the empirical shortcomings of the standard consumption model. Moreover, whereas Boguth and Kuehn (2013) argue that heteroscedasticity in consumption growth substantially improves the fit of the consumption model, we show that our SDF, which for simplicity is derived under the assumption of homoscedasticity, still yields a very good fit for the most commonly used portfolio cross sections. Finally, the empirical success of the proposed single-factor model provides support for the theoretical argument of Breeden (1979) that premia should be proportional to covariances with a function of aggregate consumption, but we crucially show that it is downside consumption risk that is priced.

Collectively, our results question the effectiveness of the standard framework of smooth utility functions that is commonly used in consumption-based asset pricing. In contrast, we argue in favor of first-order risk aversion and conclude that risk premia are, to a large extent, compensation for exposure to disappointment consumption risk.

II. The Single-Factor GDA-I SDF

In this section, we introduce our single-factor asset pricing model, termed the generalized disappointment aversion indicator (GDA-I). Our starting point is the GDA intertemporal SDF of Routledge and Zin (2010):

$$M_{t+1}^{\text{GDA}} = \beta\left(\frac{C_{t+1}}{C_t}\right)^{\rho-1}\left[\frac{V_{t+1}}{\mu_t(V_{t+1})}\right]^{\alpha-\rho}\left[\frac{1+\theta\mathbf{1}\{V_{t+1}\leq\delta\mu_t(V_{t+1})\}}{1-\theta(\delta^{\alpha}-1)\mathbf{1}\{\delta>1\}+\theta\delta^{\alpha}\mathbb{E}_t[\mathbf{1}\{V_{t+1}\leq\delta\mu_t(V_{t+1})\}]}\right], \tag{1}$$

$$\text{with } \mu_t(V_{t+1}) = \mathbb{E}_t\left[\frac{V_{t+1}^{\alpha}\,(1+\theta\mathbf{1}\{V_{t+1}\leq\delta\mu_t(V_{t+1})\})}{1-\theta(\delta^{\alpha}-1)\mathbf{1}\{\delta>1\}+\theta\delta^{\alpha}\mathbb{E}_t[\mathbf{1}\{V_{t+1}\leq\delta\mu_t(V_{t+1})\}]}\right]^{1/\alpha}. \tag{2}$$

This SDF adjusts expected values by taking into account investor preferences over the timing, risk, and disappointment of stochastic payoffs. $\mu_t(V_{t+1})$ is the GDA certainty equivalent for lifetime utility V_{t+1}, and $\mathbf{1}\{V_{t+1}\leq\delta\mu_t(V_{t+1})\}$ is the disappointment indicator. The parameter β is the rate of time preference, ρ determines the elasticity of intertemporal substitution (EIS $=1/(1-\rho)$), and α is the second-order risk-aversion parameter, which determines the piece-wise concavity of lifetime utility.

The disappointment-aversion (DA) coefficient θ is the novel parameter in the GDA SDF. This parameter determines the asymmetry in investor preferences

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around the disappointment threshold. When θ is positive, a \$1 loss in consumption during disappointment periods hurts approximately $1+\theta$ times more than a \$1 loss in consumption during normal times.[2] When θ is 0, we obtain the standard Epstein–Zin (1989) framework. In the original framework of Gul (1991) and in Delikouras (2017), disappointment events take place when lifetime utility falls below its certainty equivalent: $\mathbf{1}\{V_{t+1} \leq \mu_t(V_{t+1})\}$. In this study, however, we follow the GDA framework of Routledge and Zin (2010), where the disappointment threshold is a multiple of the certainty equivalent (i.e., $\mathbf{1}\{V_{t+1} \leq \delta\mu_t(V_{t+1})\}$). Thus, in the GDA model, the threshold for disappointment is also determined by the positive GDA parameter δ.

According to the expression in equation (1), the GDA SDF of Routledge and Zin (2010) is a function of the observable consumption growth and the unobservable lifetime utility because investor preferences are nonseparable across time. To substitute the unobservable lifetime utility out of the expression for the GDA SDF, we need to impose additional structure on the consumption growth process. In particular, we follow the consumption growth literature (see, e.g., Mehra and Prescott (1985), Routledge and Zin (2010)) and assume that consumption growth (Δc_t) is an autoregressive (AR(1)) process with normal shocks and constant volatility:

$$\Delta c_{t+1} \quad = \quad \mu_c(1-\phi_c) + \phi_c \Delta c_t + \sqrt{1-\phi_c^2}\sigma_c \epsilon_{c,t+1}.$$

The parameters μ_c, σ_c^2, and ϕ_c are the unconditional mean, variance, and first-order autocorrelation for consumption growth, respectively, whereas $\epsilon_{c,t+1}$ represents independent and identically distributed (IID) N(0, 1) shocks.

The main modeling innovation of our study is to assume that the representative investor is second-order risk neutral with time-separable preferences (i.e., $\alpha = \rho = 1$ in equation (1)). In other words, we assume that the utility function of the representative investor is piece-wise linear with a kink, the location of which is endogenously determined. The assumption that ρ equals 1 implies that in a deterministic world, intertemporal substitution is perfectly elastic. However, in a stochastic environment with nonseparable disappointment-aversion preferences, the effective EIS and risk-aversion coefficients depend on all four preference parameters: α, θ, ρ, and δ (see Routledge and Zin (2010)). Thus, in a stochastic world, the assumption that ρ is 1 does not necessarily imply an infinite EIS.

Moreover, it should be noted that this study examines the determination of unconditional asset premia, rather than interest rate dynamics, and hence the assumption $\rho = 1$ is not particularly restrictive. In fact, the assumption $\alpha = \rho = 1$ is motivated by the goal of this study, which is to show that a single-factor specification with disappointment aversion alone can sufficiently explain the cross section of expected returns. Allowing α and ρ to be free parameters would only increase the flexibility of our disappointment model.

Given these preferences assumptions, we introduce a very simple SDF that depends only on a disappointment consumption indicator. Specifically, in Section A1 of the Supplementary Material, we show that for $\alpha = \rho = 1$, we

[2]This behavior is consistent with a growing body of evidence documenting that disappointment risk is priced both at the individual and at the aggregate level (see Epstein and Zin 1990, (2001), Choi et al. (2007), Routledge and Zin (2010), and Gill and Prowse (2012)).

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can recast the GDA SDF of equation (1) as

$$(3)\quad M_{t+1}^{\text{GDA}} = \frac{\beta\left(1+\theta\mathbf{1}\left\{\Delta c_{t+1}\leq\mu_c(1-\phi_c)+\phi_c\Delta c_t+d_2\sqrt{1-\phi_c^2}\sigma_c\right\}\right)}{1-\theta(\delta-1)\mathbf{1}\{\delta>1\}+\theta\delta\mathbb{E}_t[\mathbf{1}\{\Delta c_{t+1}\leq\mu_c(1-\phi_c)+\phi_c\Delta c_t+d_2\sqrt{1-\phi_c^2}\sigma_c\}]},$$

where d_2 is the solution to a fixed-point problem (equation (A.14) in the Supplementary Material).

Motivated by the aforementioned structural GDA SDF and focusing on excess returns in our empirical analysis, we propose a linear SDF with respect to the price of disappointment risk, which reads, up to a multiplicative scalar term, as

$$(4)\quad M_{t+1}^{\text{GDA-I}} = \tilde{\theta}\mathbf{1}\left\{\Delta c_{t+1}\leq\mu_c(1-\phi_c)+\phi_c\Delta c_t+d_2\sqrt{1-\phi_c^2}\sigma_c\right\},$$

where $\tilde{\theta}$ denotes the DA coefficient (i.e., the price of disappointment risk in this linear model). We term the aforementioned SDF as GDA-I for two reasons. First, it is a special case of the GDA SDF of Routledge and Zin (2010). Second, it corresponds to a single-factor model that depends only on an indicator function.

Contrary to the general GDA SDF in equation (1) and the traditional CCAPM specification, consumption growth risk per se is not priced in our GDA-I model due to the second-order risk-neutrality assumption. Specifically, the GDA-I SDF in equation (4) indicates that risky assets should compensate investors for a single source of systematic risk, namely, disappointment consumption risk, which is captured by the disappointment indicator $\mathbf{1}\{\}$. In other words, the proposed GDA-I SDF consists of a single pricing factor, the indicator of consumption growth being less than its generalized certainty equivalent $\mu_t^{\text{GDA-I}}(\Delta c_{t+1})$, which is defined as

$$(5)\quad \mu_t^{\text{GDA-I}}(\Delta c_{t+1}) = \mu_c(1-\phi_c)+\phi_c\Delta c_t+d_2\sqrt{1-\phi_c^2}\sigma_c.$$

The disappointment indicator in equation (4) is scaled by the DA parameter $\tilde{\theta}$, which captures the price of disappointment risk.[3] The coefficient d_2 in equations (4) and (5) is the disappointment threshold coefficient. In principle, this parameter is an implicit function of consumption growth moments, the GDA coefficient δ, and the DA parameter θ (see equation (A.14) in the Supplementary Material). Because δ is a free parameter, to facilitate estimation and maintain the linearity of the GDA-I model with respect to the price of disappointment risk, we directly estimate d_2 via the generalized method of moments (GMM).

Essentially, the GDA-I SDF in equation (4) follows a bimodal distribution, exhibiting a switching-type behavior even though the consumption growth process does not (see Epstein and Zin (2001) for a related discussion). Testing the goodness of fit of this parsimonious single-factor SDF, we basically test whether disappointment consumption risk alone suffices to price the cross section of expected returns. This is an important innovation relative to existing

[3] Because the GDA-I model in equation (4) is a stripped-down version of the structural GDA discount factor in equation (3), the DA coefficient $\tilde{\theta}$ we estimate from the GDA-I model does not coincide with the structural DA coefficient θ. However, the latter could, in principle, be recovered from our $\tilde{\theta}$ estimate.

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disappointment-aversion models and a more challenging task relative to the GDA SDFs of Routledge and Zin (2010) and Delikouras (2017) because our preferences forego the flexibility that unrestricted second-order risk aversion and EIS parameters would allow.

A. Alternative Asset Pricing Models

In addition to the GDA-I model, we estimate the following set of asset pricing models:

$$M_{t+1}^{\text{CCAPM}} = -\tilde{\alpha}\,\Delta c_{t+1}, \tag{6}$$

$$M_{t+1}^{\text{CAPM}} = -b_m R_{m,t+1}^{x}, \tag{7}$$

$$M_{t+1}^{\text{FF3}} = -b_m R_{m,t+1}^{x} - b_{\text{SMB}} R_{\text{SMB},t+1} - b_{\text{HML}} R_{\text{HML},t+1}, \tag{8}$$

$$M_{t+1}^{\text{FF5}} = -b_m R_{m,t+1}^{x} - b_{\text{SMB}} R_{\text{SMB},t+1} - b_{\text{HML}} R_{\text{HML},t+1} - b_{\text{RMW}} R_{\text{RMW},t+1} - b_{\text{CMA}} R_{\text{CMA},t+1}, \tag{9}$$

$$M_{t+1}^{\text{NBER}} = b_{\text{NBER}} \mathbf{1}\{\text{more than 4 NBER recession months in year } t+1\}. \tag{10}$$

The discount factor in equation (6) is the linearized CCAPM with CRRA preferences. The coefficient $\tilde{\alpha}$ is the second-order risk-aversion parameter that captures the price of consumption risk in the linearized CCAPM. The CCAPM represents the traditional view in consumption-based asset pricing, according to which investors are second-order risk averse but disappointment risk neutral. The SDF M_{t+1}^{CAPM} corresponds to the standard CAPM, whereas M_{t+1}^{FF3} and M_{t+1}^{FF5} correspond to the Fama and French (1993), (2015) 3- and 5-factor models, respectively. Finally, M_{t+1}^{NBER} is a discount factor based on the recession indicator of the National Bureau of Economic Research (NBER). This indicator takes the value 1 when there are more than 4 NBER recession months in year $t+1$, and 0 otherwise.[4] The parameter b_{NBER} in equation (10) is the price of NBER recession risk.

The choice of these models is warranted by the fact that, with the exception of the NBER model, these specifications are the workhorses of empirical asset pricing. Additionally, we use the NBER model to show that disappointment consumption events do not simply capture NBER recessions, but they also have distinct asset pricing implications.

III. Data and Estimation Methodology

A. Data

The aim of our empirical analysis is to examine whether the GDA-I model can explain stylized facts in the cross section of equity returns. To this end, we consider the following portfolio cross sections:[5]

[4] Because NBER recessions are defined on a monthly basis, we create an annual NBER indicator by considering a 4-month cutoff for the number of NBER recession months in a year. We have considered alternative cutoffs for the NBER indicator. The 4-month cutoff yields the best cross-sectional fit. Based on this cutoff, the NBER recession years are 1937, 1938, 1945, 1949, 1953, 1954, 1960, 1970, 1974, 1980, 1981, 1982, 1990, 2001, 2008, and 2009.

[5] In Section A12 of the Supplementary Material, we report additional results at the stock level.

1) Six and 25 portfolios sorted on size/book-to-market (SIZE/BM). These portfolios capture the size and value premia, which are reflected in the SMB and HML factors of the FF3.[6]
2) Twenty-five portfolios sorted on size/operating profitability (SIZE/OP). These portfolios capture the profitability premia, which are reflected in the RMW factor of the FF5.
3) Twenty-five portfolios sorted on size/investment (SIZE/INV). These portfolios capture the investment premia, which are reflected in the CMA factor of the FF5.
4) Ten portfolios sorted on long-term reversal (LTR). These portfolios capture the long-term reversal premium documented by Jegadeesh and Titman (1993).

We use these sets of portfolios for two reasons. First, these sets of equity portfolios constitute the basis for a number of return-generated factors that are commonly used in the empirical asset pricing literature, such as the SMB, HML, CMA, and RMW factors in Fama and French (1993), (2015). Second, as shown by Harvey et al. (2015) and Hou et al. (2016), these portfolios are also the basis for a wide range of well-established patterns in the cross section of equity returns. Details on the construction of these portfolios can be found on Kenneth French's Web site (http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html), and their summary statistics are shown in Table 1.

Our benchmark analysis employs annual data, whereas the Supplementary Material contains the corresponding results using monthly and quarterly portfolio returns. Our main results for the SIZE/BM and LTR portfolios span the post–Great Depression period (1933–2012),[7] whereas profitability and investment portfolios are available since 1964.

For external validation, we additionally examine whether the GDA-I model can explain the level and cross-sectional variation of bond and equity index option premia. To this end, we use five corporate bond portfolios sorted on the basis of their credit ratings,[8] six Treasury bond portfolios sorted on maturity,[9] and six equity index option portfolios.[10] The Treasury bond portfolios are available since 1952, the corporate bond portfolios are available for the period 1976–2009, and the option portfolios are available for the period 1987–2011.

[6]In Section A8 of the Supplementary Material, we also consider 100 size/book-to-market portfolios and 10 earnings-to-price portfolios as alternative value-related cross sections.

[7]In Section A2 of the Supplementary Material, we alternatively examine the entire sample period for which aggregate consumption data are available from the Bureau of Economic Analysis (BEA) (i.e., 1930–2012), including Great Depression years.

[8]These corporate bond portfolios have been constructed by Nozawa (2012), and they are available on Michael Weber's Web site (http://faculty.chicagobooth.edu/michael.weber/).

[9]These are the Fama Maturity Portfolios provided by the Center for Research in Security Prices (CRSP).

[10]In particular, we use the 30-day put and call Standard & Poor's (S&P) 500 index option portfolios with moneyness levels of 90%, 100%, and 110%, respectively, of Constantinides, Jackwerth, and Savov (2013). These option portfolio returns are available on Alexi Savov's Web site (http://pages.stern.nyu.edu/~asavov/alexisavov/).

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TABLE 1
Summary Statistics

Table 1 shows the summary statistics of annual excess returns for a selection of the portfolios used in this study. We consider the following portfolio sorts: the 25 size/book-to-market (SIZE/BM) portfolios, the 25 size/operating profitability (SIZE/OP) portfolios, the 25 size/investment (SIZE/INV) portfolios, the 10 long-term reversal (LTR) portfolios, the 5 corporate bond portfolios of Nozawa (2012) constructed on the basis of bonds' credit ratings (CBOND), the 6 Fama Treasury bond portfolios sorted on maturity (TBOND), and the 6 equity index option portfolios of Constantinides et al. (2013) (OPTION). Panel A shows summary statistics for equal-weighted equity portfolio returns, Panel B shows statistics for value-weighted equity portfolio returns, and Panel C shows the corresponding statistics for nonequity asset classes. The sample period for the SIZE/BM and LTR portfolios is from 1933 to 2012, whereas the SIZE/OP and SIZE/INV portfolios are available since 1964. The sample period for the corporate bond portfolios is from 1976 to 2009, for the Treasury bond portfolios is from 1952 to 2012, and for the equity index option portfolios is from 1987 to 2011.

	25 SIZE/BM		25 SIZE/OP		25 SIZE/INV		10 LTR	
	SIZE1/BM5	SIZE5/BM1	SIZE1/OP1	SIZE5/OP5	SIZE1/INV1	SIZE5/INV5	LTR1	LTR10
Panel A. Equal-Weighted Equity Portfolios								
Average	27.29	7.99	14.87	7.67	21.53	4.15	27.13	9.02
Std. dev.	48.64	19.91	41.13	19.42	44.19	23.95	52.78	26.90
Panel B. Value-Weighted Equity Portfolios								
Average	24.17	11.98	13.13	12.15	18.85	10.47	19.68	12.48
Std. dev.	36.90	19.35	33.70	18.42	34.01	23.20	35.46	24.86

	5 CBOND		6 TBOND		6 OPTION	
	CBOND1	CBOND5	TBOND1	TBOND6	OTM_CALL	OTM_PUT
Panel C. Alternative Asset Classes						
Average	−0.08	5.79	0.62	2.04	1.94	25.49
Std. dev.	6.10	15.10	0.94	6.82	14.96	26.46

The GDA-I discount factor is a consumption-based model. To construct the per-capita aggregate consumption series, we use personal consumption expenditures (PCE) and PCE price index data from the BEA.[11] Aggregate consumption is defined as services plus nondurables. Each component of aggregate consumption is deflated by its corresponding PCE price index (base year is 2009). Population data are from the U.S. Census Bureau and recession dates are from the NBER. In matching consumption growth with asset returns, we follow the "beginning-of-period" convention, as in Campbell (2003), Yogo (2006), and Savov (2011), because beginning-of-period consumption growth is better aligned with asset returns.

B. Estimation Methodology

We test the competing asset pricing models in equations (6)–(10) using first-stage GMM (Hansen and Singleton (1982)) with an identity weighting matrix for the following system of Euler equations:[12]

$$\mathbb{E}[(R_{i,t} - R_{1Y,t})(1 - \mathbb{E}[M_t] + M_t)] \quad = \quad 0, \quad \text{for } i = 1,\ldots,n, \tag{11}$$

where $R_{1Y,t}$ is the 1-year interest rate, and n is the number of test assets. Consistent with Burnside's (2011) generic SDF representation for linear factor models, we

[11] Monthly consumption data used in Section A5 of the Supplementary Material are available since 1959. Quarterly consumption data used in Section A6 of the Supplementary Material are available since 1947.

[12] In Section A3 of the Supplementary Material, we alternatively estimate the various asset pricing models using a 2-stage GMM approach.

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augment the SDFs specified in equations (4) and (6)–(10) by $1-\mathbb{E}[M_t]$ to rule out a zero solution for risk prices because we are testing linear models on excess returns. Also, using the definition of covariance, we obtain the following equivalent expression for expected portfolio premia:

$$\mathbb{E}[R_{i,t} - R_{1\mathrm{Y},t}] \quad = \quad -\mathrm{cov}(R_{i,t} - R_{1\mathrm{Y},t}, M_t) \quad \text{for } i = 1,\ldots,n. \tag{12}$$

This GMM setup is equivalent to running a cross-sectional regression of portfolio premia on covariances imposing a 0 intercept. However, the critical advantage of the GMM specification is that it automatically corrects standard errors (SE) for the fact that covariances of excess returns with the SDF need also be estimated.

To test the performance of the GDA-I model in equation (4), we need to identify the set of disappointment consumption events, and hence we need to specify the values for the mean (μ_c), variance (σ_c^2), and autocorrelation (ϕ_c) of consumption growth. To this end, for the GDA-I model alone, we minimize the following augmented GMM system that fits the empirical consumption growth moments jointly with the unconditional Euler equations for excess portfolio returns:

$$\begin{bmatrix} \mathbb{E}[\Delta c_t] - \mu_c \\ \mathbb{E}[\Delta c_t^2] - \mu_c^2 - \sigma_c^2 \\ \mathbb{E}[\Delta c_t \Delta c_{t-1}] - \mu_c^2 - \phi_c \sigma_c^2 \\ \mathbb{E}\big[(R_{i,t} - R_{1\mathrm{Y},t})(1 - \mathbb{E}[M_t^{\text{GDA-I}}] + M_t^{\text{GDA-I}})\big] \text{ for } i = 1,\ldots,n \end{bmatrix} \quad = \quad \mathbf{0}. \tag{13}$$

In minimizing this GMM system, we use a diagonal weighting matrix in which the first three diagonal elements are very large numbers (10^8), and the remaining diagonal elements are 1. There are two reasons for overweighting the moment conditions for consumption growth. First, these moment conditions have a different scale from the ones for portfolio returns. For example, the annual return for the SIZE1/BM5 portfolio (=27%) is much larger than the sample autocovariance of the annual consumption growth (=0.005%). Therefore, these consumption growth moments need to be weighted accordingly. Second, by overweighting the consumption growth moments, we are not allowing the estimation procedure to fit portfolio premia at the expense of errors in the consumption growth process (e.g., inflating the variability or the persistence of consumption growth).

We assess the overall model fit using the χ^2 test (Hansen (1982)), the cross-sectional R^2, and the cross-sectional root-mean-square error (RMSE). The magnitude of the RMSE for each cross section should be compared with the corresponding average portfolio premia reported in Table 1. Finally, it should be noted that the critique of Lewellen, Nagel, and Shanken (2010) regarding the mechanical fit of asset pricing models due to the factor structure of certain cross sections of portfolios is less relevant for the GDA-I SDF because it is a single-factor model.

IV. Results

In our benchmark analysis, we estimate the various asset pricing models using annual returns for the 25 SIZE/BM, 25 SIZE/OP, 25 SIZE/INV, and 10 LTR portfolios, in turn. The results for equal-weighted portfolios are reported in Table 2, whereas the corresponding results for value-weighted portfolios are reported in Table 3.

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A. Model Fit

The results in Table 2 for equal-weighted portfolios show that the GDA-I model achieves a very good fit in absolute terms, with R^2 being as high as 90% in the cross section of the 25 SIZE/BM portfolios. The goodness of fit is also high across the 25 SIZE/OP, 25 SIZE/INV, and 10 LTR portfolios (R^2s = 88%, 74%, and 87%, respectively). Additionally, the GDA-I model yields low RMSEs. This success is particularly notable for the 25 SIZE/BM and the 25 SIZE/OP portfolios, given the substantial cross-sectional dispersion in the sample premia of

TABLE 2

GMM Results for Annual Portfolio Returns: Equal-Weighted Portfolios

Table 2 shows the generalized method of moments (GMM) results for different equal-weighted portfolio sorts and asset pricing models at the annual frequency. For this set of tests, we estimate consumption growth moments, the disappointment-aversion (DA) coefficient $\tilde{\theta}$, and the disappointment threshold d_2 for the generalized disappointment-aversion indicator (GDA-I) model from equation (4) using the augmented GMM system from equation (13). We consider four equal-weighted portfolio sorts as test assets: the 25 size/book-to-market (SIZE/BM) portfolios (Panel A), the 25 size/operating profitability (SIZE/OP) portfolios (Panel B), the 25 size/investment (SIZE/INV) portfolios (Panel C), and the 10 long-term reversal (LTR) portfolios (Panel D). GDA-I is the disappointment-aversion discount factor from equation (4), and CCAPM is the consumption-based discount factor from equation (6). CAPM is the market model from equation (7). FF3 and FF5 are the Fama–French (1993), (2015) 3- and 5-factor models in equations (8) and (9), respectively. NBER SDF is the recession-based stochastic discount factor from equation (10). GDA_IND is the disappointment indicator. μ_c, σ_c^2, and $\phi_c\sigma_c^2$ denote the consumption growth mean, variance, and autocovariance (×100), respectively. CONS is the aggregate consumption growth, MKT is the excess market return, SMB is the size factor, HML is the value factor, RMW is the profitability factor, CMA is the investment factor, and NBER_IND is the National Bureau of Economic Research (NBER) recession indicator. $\tilde{\alpha}$ is the risk-aversion parameter in the consumption-based capital asset pricing model (CCAPM). t-statistics are shown in parentheses. χ^2, DOF, and p are the first-stage χ^2 test (Hansen (1982)), degrees of freedom, and p-value that all moment conditions are jointly 0. RMSE and R^2 are the root-mean-square error (×100) and cross-sectional R-square, respectively. The sample period is 1933–2012. The sample for the FF5, the operating profitability, and investment portfolios begins in 1964.

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel A. 25 SIZE/BM Portfolios						
GDA_IND ($\tilde{\theta}$)	4.126					
	(4.207)					
d_2	−0.770					
	(−3.006)					
μ_c	2.168					
	(12.542)					
σ_c^2	0.023					
	(5.128)					
$\phi_c\sigma_c^2$	0.005					
	(0.868)					
CONS ($\tilde{\alpha}$)		57.331				
		(3.499)				
MKT			2.935	2.043	3.191	
			(4.781)	(2.375)	(2.218)	
SMB				0.335	1.912	
				(0.309)	(1.242)	
HML				3.026	−0.473	
				(3.079)	(−0.153)	
RMW					0.967	
					(0.308)	
CMA					9.321	
					(2.080)	
NBER_IND						9.157
						(1.053)
χ^2	28.795	87.580	97.940	62.460	42.570	5.543
DOF	23	24	24	22	20	24
p	0.187	0	0	0	0.002	1.000
RMSE	1.345	2.107	2.973	1.648	1.563	2.358
R^2	0.901	0.758	0.519	0.852	0.821	0.697

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TABLE 2 (continued)
GMM Results for Annual Portfolio Returns: Equal-Weighted Portfolios

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel B. 25 SIZE/OP Portfolios						
GDA_IND ($\tilde{\theta}$)	3.809					
	(1.710)					
d_2	−0.480					
	(−2.004)					
μ_c	1.945					
	(10.494)					
σ_c^2	0.016					
	(4.717)					
$\phi_c\sigma_c^2$	0.010					
	(2.309)					
CONS ($\tilde{\alpha}$)		88.188				
		(1.672)				
MKT			2.793	2.430	2.546	
			(2.698)	(1.797)	(2.051)	
SMB				1.452	2.037	
				(1.258)	(2.024)	
HML				5.156	0.805	
				(1.839)	(0.267)	
RMW					3.523	
					(1.824)	
CMA					2.142	
					(0.596)	
NBER_IND						4.527
						(1.183)
χ^2	15.506	30.481	69.366	51.500	36.759	15.714
DOF	23	24	24	22	20	24
p	0.875	0.169	0	0	0.012	0.898
RMSE	1.031	1.728	2.674	1.338	1.016	3.247
R^2	0.879	0.660	0.187	0.796	0.882	−0.197
Panel C. 25 SIZE/INV Portfolios						
GDA_IND ($\tilde{\theta}$)	3.448					
	(2.477)					
d_2	−0.584					
	(−0.031)					
μ_c	1.945					
	(10.496)					
σ_c^2	0.016					
	(56.847)					
$\phi_c\sigma_c^2$	0.010					
	(2.316)					
CONS ($\tilde{\alpha}$)		91.068				
		(1.701)				
MKT			2.889	2.841	3.713	
			(2.785)	(2.010)	(2.512)	
SMB				1.242	2.992	
				(0.921)	(1.707)	
HML				6.946	−4.664	
				(4.281)	(−0.760)	
RMW					4.686	
					(1.138)	
CMA					14.868	
					(1.811)	
NBER_IND						4.642
						(1.188)
χ^2	31.137	49.241	131.615	64.842	57.750	15.146
DOF	23	24	24	22	20	24
p	0.119	0.001	0	0	0	0.916
RMSE	1.952	3.028	3.967	2.020	1.696	4.770
R^2	0.743	0.383	−0.057	0.725	0.806	−0.529

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TABLE 2 (continued)
GMM Results for Annual Portfolio Returns: Equal-Weighted Portfolios

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel D. 10 LTR Portfolios						
GDA_IND ($\tilde{\theta}$)	4.327					
	(2.912)					
α_2	−0.752					
	(−4.682)					
μ_c	2.168					
	(12.542)					
σ_c^2	0.023					
	(5.128)					
$\phi_c\sigma_c^2$	0.005					
	(0.868)					
CONS ($\tilde{\alpha}$)		62.493				
		(3.352)				
MKT			3.189	1.042	1.558	
			(4.875)	(0.695)	(0.418)	
SMB				2.205	10.133	
				(0.830)	(1.844)	
HML				3.067	−14.952	
				(1.330)	(−1.052)	
RMW					12.369	
					(1.121)	
CMA					31.899	
					(1.248)	
NBER_IND						9.797
						(1.054)
χ^2	13.411	24.233	37.292	25.435	2.763	2.943
DOF	8	9	9	7	5	9
p	0.098	0.003	0	0	0.736	0.966
RMSE	1.640	1.658	2.213	0.954	0.708	2.450
R^2	0.868	0.865	0.759	0.960	0.967	0.705

these portfolios. Finally, on the basis of the χ^2 test of overidentifying restrictions, the GDA-I model cannot be formally rejected in the examined sets of portfolios.

The empirical success of the GDA-I model is also illustrated in Figure 1, which plots sample portfolio premia versus fitted expected excess returns for all models. These scatterplots show that the GDA-I model can successfully align when fitted with sample premia across all portfolio sorts.

The goodness of fit of the GDA-I model is even more striking when assessed in relation to the alternative, commonly used asset pricing models. The GDA-I model outperforms the CCAPM in terms of goodness of fit. This outperformance is more pronounced among the 25 SIZE/INV portfolios, where the CCAPM performs very poorly. This is also evident from Panel C of Figure 1. It should also be mentioned that the CCAPM is formally rejected by the χ^2 test in all but one of the portfolio sets considered. Equally importantly, the GDA-I model outperforms the CAPM, which yields much lower R^2s and much higher RMSEs across the board. Interestingly, the CAPM not only badly fails to price the 25 SIZE/INV and the 25 SIZE/OP portfolios, but it also yields a twice-as-high RMSE as the GDA-I model for the 25 SIZE/BM portfolios. Moreover, the CAPM is rejected by the χ^2 test for all sets of portfolios considered.

A challenging benchmark for the GDA-I model is the FF3. Despite the fact that our model utilizes a single macroeconomic indicator variable, its goodness of fit is very similar to that of the Fama–French model. In fact, in the cross section

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of the 25 SIZE/BM portfolios, from which the SMB and HML factors are constructed, the GDA-I model yields a higher R^2 and a lower RMSE relative to this model. The same is true for the 25 SIZE/INV and the 25 SIZE/OP portfolios. In contrast, the 3-factor model outperforms the GDA-I model in the cross section of the 10 LTR portfolios, where it achieves an almost perfect fit. The overall relative parity in terms of goodness of fit among these two models is illustrated in Figure 1. Finally, it should be mentioned that the Fama–French model is rejected by the χ^2 test for all sets of portfolios considered.

The GDA-I model also compares very well with respect to the FF5. The corresponding estimation results in Table 2 show that the 5-factor model outperforms the GDA-I model across the 25 SIZE/INV portfolios, but it underperforms our model across the 25 SIZE/BM portfolios. Both models achieve an equally good fit across the 25 SIZE/OP portfolios.

TABLE 3

GMM Results for Annual Portfolio Returns: Value-Weighted Portfolios

Table 3 shows the generalized method of moments (GMM) results for different value-weighted portfolio sorts and asset pricing models at the annual frequency. For this set of tests, we estimate consumption growth moments, the disappointment-aversion (DA) coefficient $\tilde{\theta}$, and the disappointment threshold d_2 for the generalized disappointment-aversion indicator (GDA-I) model from equation (4) using the augmented GMM system from equation (13). Table 3 does not report estimation results for the consumption growth moments. We consider four value-weighted portfolio sorts as test assets: the 25 size/book-to-market (SIZE/BM) portfolios (Panel A), the 25 size/operating profitability (SIZE/OP) portfolios (Panel B), the 25 size/investment (SIZE/INV) portfolios (Panel C), and the 10 long-term reversal (LTR) portfolios (Panel D). GDA-I is the disappointment aversion discount factor, and CCAPM is the consumption-based discount factor. CAPM is the market model. FF3 and FF5 are the Fama–French (1993), (2015) 3- and 5-factor models. NBER SDF is the recession-based discount factor. GDA_IND is the disappointment indicator, CONS is the aggregate consumption growth, MKT is the excess market return, SMB is the size factor, HML is the value factor, RMW is the profitability factor, CMA is the investment factor, and NBER_IND is the National Bureau of Economic Research (NBER) recession indicator. $\tilde{\alpha}$ is the risk-aversion parameter in the consumption-based capital asset pricing model (CCAPM). t-statistics are shown in parentheses. χ^2, DOF, and p are the first-stage χ^2 test, degrees of freedom, and p-value that all moment conditions are jointly 0. RMSE and R^2 are the root-mean-square error (×100) and cross-sectional R-square, respectively. The sample period is 1933–2012. The sample for the FF5, the operating profitability, and the investment portfolios begins in 1964.

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel A. 25 SIZE/BM Portfolios						
GDA_IND ($\tilde{\theta}$)	4.447					
	(3.593)					
d_2	−0.872					
	(−4.681)					
CONS ($\tilde{\alpha}$)		56.398				
		(3.532)				
MKT			2.861	2.221	3.384	
			(4.599)	(2.776)	(2.734)	
SMB				−0.008	1.470	
				(−0.007)	(1.042)	
HML				2.856	0.750	
				(2.891)	(0.252)	
RMW					4.113	
					(1.260)	
CMA					5.478	
					(1.470)	
NBER_IND						7.659
						(1.192)
χ^2	34.238	104.775	120.754	100.862	71.257	7.978
DOF	23	24	24	22	20	24
p	0.061	0	0	0	0	0.999
RMSE	1.358	2.165	2.819	1.591	1.497	2.867
R^2	0.827	0.561	0.256	0.763	0.754	0.231

(continued on next page)

TABLE 3 (continued)
GMM Results for Annual Portfolio Returns: Value-Weighted Portfolios

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel B. 25 SIZE/OP Portfolios						
GDA_IND ($\tilde{\theta}$)	3.350					
	(1.987)					
d_2	−0.393					
	(−2.268)					
CONS ($\tilde{\alpha}$)		81.574				
		(1.762)				
MKT			2.623	2.703	2.566	
			(2.594)	(2.073)	(1.831)	
SMB				0.862	1.047	
				(0.703)	(0.844)	
HML				4.555	2.348	
				(1.720)	(0.548)	
RMW					3.642	
					(1.734)	
CMA					−2.124	
					(−0.372)	
NBER_IND						3.550
						(1.467)
χ^2	19.718	52.336	77.643	60.959	48.279	24.582
DOF	23	24	24	22	20	24
p	0.658	0	0	0	0	0.431
RMSE	1.151	1.939	2.641	1.651	0.942	2.965
R^2	0.791	0.409	−0.095	0.571	0.860	−0.380
Panel C. 25 SIZE/INV Portfolios						
GDA_IND ($\tilde{\theta}$)	3.622					
	(1.466)					
d_2	−0.609					
	(−5.285)					
CONS ($\tilde{\alpha}$)		84.516				
		(1.803)				
MKT			2.726	3.335	3.606	
			(2.676)	(2.555)	(2.782)	
SMB				0.169	0.820	
				(0.124)	(0.505)	
HML				5.785	2.246	
				(3.765)	(0.450)	
RMW					2.005	
					(0.677)	
CMA					4.553	
					(0.809)	
NBER_IND						3.669
						(1.489)
χ^2	21.249	42.580	122.485	101.728	73.478	23.746
DOF	23	24	24	22	20	24
p	0.565	0.011	0	0	0	0.476
RMSE	1.460	2.283	3.277	1.360	1.330	3.605
R^2	0.707	0.286	−0.471	0.746	0.757	−0.780

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TABLE 3 (continued)
GMM Results for Annual Portfolio Returns: Value-Weighted Portfolios

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel D. 10 LTR Portfolios						
GDA_IND ($\tilde{\theta}$)	4.829					
	(1.791)					
d_2	−0.780					
	(−6.470)					
CONS ($\tilde{\alpha}$)		57.532				
		(3.221)				
MKT			2.734	2.565	3.434	
			(4.640)	(2.729)	(2.301)	
SMB				−1.077	0.513	
				(−0.539)	(0.180)	
HML				3.246	0.094	
				(1.879)	(0.021)	
RMW					0.644	
					(0.144)	
CMA					6.057	
					(0.929)	
NBER_IND						6.890
						(1.317)
χ^2	4.121	8.592	10.816	6.391	6.018	3.059
DOF	8	9	9	7	5	9
p	0.845	0.475	0.288	0.494	0.304	0.961
RMSE	0.626	0.742	1.356	0.513	0.717	1.374
R^2	0.910	0.874	0.581	0.939	0.844	0.570

To show that the empirical success of the GDA-I model does not derive from naively mimicking recession periods, we also compare it with the NBER model. In fact, we find that a simple NBER recession indicator cannot price at all the SIZE/INV and SIZE/OP portfolios because it yields very high RMSEs and very low R^2s. The very poor performance of the NBER model can be visualized by the scatterplots in the fourth column of Figure 1. These results imply that NBER recessions are not particularly important for risk premia.

Overall, the performance of the GDA-I model indicates that investors are sensitive to disappointment consumption years and require a higher premium for holding assets that perform badly during these years. In fact, the results in Table 2 suggest that downside consumption risk, as measured by disappointment events in consumption growth, can explain both the level and the cross-sectional variation in portfolio premia. Hence, we argue that to a large extent, SIZE, BM, OP, INV, and LTR premia are compensation for exposure to downside consumption risk.

Contrary to a number of previous studies that propose mispricing as an explanation for equity premia, our results support a risk-based rationalization for the cross section of stock returns. Specifically, the successful performance of the GDA-I model in the SIZE/BM cross section challenges the conclusions of Lakonishok, Shleifer, and Vishny (1994), who suggest that book-to-market effects are a result of investor overreaction to past firm performance. Lakonishok et al. (1994) and De Bondt and Thaler (1985) propose an overreaction explanation for the long-term reversal premium as well. Yet, the reported results for the LTR portfolios in Table 2 indicate that disappointment risk can explain the long-term reversal puzzle too.

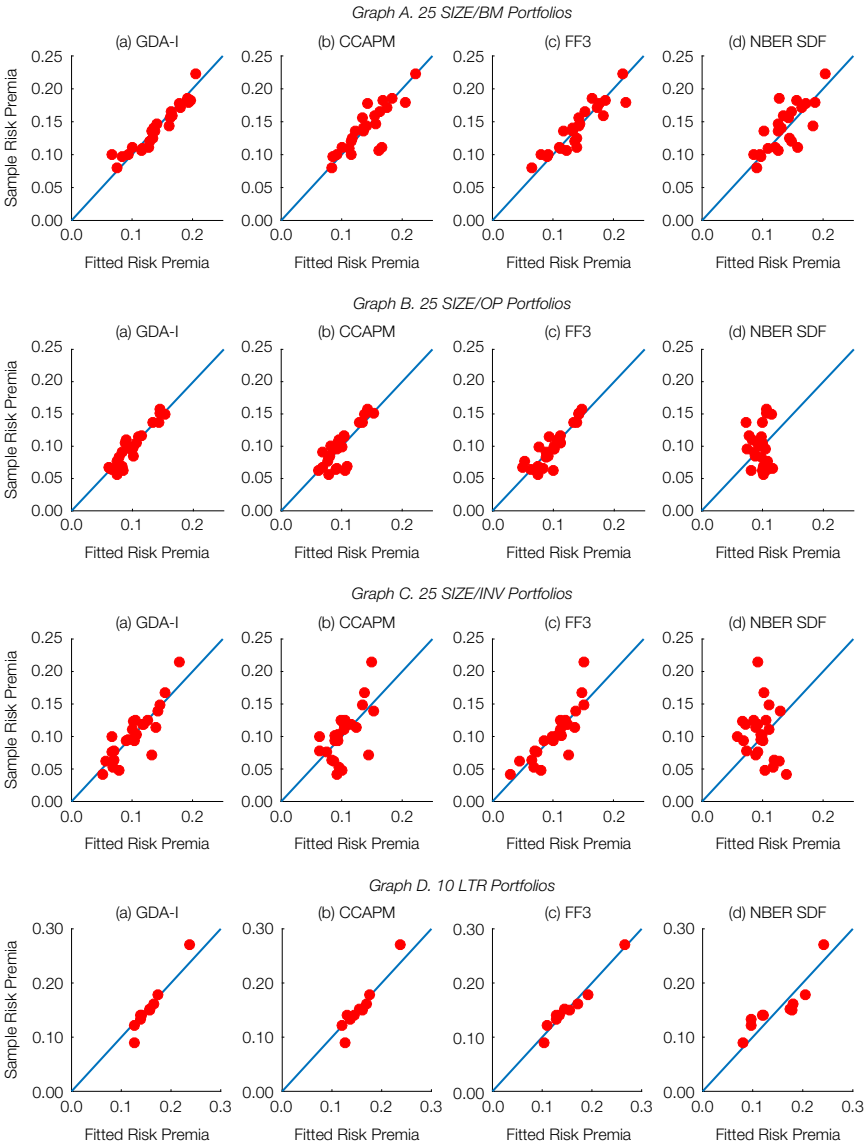
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FIGURE 1

Sample and Fitted Risk Premia for Equity Portfolios: Equal-Weighted Returns

Figure 1 shows sample and fitted annual risk premia for the 25 size/book-to-market (SIZE/BM) portfolios (Graph A), the 25 size/operating profitability (SIZE/OP) portfolios (Graph B), the 25 size/investment (SIZE/INV) portfolios (Graph C), and the 10 long-term reversal (LTR) portfolios (Graph D). All portfolios are equal weighted. Fitted risk premia are estimated according to the expression in equation (12) for the generalized disappointment-aversion indicator (GDA-I), consumption-based capital asset pricing model (CCAPM), Fama–French (1993) 3-factor model (FF3), and National Bureau of Economic Research (NBER) discount factors. Estimation results are shown in Table 2. The sample period is from 1933 to 2012, with the exception of the operating profitability and investment portfolios that start in 1964.



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B. Prices of Risk and Disappointment Threshold Coefficient

In addition to model fit, Table 2 reports the estimated prices of risk across the various models and portfolio sorts considered. For the GDA-I and CCAPM specifications, the prices of risk have a structural interpretation and can be directly mapped into preference parameters. Specifically, the price of risk for the GDA-I model is the DA coefficient $\tilde{\theta}$ from equation (4), whereas the price of risk for the CCAPM is the second-order risk-aversion parameter $\tilde{\alpha}$ from equation (6). Hence, we can also assess the plausibility of these two models by examining their implied preference parameters.

Regarding the CCAPM, the estimated risk-aversion coefficients reported in Table 2 range from 57 for the SIZE/BM portfolios to 91 for the SIZE/INV portfolios. These magnitudes are consistent with the ones reported by Mehra and Prescott (1985), Cochrane (2001), and Savov (2011) and reflect the equity premium puzzle. In other words, the CCAPM requires extremely large risk-aversion parameters to match equity premia. Echoing Rabin (2000), accepting such a high degree of concavity in the representative investor's utility function is equivalent to accepting that this investor would paradoxically reject even extremely favorable larger-stake gambles, rendering the investor implausibly risk averse.

In contrast, the estimated DA coefficients reported in Table 2 range from 3.4 for the SIZE/INV portfolios to 4.3 for the LTR portfolios, and they remain fairly stable across the examined cross sections. This range of values for the DA coefficient is very close to the value required by Ang et al. (2005) to explain the historical equity premium and implies a very reasonable risk-taking behavior in the Rabin (2000) games (see Ang et al. (2005)). It should also be noted that the estimated factor coefficients reported in Table 2 reveal an important limitation of the Fama–French (1993), (2015) models. These multifactor models achieve relatively low RMSEs in each set of test assets at the expense of strikingly different factor coefficient estimates across these cross sections, hindering their interpretation in a theoretically consistent fashion.

Finally, Table 2 also reports estimates for the coefficient d_2 in each cross section. According to equation (5), the parameter d_2 characterizes the threshold for disappointment in terms of standard deviations away from the conditional mean of consumption growth. For example, we find that for the SIZE/BM (LTR) portfolios in the full sample period, disappointment consumption events occur when realized consumption growth is 0.77 (0.75) standard deviations below its conditional mean. The threshold is somewhat higher for the SIZE/INV and SIZE/OP portfolios in the post-1964 period, but the corresponding estimates for d_2 are not directly comparable because the consumption growth moments are also different in this subperiod. In particular, the estimates of the annual consumption growth moments for the period 1933–2012 are 2.17% for average consumption growth, 0.024% for consumption growth variance, and 0.005% for consumption growth autocovariance. For the period 1964–2012, the corresponding estimates are 1.94%, 0.016%, and 0.010%, respectively.

C. Value-Weighted Portfolios

Following Liu et al. (2009), the previous empirical tests were based on equal-weighted portfolio returns because they exhibit more pronounced cross-sectional

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variation. For robustness, in this section, we alternatively use value-weighted portfolio returns. Table 3 and Figure 2 report the corresponding results for the various asset pricing models at the annual frequency.

Comparing the results reported in Table 3 and Figure 2 for value-weighted returns with the ones reported in Table 2 and Figure 1 for equal-weighted returns, we can make the following observations. First, the GDA-I model maintains its very good fit both in absolute and in relative terms, yielding low pricing errors across the four sets of portfolios used in our benchmark analysis. The GDA-I model outperforms the competing single-factor models, and its goodness of fit is comparable to that of the multifactor Fama–French (1993), (2015) models.

Second, the DA coefficients estimated from the GDA-I model are very close to the ones derived using equal-weighted portfolios, implying again a very reasonable risk-taking behavior for the representative investor. The same is true for the corresponding estimates of the disappointment threshold coefficient. Moreover, the second-order risk-aversion coefficients estimated from the CCAPM are again implausibly high.

Third, the goodness of fit that the multifactor Fama–French (1993), (2015) models achieve comes at the expense of strikingly different factor coefficient estimates, both across portfolio cross sections and across weighting schemes. In sum, the results presented in Table 3 and Figure 2 confirm that the pricing ability of the proposed single-factor GDA-I model is not sensitive to the choice of the portfolio weighting scheme.

D. Additional Results

In the Supplementary Material, we repeat our benchmark analysis following a series of alternative empirical approaches using i) the full sample period for which aggregate consumption data are available from the BEA (i.e., 1930–2012) (Section A2), ii) a 2-stage GMM estimation procedure (Section A3), iii) a recursive GMM estimation approach with expanding time windows (Section A4), iv) monthly portfolio returns (Section A5), v) quarterly portfolio returns (Section A6), vi) a joint cross section combining all equity portfolios (Section A7), and vii) alternative value-related cross sections (100 size/book-to-market and 10 earnings-to-price portfolios in Section A8). The results from these alternative empirical approaches are similar to the ones reported in Tables 2 and 3. In particular, the GDA-I model yields a very good fit across portfolio sorts, both in absolute and in relative terms, and the disappointment-aversion coefficient estimates are rather stable and imply reasonable risk-taking behavior.[13]

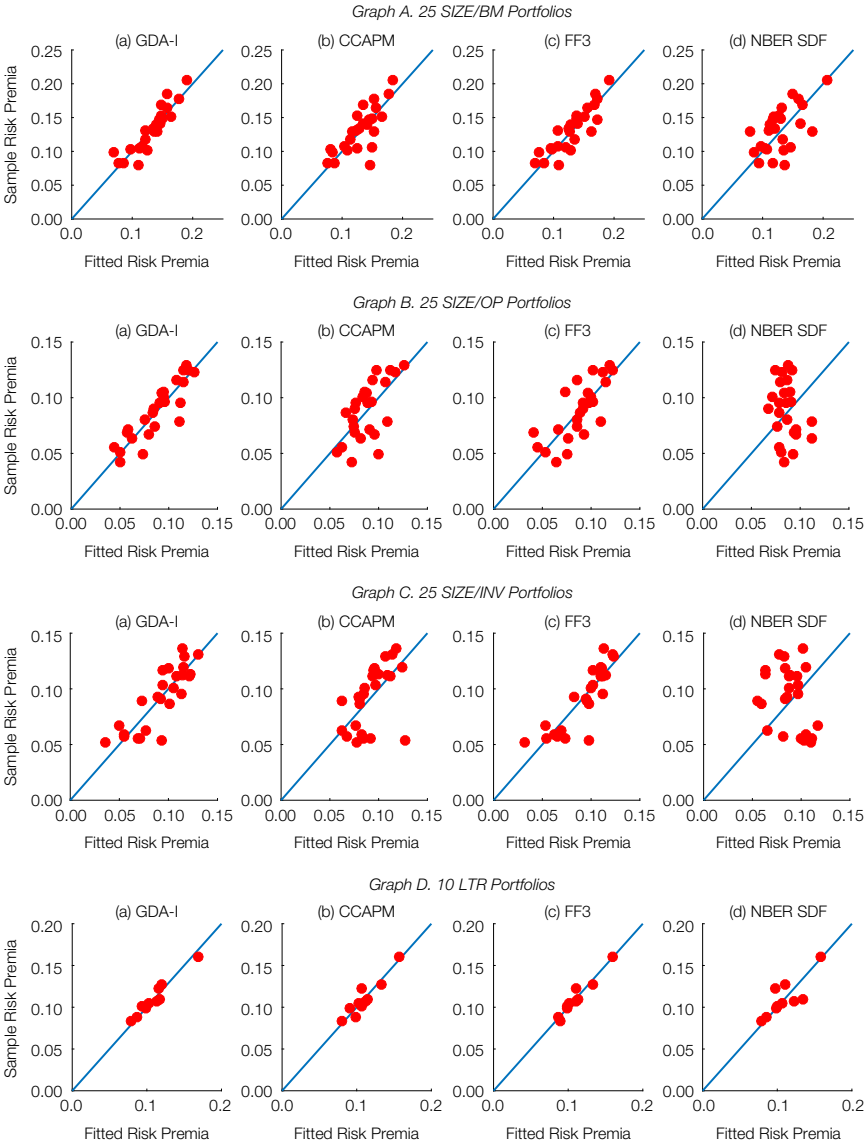
Another important point regarding the empirical tests in Tables 2 and 3 is that in fitting the three return-based SDFs (CAPM, FF3, FF4), the prices of risk are free parameters estimated by the GMM. This approach guarantees maximum flexibility for these models, whose prices of risk vary substantially across the different test assets. However, the factors in these SDFs are themselves excess returns of

[13] We have also repeated our benchmark analysis for the post-1945 period. The explanatory power of the GDA-I model remains strong during this sample period. In particular, the R^2s for the GDA-I model with respect to the 25 SIZE/BM and 10 LTR portfolios are 86% and 87%, respectively. The corresponding R^2s for the CCAPM are 63% and 46%, and for the FF3 specification, they are 75% and 94%.

FIGURE 2

Sample and Fitted Risk Premia for Equity Portfolios: Value-Weighted Returns

Figure 2 shows sample and fitted annual risk premia for the 25 size/book-to-market (SIZE/BM) portfolios (Graph A), the 25 size/operating profitability (SIZE/OP) portfolios (Graph B), the 25 size/investment (SIZE/INV) portfolios (Graph C), and the 10 long-term reversal (LTR) portfolios (Graph D). All portfolios are value weighted. Fitted risk premia are estimated according to the expression in equation (12) for the generalized disappointment-aversion indicator (GDA-I), consumption-based capital asset pricing model (CCAPM), Fama–French (1993) 3-factor model (FF3), and National Bureau of Economic Research (NBER) discount factors. Estimation results are shown in Table 3. The sample period is from 1933 to 2012, with the exception of the operating profitability and investment portfolios that start in 1964.



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traded assets. Thus, in principle, the prices of risk in these models should be constrained by the expected returns of the factors, as suggested by Cochrane ((2001), p. 107). Therefore, in Section A9 of the Supplementary Material, we repeat our benchmark analysis imposing price-of-risk restrictions on the return-based SDFs. These restrictions are derived from the additional condition that the return-based models should be able to perfectly fit the corresponding return-based factors.

According to the results in Section A9 of the Supplementary Material, when we require the return-based SDFs to perfectly price their factors, their goodness of fit considerably deteriorates. For instance, the R^2s of the restricted FF5 model in the four portfolio sorts of our benchmark analysis become 72%, 62%, 66%, and 40%, respectively. These R^2s are substantially worse than those of the single-factor GDA-I model reported in Table 2. This is an alternative way of interpreting the instability of the prices of risk that the return-based SDFs exhibit across the various test assets. In contrast, the single-factor GDA-I model can explain expected returns with a price of disappointment risk that remains stable across the alternative cross sections.

V. Alternative Asset Classes

Up to this point, we have examined an extensive set of equity portfolio sorts. However, the proposed consumption-based GDA-I model should be able to price all assets in the economy. To this end, this section examines the pricing ability of the GDA-I model with respect to other asset classes. In particular, we consider the five corporate bond portfolios of Nozawa (2012) that have been constructed on the basis of bonds' credit ratings; the six Fama maturity-sorted Treasury bond portfolios; and the six 30-day put and call S&P 500 index option portfolios, with moneyness levels of 90%, 100%, and 110%, respectively, of Constantinides et al. (2013).

First, we estimate the GDA-I and the competing pricing models for each of these three sets of portfolios. To ensure that the estimated model coefficients simultaneously explain the equity premium apart from fitting the corresponding cross section, we also add the equity market portfolio to each of these three sets of portfolios. Second, along the lines of Lettau, Maggiori, and Weber (2014), we test the pricing ability of the various models using a joint cross section of equity, corporate bond, Treasury bond, and equity index option portfolios. In this way, we examine whether the same risk price(s) can simultaneously explain the cross-sectional premia that different asset classes yield.

Corporate bond, Treasury bond, and equity index option portfolio returns are jointly available only for the period 1987–2009. Due to this short sample period, and to ensure that each of the asset classes is evenly represented in the joint cross section, the equity portfolios we use are the six Fama–French SIZE/BM portfolios. We present results for the annual portfolio returns of the alternative test assets in Table 4, whereas Section A10 in the Supplementary Material contains the corresponding results using monthly portfolio returns.

Panel A of Table 4 presents the results for the corporate bond portfolios. We find that the GDA-I model can almost perfectly explain the cross section of their premia. This goodness of fit can be visualized in Panel A of Figure 3. In fact,

the corporate bonds with the highest credit spreads are the most exposed ones to disappointment consumption risk. Hence, the high premia they yield can be regarded as compensation for this aspect of risk. Interestingly, the estimated price of disappointment risk (estimate = 3.7) is very similar in magnitude to the ones derived from equity portfolio sorts. In relative terms, the GDA-I model outperforms the competing single-factor models, and the CCAPM still implies an implausibly high risk-aversion coefficient. Equally importantly, the multifactor Fama–French (1993), (2015) models can only fit this cross section through extremely large or even negative factor coefficient estimates, which are at odds with the corresponding estimates derived from equity portfolio sorts.

TABLE 4
GMM Results for Alternative Asset Classes

Table 4 shows the generalized method of moments (GMM) results for different asset classes and asset pricing models at the annual frequency. For this set of tests, we estimate consumption growth moments, the disappointment-aversion (DA) coefficient $\tilde{\theta}$, and the disappointment threshold d_2 for the generalized disappointment-aversion indicator (GDA-I) model from equation (4) using the augmented GMM system from equation (13). Table 4 does not report estimation results for the consumption growth moments. We consider four sets of portfolios as test assets: the five corporate bond portfolios of Nozawa (2012) constructed on the basis of bonds' credit ratings (Panel A), the six Fama Treasury bond portfolios sorted on maturity (Panel B), the six equity index option portfolios of Constantinides et al. (2013) (Panel C), and a joint cross section of the aforementioned portfolios together with the six Fama–French equal-weighted size/book-to-market portfolios (Panel D). In each of these cross sections, we also include the equity market portfolio as a test asset. GDA-I is the disappointment-aversion discount factor, and CCAPM is the consumption-based discount factor. CAPM is the market model. FF3 and FF5 are the Fama–French (1993), (2015) 3- and 5-factor models. NBER SDF is the recession-based discount factor. GDA_IND is the disappointment indicator, CONS is the aggregate consumption growth, MKT is the excess market return, SMB is the size factor, HML is the value factor, RMW is the profitability factor, CMA is the investment factor, and NBER_IND is the National Bureau of Economic Research (NBER) recession indicator. $\tilde{\alpha}$ is the risk-aversion parameter in the consumption-based capital asset pricing model (CCAPM). t-statistics are shown in parentheses. χ^2, DOF, and p are the first-stage χ^2 test, degrees of freedom, and p-value that all moment conditions are jointly 0. Panel D does not show t-statistics or p-values for the χ^2 test due to the limited time-series observations relative to the number of portfolios. RMSE and R^2 are the root-mean-square error (×100) and cross-sectional R-square, respectively. The sample period is 1976–2009 for the corporate bond portfolios, 1952–2012 for the Treasury bond portfolios (1964–2012 in the FF5 model), 1987–2011 for the equity index option portfolios, and 1987–2009 for the joint cross section.

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel A. 5 Corporate Bond Portfolios						
GDA_IND ($\tilde{\theta}$)	3.721					
	(1.442)					
d_2	−0.784					
	(−4.700)					
CONS ($\tilde{\alpha}$)		100.531				
		(1.126)				
MKT			2.763	5.508	18.134	
			(2.157)	(2.076)	(0.698)	
SMB				1.677	9.090	
				(0.380)	(0.498)	
HML				11.052	−11.009	
				(1.477)	(−0.373)	
RMW					37.755	
					(0.540)	
CMA					48.154	
					(0.632)	
NBER_IND						2.298
						(1.501)
χ^2	2.755	2.415	15.090	3.117	0.026	4.462
DOF	4	5	5	3	1	5
p	0.599	0.789	0.009	0.373	0.871	0.484
RMSE	0.689	1.096	1.486	0.729	0.124	3.150
R^2	0.939	0.847	0.719	0.932	0.998	−0.259

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TABLE 4 (continued)
GMM Results for Alternative Asset Classes

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel B. 6 Treasury Bond Portfolios						
GDA_IND ($\tilde{\theta}$)	4.776					
	(2.785)					
d_2	−0.626					
	(−1.619)					
CONS ($\tilde{\alpha}$)		74.295				
		(2.016)				
MKT			2.243	3.040	4.168	
			(2.851)	(0.964)	(0.922)	
SMB				5.546	3.579	
				(0.425)	(0.250)	
HML				22.027	27.266	
				(1.099)	(1.217)	
RMW					−7.379	
					(−0.561)	
CMA					−9.685	
					(−0.398)	
NBER_IND						2.916
						(1.180)
χ^2	8.058	25.917	36.151	2.360	0.834	15.032
DOF	5	6	6	4	2	6
p	0.153	0	0	0.669	0.657	0.020
RMSE	0.242	1.370	1.375	0.154	0.153	2.290
R^2	0.986	0.581	0.578	0.994	0.990	−0.169
Panel C. 6 Equity Index Option Portfolios						
GDA_IND ($\tilde{\theta}$)	4.433					
	(1.421)					
d_2	−2.054					
	(−4.271)					
CONS ($\tilde{\alpha}$)		77.859				
		(1.369)				
MKT			3.115	1.757	9.963	
			(1.965)	(0.429)	(0.611)	
SMB				19.776	3.198	
				(1.411)	(0.104)	
HML				−0.505	4.191	
				(−0.049)	(0.130)	
RMW					−7.107	
					(−0.167)	
CMA					44.848	
					(0.702)	
NBER_IND						2.810
						(1.064)
χ^2	12.561	44.925	28.714	3.209	0.853	17.426
DOF	5	6	6	4	2	6
p	0.027	0	0	0.523	0.652	0.007
RMSE	4.500	6.919	6.964	2.107	1.762	8.511
R^2	0.751	0.411	0.404	0.945	0.961	0.109

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TABLE 4 (continued)
GMM Results for Alternative Asset Classes

	GDA-I	CCAPM	CAPM	FF3	FF5	NBER SDF
Panel D. Joint Cross Section						
GDA_IND ($\tilde{\theta}$)	4.120					
d_2	−1.634					
CONS ($\tilde{\alpha}$)		80.403				
MKT			2.826	3.205	−0.571	
SMB				0.819	−2.269	
HML				4.342	13.336	
RMW					−14.478	
CMA					−6.595	
NBER_IND						3.192
χ^2	11.945	584.405	113.709	104.126	32.693	15.708
DOF	22	23	23	21	19	23
p						
RMSE	3.402	4.447	4.749	4.137	3.717	6.358
R^2	0.719	0.523	0.454	0.585	0.665	0.021

The results for the Treasury bond portfolios (Panel B in Table 4 and in Figure 3) are very similar to the ones for the corporate bond portfolios, even though the cross-sectional dispersion of Treasury bond premia is much less pronounced than the one of the corporate bond premia.

Panel C of Table 4 presents the corresponding results for the equity index option portfolios. The GDA-I model again achieves a very good fit, both in absolute and in relative terms. The DA coefficient estimate is also very similar in magnitude to the ones derived from equity portfolios as well as from corporate and Treasury bond portfolios (estimate = 4.4). Panel C of Figure 3 shows that the pricing errors for most of these option portfolios are low. Most importantly, the GDA-I model goes a long way in explaining the very high premium earned by the 30-day deep out-of-the-money (OTM) equity index put option portfolio, which has hitherto proven notoriously difficult to rationalize (see the discussion in Constantinides et al. (2013)).[14]

Contributing to this strand of the literature, we find that this 30-day OTM index put option portfolio is severely exposed to disappointment consumption risk, explaining why it commands a very high risk premium. A natural interpretation of this result is that deep OTM index put options provide insurance against disappointment consumption events, so a sufficiently disappointment-averse investor is willing to pay a high premium to the counterparty who provides this insurance (see Schreindorfer (2014) for a similar argument). In relative terms, the GDA-I model clearly outperforms the competing single-factor models. Again, the Fama–French (1993), (2015) models can only fit this cross section through factor coefficient estimates whose signs and magnitudes are strikingly inconsistent.

Finally, Panel D of Table 4 presents the model estimates for the joint cross section of equity, corporate bond, Treasury bond, and equity index option portfolios. Interestingly, we find that the GDA-I model can correctly price this joint cross section ($R^2 = 72\%$), yielding the best fit across the competing models (see also Panel D of Figure 3). In fact, the FF3 performs poorly because it

[14]Recall that, according to the definition of Constantinides et al. (2013), put portfolios have a short position in puts.

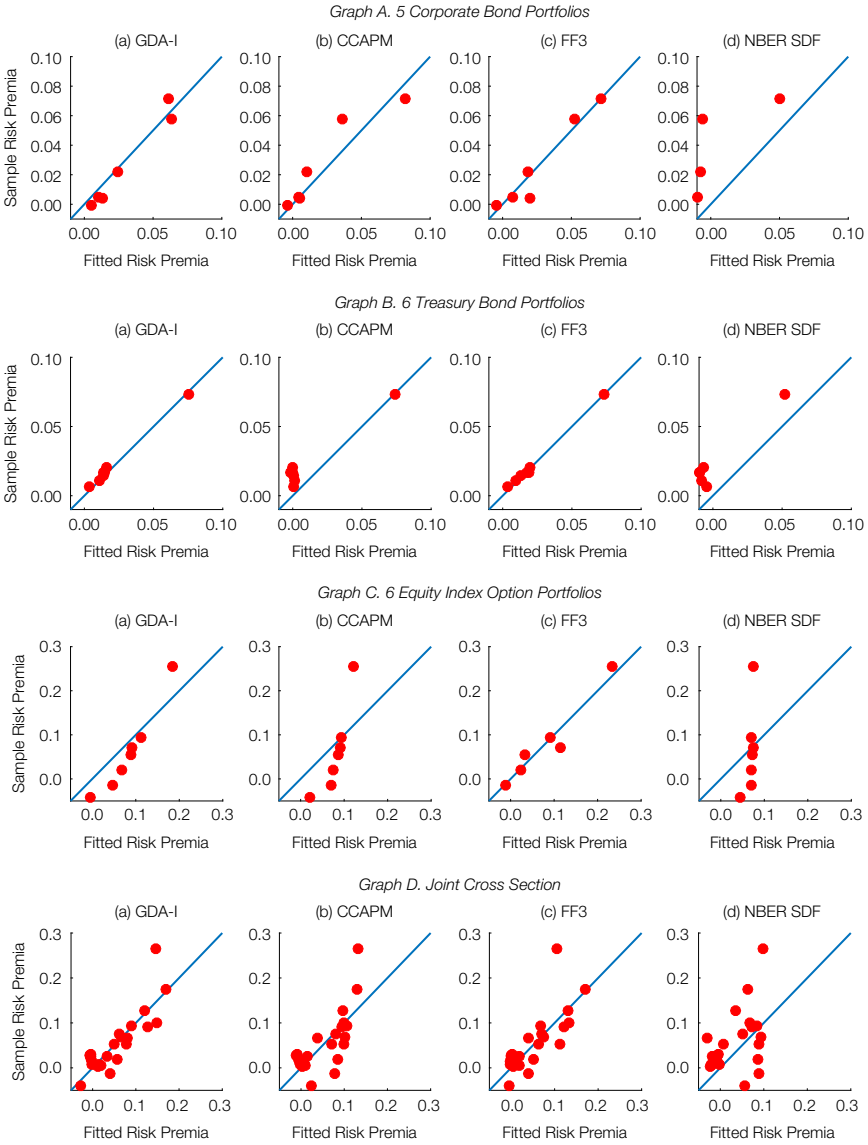
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FIGURE 3

Sample and Fitted Risk Premia for Alternative Asset Classes

Figure 3 shows sample and fitted annual risk premia for the five corporate bond portfolios of Nozawa (2012) constructed on the basis of bonds' credit ratings (Graph A), the six Fama Treasury bond portfolios sorted on maturity (Graph B), the six equity index option portfolios of Constantinides et al. (2013) (Graph C), and the joint cross section of the aforementioned portfolios together with the six Fama–French SIZE/BM portfolios (Graph D). Each cross section also includes the equity market portfolio. Fitted risk premia are estimated according to the expression in equation (12) for the generalized disappointment-aversion indicator (GDA-I), consumption-based capital asset pricing model (CCAPM), FF3, and National Bureau of Economic Research (NBER) discount factors. The corresponding estimation results are shown in Table 4. The sample period is 1976–2009 for the corporate bond portfolios, 1952–2012 for the Treasury bond portfolios, 1987–2011 for the equity index option portfolios, and 1987–2009 for the joint cross section.



cannot sufficiently explain these different asset classes with the same set of factor coefficient estimates. Even worse, the FF5 yields predominantly negative factor risk prices. To the contrary, the GDA-I model achieves a very good fit with a DA coefficient estimate that is very similar to the ones derived from the previously examined cross sections (estimate = 4.1) and implies a very reasonable risk-taking behavior. In conclusion, the notable pricing ability of the GDA-I model, and the stability of the DA coefficient estimates across different asset classes and subperiods, can be regarded as an external validation of our benchmark analysis.

In sum, we find that the GDA-I model clearly outperforms the competing single-factor models, whereas the multifactor Fama–French (1993), (2015) models only achieve a good fit at the expense of extreme factor coefficient estimates. Further, the DA coefficient estimates are remarkably stable across the different asset classes, and they are consistent with the estimates we derived from the equity portfolio cross sections.

VI. The DA-I Model with $\delta = 1$

Our single-factor model has been developed under the GDA framework of Routledge and Zin (2010), where the disappointment threshold is a multiple of the certainty equivalent ($\mathbf{1}\{V_{t+1} \leq \delta \mu_t(V_{t+1})\}$ in equation (1)), and it is determined by the GDA parameter δ. In the special case where δ is restricted to be equal to 1, the GDA framework becomes equivalent to the original DA framework of Gul (1991), where the disappointment threshold is equal to the certainty equivalent ($\mathbf{1}\{V_{t+1} \leq \mu_t(V_{t+1})\}$).

To assess the importance of δ being a free parameter in the GDA framework, this section examines a restricted version of the GDA-I SDF, setting δ equal to 1. We term the restricted disappointment model as the disappointment-aversion indicator (DA-I) SDF. The specification of the DA-I model and the empirical methodology used to estimate it are identical to those for the GDA-I model in equations (4) and (13). The key difference is that in the DA-I model, the corresponding disappointment threshold $d_2^{\text{DA-I}}$ in equations (4) and (5) cannot be estimated as a free parameter. Instead, $d_2^{\text{DA-I}}$ is an implicit nonlinear function of the consumption growth moments and the DA parameter (see equation (A.15) in the Supplementary Material). Hence, to compute $d_2^{\text{DA-I}}$ in the DA-I model, one needs to solve the corresponding fixed-point problem while estimating the unknown parameters (μ_c, ϕ_c, σ_c^2, and $\tilde{\theta}$).

Estimation results for the DA-I model are reported in Table 5. Panel A shows the results for the equity portfolio cross sections (SIZE/BM, SIZE/OP, SIZE/INV, LTR). We find that the DA coefficient estimates from the DA-I SDF are similar, albeit slightly lower, to the corresponding estimates reported in Table 2 for the GDA-I SDF. In addition, the threshold coefficient estimates $d_2^{\text{DA-I}}$ we back out from the DA-I model are typically lower than the ones directly estimated from the GDA-I specification, leading to a larger number of disappointment consumption events. For instance, in the cross section of the SIZE/BM portfolios, we get 20 disappointment consumption years from the DA-I model, whereas the GDA-I model yields only 13 such events (see Section VII.A). Finally, in terms of model

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TABLE 5

GMM Results for Annual Portfolio Returns:
DA-I Model with Restricted GDA Coefficient ($\delta = 1$)

Table 5 shows the generalized method of moments (GMM) results for the disappointment-aversion indicator (DA-I) model across different test assets at the annual frequency. For these tests, we estimate consumption growth moments and the disappointment-aversion (DA) coefficient $\tilde{\theta}$ for the DA-I model using the augmented GMM system specified in equation (13). The specification of the DA-I model is identical to that of the generalized disappointment-aversion indicator (GDA-I) model in equation (4). However, in the DA-I model, the disappointment threshold coefficient $d_2^{\text{DA-I}}$ in equations (4) and (5) is not estimated as a free parameter. Instead, it is computed as an implicit function of the DA coefficient and the consumption growth moments (see equation (A.15) in the Supplementary Material). Table 5 does not report estimation results for the consumption growth moments. In Panel A, we consider four equal-weighted portfolio sorts: the 25 size/book-to-market (SIZE/BM) portfolios, the 25 size/operating profitability (SIZE/OP) portfolios, the 25 size/investment (SIZE/INV) portfolios, and the 10 long-term reversal (LTR) portfolios. In Panel B, we consider the five corporate bond portfolios of Nozawa (2012) constructed on the basis of bonds' credit ratings, the six Fama Treasury bond portfolios sorted on maturity, the six equity index option portfolios of Constantinides et al. (2013), and a joint cross section of the aforementioned portfolios together with the six Fama–French equal-weighted size/book-to-market portfolios. In each of the cross sections in Panel B, we also include the equity market portfolio as a test asset. t-statistics are shown in parentheses. χ^2, DOF, and p are the first-stage χ^2 test, degrees of freedom, and p-value that all moment conditions are jointly 0. The results for the joint cross section in Panel B do not include t-statistics or p-values for the χ^2 test due to the limited time-series observations relative to the number of portfolios. RMSE and R^2 are the root-mean-square error (×100) and cross-sectional R-square, respectively. In Panel A, the sample period for the size/book-to-market and long-term reversal portfolios is 1933–2012, whereas the sample for the operating profitability and the investment portfolios begins in 1964. In Panel B, the sample period is 1976–2009 for the corporate bond portfolios, 1952–2012 for the Treasury bond portfolios, 1987–2011 for the equity index option portfolios, and 1987–2009 for the joint cross section.

Panel A. Equal-Weighted Equity Portfolios

	25 SIZE/BM	25 SIZE/OP	25 SIZE/INV	10 LTR
$\tilde{\theta}$	3.090	3.485	3.510	3.253
	(3.862)	(2.648)	(2.142)	(6.645)
$d_2^{\text{DA-I}}$	−0.548	−0.581	−0.583	−0.563
χ^2	35.598	43.429	79.456	13.470
DOF	24	24	24	9
p	0.060	0.008	0	0.142
RMSE	1.542	1.010	1.962	1.857
R^2	0.870	0.883	0.741	0.830

Panel B. Alternative Asset Classes

	5 Corporate Bond Portfolios	6 Treasury Bond Portfolios	6 Equity Index Option Portfolios	Joint Cross Section
$\tilde{\theta}$	3.561	4.776	3.912	2.897
	(8.159)	(2.192)	(1.362)	
$d_2^{\text{DA-I}}$	−0.587	−0.681	−0.616	−0.526
χ^2	6.612	14.533	8.833	29.487
DOF	5	6	6	23
p	0.251	0.024	0.183	
RMSE	1.584	0.242	5.457	3.907
R^2	0.681	0.986	0.633	0.630

fit, the results in Panel A of Table 5 show that when the test assets are equity portfolios, the empirical fit of the DA-I SDF is very similar to the fit of our benchmark GDA-I SDF.

In contrast, the estimation results for the alternative asset classes (corporate bond, Treasury bond, and equity index option portfolios) reported in Panel B of Table 5 indicate that, overall, the DA-I SDF performs worse than the GDA-I SDF. Moreover, the corresponding DA coefficient estimates from the DA-I model are typically lower than the ones reported in Table 4 for the GDA-I model. Similarly, the threshold coefficient values $d_2^{\text{DA-I}}$ we back out from the DA-I specification are mostly lower than the corresponding ones estimated from the GDA-I framework.

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The fact that the empirical fit of the DA-I model in the sample of equity portfolios is very similar to that of the GDA-I model leads to two important conclusions. First, the goodness of fit of the disappointment model in the equities sample is not particularly sensitive to small variations in the threshold coefficient. This finding is consistent with the sensitivity analysis we carry out in Section VIII.B. Second, as suggested by Routledge and Zin (2010), the identification of the GDA parameter δ is rather challenging. In other words, one cannot confidently argue whether δ is equal to 1 or not solely on the basis of comparing the cross-sectional fit of the GDA-I and DA-I models.

Based on the previous discussion, we consider the benchmark GDA-I model preferable to the DA-I model for three reasons. First, from a theoretical point of view, the flexibility of the GDA framework has been shown to be important in capturing a series of stylized facts in asset pricing (see Routledge and Zin (2010) and Bonomo et al. (2011)). Given the challenging identification of δ, there is no obvious benefit in restricting it to be exactly equal to 1. Second, from an empirical point of view, allowing δ to be a free parameter permits the estimation of the disappointment threshold coefficient d_2 as a free parameter and hence enables the GDA-I model to maintain its linearity with respect to the price of disappointment risk. This feature substantially facilitates its estimation. Last, the cross-sectional fit of the benchmark GDA-I SDF is overall better than that of the restricted DA-I specification, particularly for the alternative asset classes.

VII. Disappointment Events and High Marginal Utility States

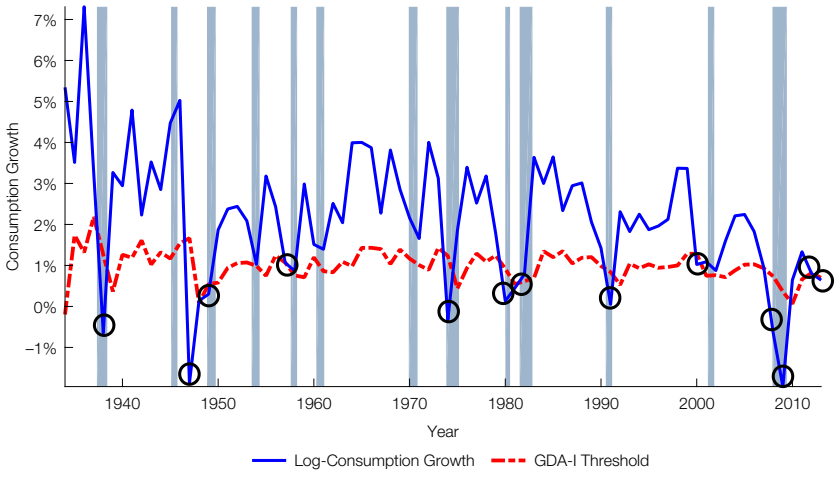
A. Disappointment Events and Their Characteristics

A successful consumption-based asset pricing model needs to accurately map "bad economic times" to states of high marginal utility for the representative investor. To this end, we examine how some key economic and financial variables behave during the disappointment consumption years extracted from the cross section of the 25 SIZE/BM portfolios (see Table 2). In particular, these disappointment years occurred in 1937, 1946, 1948, 1956, 1973, 1979–1980, 1990, 1999, 2007–2008, and 2011–2012, and they are illustrated in Figure 4. Table 6 reports the average values of these variables in the full sample period as well as during disappointment years. Moreover, to highlight that the identified disappointment events do not simply capture recession periods, we also report the average values of these variables during NBER recession years.

With respect to stock market performance, we find that during disappointment consumption years, the MARKET_PREMIUM is negative, the SMB_PREMIUM is highly negative, and the VALUE_PREMIUM is quite low relative to its full sample average. These effects are much less pronounced when we split the sample on the basis of NBER recession years, indicating that the stock market leads NBER recessions and that the disappointment indicator is successful in capturing this behavior. For instance, the HML_PREMIUM during NBER recessions (8.02%) is marginally different from the HML_PREMIUM during nonrecession years (9.28%). In contrast, the HML_PREMIUM during disappointment years is much lower than during nondisappointment periods (4.23% vs. 9.95%).

FIGURE 4
Consumption Growth, Disappointment Events, and NBER Recessions

Figure 4 shows annual disappointment events in consumption growth for the generalized disappointment-aversion indicator (GDA-I) discount factor of equation (4). Disappointment events are determined using the estimation results for the 25 size/book-to-market portfolios reported in Panel A of Table 2. The solid line denotes consumption growth, and the dashed line is the time-varying GDA-I certainty equivalent of consumption growth from equation (5). Disappointment events are highlighted by circles, and shaded areas are National Bureau of Economic Research (NBER) recessions. The sample period is 1933–2012.



Disappointment years are associated with a lower average TERM_SPREAD, supporting the argument that the disappointment indicator can be considered as a recession leading indicator. Verifying that disappointment years are "bad times" indeed, REAL_CONS_GROWTH_RATE is almost 0, and the corresponding EARNINGS_GROWTH_RATE is less than a third of its full sample average. Confirming that consumption is low during these years, the CAY variable of Lettau and Ludvigson (2001) takes a negative average value, and CONSUMER_CONFIDENCE is dramatically reduced. In contrast, we do not find disappointment years to be associated with a particularly negative INVESTOR_SENTIMENT, once this is orthogonalized with respect to macroeconomic conditions. Interestingly, NET_EQUITY_EXPANSION is also very low during disappointment years, indicating that corporations anticipate the economic slowdown. Further, we find that disappointment consumption years are also associated with a high and increasing INFLATION_RATE. Finally, supporting the argument that the disappointment indicator anticipates rather than coincides with recessions, we find that these years are followed by, but do not coincide with, a higher and increasing UNEMPLOYMENT_RATE.

B. Expected Returns According to the GDA-I Model

To explain the superior performance of the GDA-I model relative to the traditional CCAPM framework, we consider the expression for risk premia in equation (12). The poor performance of the CCAPM, especially at the monthly and quarterly frequencies (see Sections A5 and A6 in the Supplementary Material), is due to the inability of aggregate consumption growth to align asset

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TABLE 6

Key Economic Variables during Disappointment and Recession Years

Table 6 shows averages for key economic and financial variables during disappointment and National Bureau of Economic Research (NBER) recession years as well as over the full sample period, 1933–2012. The set of disappointment years is determined using the estimation results for the 25 size/book-to-market portfolios reported in Panel A of Table 2. The CAY variable of Lettau and Ludvigson (2001) is available since 1945, UNEMPLOYMENT_RATE is available since 1947, CONSUMER_CONFIDENCE is available since 1961, and MARKET_SENTIMENT is available since 1965. Factor returns are taken from Kenneth French's Web site. The economic and financial variables are computed using the updated data set of Welch and Goyal (2008), with the exception of the UNEMPLOYMENT_RATE, which is taken from the Bureau of Labor Statistics, and the REAL_CONS_GROWTH_RATE, which is from the Bureau of Economic Analysis (BEA). CONSUMER_CONFIDENCE is sourced from the University of Michigan survey of consumer confidence, and INVESTOR_SENTIMENT is the Baker and Wurgler (2006) index of investor sentiment orthogonalized with respect to a set of macroeconomic conditions.

		Years			
	Full Sample	Disappointment	Nondisappointment	NBER Recession	Nonrecession
MARKET_PREMIUM	9.16%	−3.05%	11.52%	1.93%	10.96%
SMB_PREMIUM	2.12%	−5.68%	3.63%	0.95%	2.41%
HML_PREMIUM	9.02%	4.23%	9.95%	8.02%	9.28%
S&P_500_DAILY_STANDARD_DEVIATION	15.43%	19.22%	14.70%	19.26%	14.48%
DEFAULT_SPREAD	1.12%	1.33%	1.08%	1.55%	1.01%
TERM_SPREAD	1.62%	0.99%	1.74%	1.83%	1.57%
REAL_CONS_GROWTH_RATE	2.21%	−0.07%	2.65%	1.47%	2.39%
EARNINGS_GROWTH_RATE	11.23%	3.20%	12.78%	4.65%	12.87%
NET_EQUITY_EXPANSION	1.52%	0.77%	1.66%	1.48%	1.53%
CAY	0.00%	−0.16%	0.03%	−0.08%	0.02%
INFLATION_RATE	3.68%	6.03%	3.23%	3.47%	3.74%
CHANGE_IN_INFLATION_RATE	0.15%	1.75%	−0.16%	−1.13%	0.47%
UNEMPLOYMENT_RATE	5.77%	5.72%	5.78%	6.16%	5.67%
CHANGE_IN_UNEMPLOYMENT_RATE	0.06%	−0.04%	0.09%	1.25%	−0.23%
UNEMPLOYMENT_RATE, $t+1$	5.77%	6.22%	5.68%	7.28%	5.43%
CHANGE_IN_UNEMPLOYMENT_RATE, $t+1$	0.06%	0.97%	−0.10%	1.12%	−0.20%
CONSUMER_CONFIDENCE	86.31	76.44	88.38	71.52	89.41
%_CHANGE_IN_CONSUMER_CONFIDENCE	0.16%	−7.49%	1.80%	−8.80%	2.08%
INVESTOR_SENTIMENT (orthogonalized)	0.012	−0.028	0.019	0.364	−0.074

returns with marginal utility. In contrast, the GDA-I factor is a nonlinear transformation of the consumption growth process, for which periods of low returns are aligned with periods of high marginal utility (disappointment consumption events).

More specifically, the GDA-I SDF in equation (4) implies that asset premia should be linearly related to their expected losses during disappointment consumption events. In particular, substituting equation (4) into equation (12), we get:

$$\begin{aligned}\mathbb{E}[R_{i,t} - R_{1\mathrm{Y},t}] &= -\frac{\tilde{\theta}}{1-\tilde{\theta}\mathbb{E}[\mathbf{1}_t]}\mathbb{E}[\mathbf{1}_t(R_{i,t} - R_{1\mathrm{Y},t})] \\ &= -\frac{\tilde{\theta}\mathbb{E}[\mathbf{1}_t]}{1-\tilde{\theta}\mathbb{E}[\mathbf{1}_t]}\mathbb{E}[(R_{i,t} - R_{1\mathrm{Y},t})|\mathbf{1}_t{=}1],\end{aligned}$$

where $\mathbf{1}_t$ is the disappointment indicator. By correctly identifying disappointment consumption events, the GDA-I model is able to align the full sample premia of risky assets with the average losses they incur during these events. To this end, Delikouras (2017) shows that the location of the reference point for gains and losses (i.e., the certainty equivalent of consumption growth) is crucial for this successful alignment.

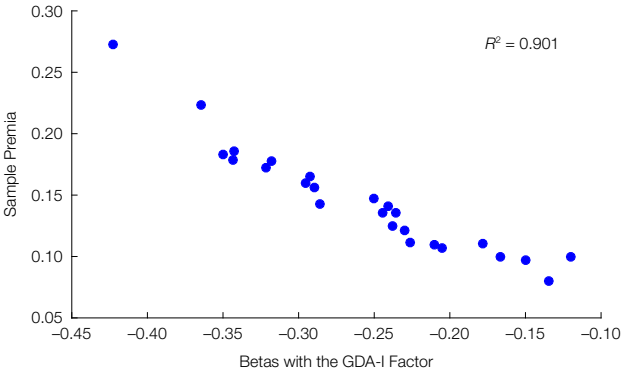
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Confirming these arguments, Figure 5 plots the full sample annual premia of the 25 SIZE/BM portfolios versus their full sample disappointment betas. These betas are estimated from a model where the only regressor is the disappointment indicator (GDA-I factor) extracted from the cross section of the 25 SIZE/BM portfolios in Table 2. In fact, Figure 5 shows that portfolio premia are almost perfectly aligned with their disappointment betas ($R^2 = 90.1\%$). In other words, the more sensitive portfolio returns are to the occurrence of a disappointment event (i.e., the greater its losses are during this period), the higher the premium that this portfolio yields.

FIGURE 5
Premia versus GDA-I Betas for the 25 SIZE/BM Portfolios

Figure 5 plots sample annual premia versus time-series betas estimated with respect to the generalized disappointment-aversion indicator (GDA-I) factor for the set of the 25 size/book-to-market (SIZE/BM) portfolios. The GDA-I factor is determined using the estimation results for the 25 size/book-to-market portfolios reported in Panel A of Table 2. R^2 is computed from the cross-sectional regression of portfolio premia on GDA-I factor betas without intercept. The sample period is 1933–2012.



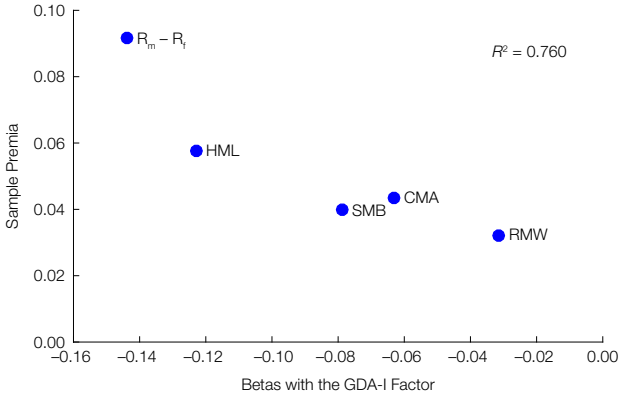
C. The GDA-I Model and the Fama–French Factor Premia

The results reported in Section IV show that the proposed GDA-I model can fit the premia of the size, value, profitability, and investment portfolio cross sections as accurately as the multifactor Fama–French (1993), (2015) models. This striking finding implies that the GDA-I SDF is a single-factor, consumption-based equivalent representation of the Fama–French multifactor SDFs. If this holds true, then the GDA-I model should also be able to explain the premia that the five Fama and French (2015) factors yield.

To examine the validity of this conjecture, Figure 6 plots the full sample premia of the market, SMB, HML, RMW, and CMA factors versus their disappointment betas. Again, these betas are estimated from a model where the only regressor is the disappointment indicator (GDA-I factor) extracted from the cross section of the 25 SIZE/BM portfolios in Table 2. In line with our conjecture, we find that the factor premia are well aligned with their GDA-I betas (R^2 = 76%). Moreover, the slope of the linear relationship between the factor premia and their GDA-I betas implies a DA coefficient of 4.2, which is very close to the corresponding estimates from the portfolio cross sections (see Table 2). These results

FIGURE 6
Premia versus GDA-I Betas for the Five Fama–French Factors

Figure 6 plots sample annual premia versus time-series betas estimated with respect to the generalized disappointment-aversion indicator (GDA-I) factor for the 5 Fama and French (2015) factors. The GDA-I factor is determined using the estimation results for the 25 size/book-to-market (SIZE/BM) portfolios reported in Panel A of Table 2. R^2 is computed from the cross-sectional regression of factor premia on GDA-I factor betas without intercept. The sample period for the excess market ($R_m - R_t$), small-minus-big (SMB), and high-minus-low (HML) factors is 1933–2012, whereas the sample period for the robust-minus-weak (RMW) and conservative-minus-aggressive (CMA) factors is 1964–2012.



point to the conclusion that the premia that the Fama–French (2015) factors yield reflect compensation for exposure to disappointment consumption risk.

VIII. Identification and Placebo Tests

A. Identification Tests for GDA-I Betas

Burnside (2011) and Bryzgalova (2015) convincingly show that if asset returns are weakly correlated with the candidate pricing factor, then the corresponding risk premium may be weakly identified, leading to spurious inference. To address the potential concern that the proposed GDA-I model may suffer from weak identification, we conduct a series of Wald tests regarding the joint significance and the cross-sectional dispersion of the GDA-I betas.

To this end, we jointly estimate GDA-I betas utilizing a system of seemingly unrelated regressions (SUR) of the 25 SIZE/BM portfolio excess returns as well as the market portfolio excess return on the disappointment indicator (GDA-I factor), which has been extracted from the cross section of the 25 SIZE/BM portfolios in Table 2. The point estimates of GDA-I betas for these 25 portfolios are illustrated in Figure 5. First, we test the null hypothesis that these 25 GDA-I betas are jointly equal to 0 ($H_0 : \widehat{\beta}_i^{\text{GDA-I}} = 0, \forall i$). Second, we test whether the 25 GDA-I betas are jointly equal to the GDA-I beta of the market portfolio ($H_0 : \widehat{\beta}_i^{\text{GDA-I}} = \widehat{\beta}_m^{\text{GDA-I}}, \forall i$). Third, we test whether the 25 GDA-I betas are jointly equal to their average estimate ($H_0 : \widehat{\beta}_i^{\text{GDA-I}} = \overline{\widehat{\beta}}^{\text{GDA-I}}, \forall i$). Finally, we test whether the GDA-I beta of the small/value portfolio (S1B5), which yields the highest premium (see Table 1), is equal to that of the big/growth portfolio (S5B1), which yields the lowest premium ($H_0 : \widehat{\beta}_{S1B5}^{\text{GDA-I}} = \widehat{\beta}_{S5B1}^{\text{GDA-I}}$).

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The Wald statistics reported in Table 7 indicate that we can reject each of these four null hypotheses at any conventional significance level. These results alleviate the potential concern that the GDA-I model may be weakly identified because the 25 SIZE/BM portfolio betas with respect to the GDA-I factor are both individually and jointly significant, and they exhibit significant cross-sectional dispersion in the statistical sense too. Hence, we conclude that the 25 SIZE/BM portfolio returns significantly covary with the GDA-I SDF and that these covariances exhibit a significant cross-sectional dispersion, reflecting the dispersion across portfolio premia.

TABLE 7
Identification Tests for GDA-I Betas

Table 7 presents the results from four Wald tests regarding the joint significance and the cross-sectional dispersion of generalized disappointment-aversion indicator (GDA-I) factor betas. We estimate a system of seemingly unrelated regressions (SUR) of the 25 size/book-to-market (SIZE/BM) portfolio excess returns as well as the market portfolio excess return on the GDA-I factor. The GDA-I factor is determined using the estimation results for the 25 SIZE/BM portfolios reported in Panel A of Table 2. Based on these estimates, we conduct the following Wald tests: i) we test whether the 25 portfolio betas are jointly equal to 0 ($H_0: \widehat{\beta}_i^{\text{GDA-I}} = 0, \forall i$), ii) we test whether the 25 portfolio GDA-I betas are jointly equal to the market portfolio beta ($H_0: \widehat{\beta}_i^{\text{GDA-I}} = \widehat{\beta}_m^{\text{GDA-I}}, \forall i$), iii) we test whether the 25 portfolio betas are jointly equal to their average estimate ($H_0: \widehat{\beta}_i^{\text{GDA-I}} = \overline{\widehat{\beta}}^{\text{GDA-I}}, \forall i$), and iv) we test whether the GDA-I beta of the small/value portfolio is equal to that of the big/growth portfolio ($H_0: \widehat{\beta}_{S1B5}^{\text{GDA-I}} = \widehat{\beta}_{S5B1}^{\text{GDA-I}}$). Wald and p denote the Wald statistic and the corresponding p-value, respectively. The sample period is 1933–2012.

	Null Hypothesis			
	$\widehat{\beta}_i^{\text{GDA-I}} = 0, \forall i$	$\widehat{\beta}_i^{\text{GDA-I}} = \widehat{\beta}_m^{\text{GDA-I}}, \forall i$	$\widehat{\beta}_i^{\text{GDA-I}} = \overline{\widehat{\beta}}^{\text{GDA-I}}, \forall i$	$\widehat{\beta}_{S1B5}^{\text{GDA-I}} = \widehat{\beta}_{S5B1}^{\text{GDA-I}}$
Wald	71.68	63.32	68.97	6.24
p	0.000	0.000	0.000	0.002

B. Sensitivity Analysis with Respect to the Disappointment Threshold Coefficient

In this section, we examine how sensitive the explanatory ability of the GDA-I model is with respect to the disappointment threshold coefficient d_2, which, together with the consumption growth moments, determines the GDA-I certainty equivalent for consumption growth (see equation (5)). To this end, we arbitrarily set this coefficient equal to $d_2 \pm \text{se}(d_2)$, based on the estimates of d_2 ($= -0.770$) and $\text{se}(d_2)$ ($=0.256$) for the 25 SIZE/BM portfolios from Panel A of Table 2. Each of these two alternative values for d_2 yields a different certainty equivalent and hence a different set of disappointment consumption events.

Using each of these two modified disappointment indicators, we reestimate the corresponding betas for the 25 SIZE/BM portfolios and reexamine their explanatory ability with respect to the annual portfolio premia. Panel A of Figure 7 contains the scatterplot of these portfolio premia versus their GDA-I betas, where the modified GDA-I factor has been determined using a disappointment threshold coefficient equal to $d_2 + \text{se}(d_2)$ ($=-0.513$). In this case, we get 21 disappointment consumption years in the 1933–2012 sample period, and the cross-sectional R^2 of the modified GDA-I betas with respect to the 25 SIZE/BM portfolio premia is 82.2%. Moreover, the implied DA coefficient is 2.9.

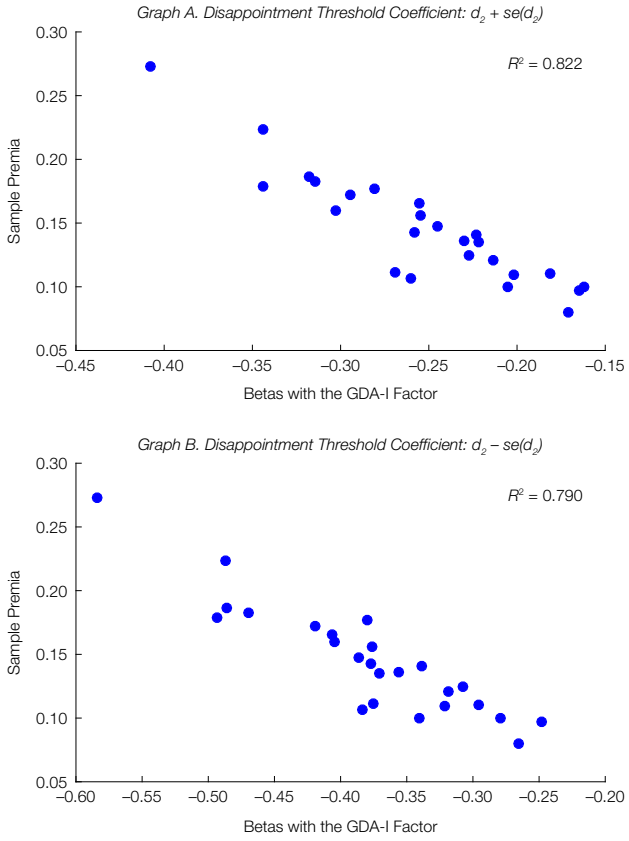
Panel B of Figure 7 contains the scatterplot of the 25 SIZE/BM portfolio premia versus their GDA-I betas, where the modified GDA-I factor has been determined using a disappointment threshold coefficient equal to $d_2 - \text{se}(d_2)$

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FIGURE 7

Sensitivity Analysis for the Disappointment Threshold Coefficient

Figure 7 plots sample annual premia versus time-series betas estimated with respect to a modified generalized disappointment-aversion indicator (GDA-I) factor for the set of 25 equal-weighted size/book-to-market (SIZE/BM) portfolios. In Graph A, the modified GDA-I factor is defined relative to the certainty equivalent for consumption growth given by equation (5) and is computed using a disappointment threshold coefficient equal to $d_2+\text{se}(d_2)$ $(=-0.513)$. In Graph B, the corresponding modified GDA-I factor is defined using a disappointment threshold coefficient equal to $d_2-\text{se}(d_2)$ $(=-1.026)$. The estimates of d_2 $(=-0.770)$ and $\text{se}(d_2)$ $(=0.256)$ are taken from Panel A of Table 2 for the 25 SIZE/BM portfolios. R^2 is computed from the cross-sectional regression of portfolio premia on GDA-I factor betas without intercept. The sample period is 1933–2012.



$(=-1.026)$. In this case, we get only 7 disappointment consumption years. For this set of disappointment events, the cross-sectional R^2 of the modified GDA-I betas with respect to the portfolio premia is 79%, and the implied DA coefficient is 4.8.

This sensitivity analysis leads to some interesting conclusions. First, the explanatory power of the GDA-I model in the SIZE/BM sample remains quite high for a wide range of values of the disappointment threshold coefficient. Therefore, its successful empirical performance in our benchmark results is not particularly sensitive to the characterization as disappointing or not of those consumption growth observations that lie marginally around the GDA-I certainty equivalent

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(see Figure 4). This result is also consistent with the findings in Section VI regarding the performance of the restricted DA-I model across equity portfolios.

Second, an alternative way to interpret these results is that the set of the 7 lowest consumption growth years (1937, 1946, 1973, 1979, 1990, and 2007–2008), which coincides with the set of the disappointment events derived when the threshold coefficient is equal to −1.026, goes a long way in explaining the cross-sectional dispersion in portfolio premia. Finally, the price of disappointment risk estimated from each of these two alternative sets of disappointment consumption events remains relatively stable around the benchmark estimates reported in Table 2.

It should also be noted that in Section A11 of the Supplementary Material, we conduct a number of further tests to examine the importance of key parameters of the GDA-I model. Specifically, to examine the importance of persistence in consumption growth, we first estimate a variant of the GDA-I model, where we impose an IID consumption growth process (i.e., $\phi_c = 0$ in equation (4)). Secondly, to assess the importance of the time variation in the GDA-I threshold, we alternatively estimate indicator models where the disappointment threshold is time invariant, and it is specified in an ad hoc fashion. In particular, we estimate an indicator model where the disappointment threshold is the unconditional mean of consumption growth ($\mathbf{1}\{\Delta c_{t+1} \leq \mu_c\}$) as well as a model where the threshold is 0 consumption growth ($\mathbf{1}\{\Delta c_{t+1} \leq 0\}$). Finally, to highlight the role of the DA parameter $\tilde{\theta}$, we also estimate a variant of the GDA-I model, where we restrict the price of disappointment risk to be equal to 1 (i.e., $\tilde{\theta} = 1$ in equation (4)).

The results from these tests show that these alternative models perform much worse relative to our benchmark GDA-I model, yielding higher average pricing errors and lower cross-sectional R^2s. Based on these findings, we conclude that the success of the GDA-I model should be attributed to a theoretically consistent time-varying reference point (see equation (5)) and a DA parameter that is greater than 1.

C. Placebo Indicators

Last but not least, we address the potential concern that the high explanatory power of the GDA-I model may be spuriously driven by the fact that the proposed factor is an indicator function. In particular, we examine whether a randomly generated (placebo) indicator could also yield a similarly high explanatory power and hence whether our benchmark results are driven by "luck." To this end, we generate 1,000,000 time series, each consisting of 80 observations drawn from a Bernoulli distribution, containing exactly 67 zeros and 13 ones, and hence characterized by the same mean and variance as the GDA-I factor extracted from the cross section of the 25 SIZE/BM portfolios in Table 2. Subsequently, we estimate the betas of the 25 SIZE/BM portfolios with respect to each of these placebo indicators and run a cross-sectional regression (without intercept) of the annual portfolio premia on the corresponding placebo indicator betas.

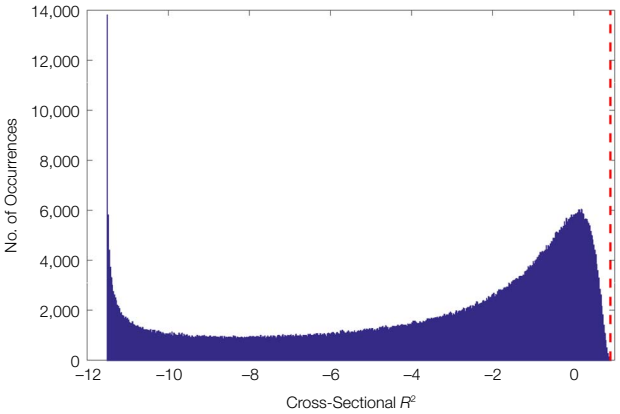
Figure 8 shows the histogram of the cross-sectional R^2s yielded by the placebo indicators, whereas the dashed vertical line indicates the R^2 (= 90.1%) of the actual GDA-I factor. It is evident that the placebo indicators have a dismally low explanatory power with respect to the 25 SIZE/BM portfolio premia.

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In fact, only 15 out of the 1,000,000 placebo indicators (= 0.0015%) yield a cross-sectional R^2 that is equal to or greater than the one of the GDA-I factor. Hence, we can confidently reject at any conventional significance level the null hypothesis that the explanatory power of the proposed factor is driven by "luck."[15]

FIGURE 8
Explanatory Power of Placebo Indicators

Figure 8 shows the histogram of cross-sectional R^2s yielded by a no-intercept regression of the annual premia of the 25 equal-weighted size/book-to-market (SIZE/BM) portfolios on their betas estimated with respect to a placebo indicator. Each of the 1,000,000 placebo indicators that we generate consists of 80 observations drawn from a Bernoulli distribution, containing exactly 67 zeros and 13 ones, and thus is characterized by the same mean and variance as the generalized disappointment-aversion indicator (GDA-I) factor extracted from the cross section of the 25 SIZE/BM portfolios. The vertical dashed line indicates the corresponding cross-sectional R^2 yielded by the actual GDA-I factor ($R^2 = 0.901$).



IX. Conclusion

Kocherlakota (1996) argues that to improve the performance of consumption-based models (at least) one of the following three assumptions needs to be relaxed: i) CRRA preferences, ii) market completeness, iii) transaction costs. In this article, we relax the CRRA assumption and propose a single-factor asset pricing model based on an indicator function of aggregate consumption growth being less than its certainty equivalent. This certainty equivalent is derived from generalized disappointment-aversion preferences, and it is located approximately 1 standard deviation below the expected consumption growth.

Our single-factor model can sufficiently explain the cross section of expected returns for the size, value, profitability, investment, reversal, corporate bond, Treasury bond, and equity index option portfolios. In terms of relative

[15] Another potential concern is that the disappointment indicator may be highly persistent, leading to spurious inference. The autocorrelation coefficient estimate for the disappointment indicator extracted from the cross section of the 25 SIZE/BM portfolios in Table 2 is 0.09 (p-value = 0.395). Similar results hold for the other cross sections, indicating that the GDA-I factor exhibits insignificant persistence.

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performance, the proposed model outperforms traditional asset pricing models (CCAPM, CAPM). Moreover, the fit of our model is at least as good as the fit of the FF3, and its performance is comparable to that of the FF5. In fact, our single-factor model can also explain the premia that the 5 Fama–French (2015) factors yield. Finally, the estimated prices of disappointment risk are plausible and much more stable across test assets and frequencies relative to the risk-aversion coefficient estimates derived from the CCAPM.

Collectively, our results indicate that risk premia are, to a large extent, compensation for exposure to disappointment (i.e., downside) consumption risk, and they challenge the overreaction hypothesis suggested by a number of previous studies. Our findings also highlight that disappointment aversion plays a crucial role in understanding the cross section of expected returns from a consumption-based perspective, and they question the effectiveness of the second-order risk-aversion framework that is commonly used in consumption-based asset pricing.

Supplementary Material

Supplementary Material for this article is available at https://doi.org/10.1017/S0022109018000819.

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