Financial Integration and the Equity Term Structure: Insights from 150 Years of UK Data

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Abstract

We recover the term structure of equity risk premia in the UK from 1870 to 2019 and develop an international asset pricing model to interpret the results. Four main findings emerge. First, risk premia have declined markedly since the 1980s, particularly at short maturities, leading to a steeper equity term structure in the post-1980 period. Second, over the same time period, correlations in risk premia between the UK and the U.S. have risen substantially, particularly at short horizons. Third, viewed over the full 150-year sample, the equity term structure is upward-sloping on average, and investors demand higher compensation for bearing long-term than short-term equity risk. Fourth, the term structure flattens during periods of financial distress, driven by a disproportionate increase in short-term risk premia. Our international asset-pricing model attributes the post-1980 steepening of the term structure to greater financial integration, via enhanced cross-border diversification, which reduces short-term risk premia more than long-term premia.

JEL Codes: D51, E21, G12, G13, G15

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1 Introduction

The equity term structure is fundamental to asset pricing, portfolio choice and firm valuation because it determines how cash flows of different maturities are discounted. Yet empirical evidence on its long-run behavior remains scarce because existing data are short and concentrated around a few recent crises. For instance, forward equity yields—the sum of maturity-specific dividend-growth expectations and risk premia—are typically computed from dividend futures (Van Binsbergen et al., 2013; Van Binsbergen and Koijen, 2017; Gormsen, 2021; Bansal et al., 2021), which exist only for the last two decades. Although recent work has broadened the available evidence by leveraging the cross-section of firms (Giglio et al., 2024), the combined coverage still represents only a limited fraction of modern financial history, as Figure 1 shows. Cochrane (2017) emphasizes that the "facts" about the equity term structure are contentious precisely because the available data are short and clustered around crisis periods.

We make three contributions. First, we construct a new 150-year dataset of equity yields for the UK and recover the whole term structure of equity risk premia using the method of Giglio et al. (2024). Second, we document four key facts about the term structure of equity risk premia, including a sharp post-1980 decline—especially at short maturities—that coincides with the 1979 abolition of UK capital controls and the rise in financial market integration. We also provide 150 years of evidence on its level, slope, and time variation. Third, we develop an international asset-pricing model that reproduces the key facts of the equity term structure and use it to identify a new channel through which financial integration shapes the term structure of risk premia.

Our analysis starts by testing whether the methodology of Giglio et al. (2024)—which recovers equity yields from the cross-section of firms—is applicable outside the U.S. We find that the model-implied yields closely track those extracted from FTSE 100 Dividend Index Futures, the most liquid non-U.S. dividend contract. Moreover, the model's recovered dividend-growth expectations exhibit pronounced co-movement with major macroeconomic and financial episodes, including the Great Depression, the postwar expansion, the 1970s inflation shock, and the Global Financial Crisis. The tight correspondence between model-implied and market-based equity yields, together with the alignment between model-implied growth expectations and realized economic conditions, strengthens confidence that the resulting equity risk premia—defined as the difference between forward yields and expected dividend growth—provide a credible representation of reality.

Our 150-year sample of the equity term structure delivers new insights and allows precise inference about previously debated "facts." A key result is that hold-to-maturity (HTM) risk premia decline sharply after 1980, with the drop most pronounced at short maturities, producing a much steeper equity term structure in the post-1980 period.[1] Over the same horizon, the contribution of risk-premia variation to short-term equity yield fluctuations falls substantially, whereas for long-maturity yields, the relative roles of discount rates and growth expectations remain essentially unchanged. In addition, UK–U.S. correlations in risk premia rise markedly after 1980—again most strongly at short maturities—indicating tighter cross-country co-movement in short-horizon premia. These shifts align closely with the 1979 abolition of UK capital controls, which ended the postwar regime of financial segmentation.[2] The post-1980 period is therefore characterized by deeper financial integration, and we use our model to show that enhanced international risk sharing—particularly for short-run risks—plays a central role in reshaping the equity term structure.

We make two contributions to the debate on the slope and cyclicality of the equity term structure. First, we find that the term structure of HTM equity risk premia is upward sloping on average: long-horizon dividend strips earn higher returns, with the average premium on a 30-year strip about 5%, compared with about 2.5% for a 1-year strip. Second, we document a clear cyclical pattern: during recessions, short-maturity risk premia rise more than long-

[1]Throughout the paper, we focus on HTM premia because they map directly into dividend-strip discount rates and avoid the measurement issues affecting one-period returns on long-maturity strips (Bansal et al., 2021). See Kvaerner et al. (2025) for a detailed reconciliation of HTM and one-period risk premia.

[2]Before 1979, regulations restricted cross-border capital flows, and the sterling was non-convertible for capital-account transactions. Appendix D provides supporting evidence for the 1980 breakpoint.

maturity premia, which produces a sharp flattening of the HTM term structure. This pattern reflects the procyclical nature of dividend-growth expectations, which fall in downturns and rise in expansions. Our findings are consistent with those of Bansal et al. (2021), who rely on futures prices and a shorter U.S. sample. In Kvaerner et al. (2025), we provide a detailed reconciliation of the main U.S. evidence on the equity term structure, with a focus on hold-to-maturity and one-period returns.

To understand the dynamics of the term structure of equity premia and the sharp decline in short-term risk premia since the 1980s, we develop an international asset-pricing model with Epstein-Zin preferences, building on Colacito and Croce (2013).[3] Our model integrates country-specific disaster risks alongside a global long-run risk factor in the spirit of Bansal and Yaron (2004); Barro (2006); Wachter (2013). This allows us to analyze differences in equity yields, growth expectations, and risk premia across maturities under financial autarky versus full international integration. Opening to global capital markets substantially lowers short-term risk premia as well as their variation by enhancing international risk sharing. In contrast, long-term premia remain less affected by market integration due to shared exposure to persistent global shocks.

Quantitatively, the model replicates key empirical features of the data: an upward-sloping equity term structure; a disproportionate rise in short-maturity risk premia during periods of financial distress; and a sharp post-1980 decline in both the level and volatility of short-horizon premia driven by financial market integration. At the same time, it matches standard asset-pricing moments—including the market risk premium, a low risk-free rate, and a volatile price–dividend ratio—under conventional parameter values for preferences and endowments. Thus, the model's predictions for the equity term structure emerge entirely from standard parameter choices together with our novel risk-sharing channel. Taken together, the results show that financial integration has shifted the distribution of risk across

[3]Methodologically, we separate measurement from interpretation: a reduced-form SDF recovers dividend-strip prices and premia under no-arbitrage, and a distinct structural kernel is used to rationalize the measured term-structure facts; alignment is evaluated at the level of implied moments.

time, with important implications for long-horizon investors, sovereign wealth funds, and policymakers assessing the cost of capital in globally integrated markets.

Our paper is related to several strands of literature. First, we build on the literature that applies affine models to understand the risk-return properties of financial assets (Lemke and Werner, 2009; Cochrane and Piazzesi, 2005; Giglio et al., 2024). We closely follow the setup in Giglio et al. (2024) but apply it to a new dataset of the cross-section of UK stocks from 1870 until 2019. The central idea is to use both time-series and cross-sectional variation in equity portfolios to identify a set of model parameters that makes it possible to value any equity portfolio "dividend-by-dividend" in the spirit of Brennan (1998). Previous literature has also attempted to measure the term structure of equity premia from the cross-section of equities (Bansal et al., 2005; Hansen et al., 2008; Lettau and Wachter, 2007).[4]

Second, we contribute to the literature aiming to understand forward equity yields and their components following the seminal work by Van Binsbergen et al. (2012).[5] The somewhat "irregular" sample of mostly US data, covering the last 20 years, with frequent economic crises, has led to a dispute about the dynamics of forward equity yields. Critics of the "new facts," which are inconsistent with a host of asset pricing models, argue that the sample is unrepresentative and inference from derivative contracts is unreliable due to low market liquidity (Kirshon, 2020; Bansal et al., 2021). In addition, the prices of these contracts may be influenced by the views and preferences of financial intermediaries, rather than the preferences of the aggregate household sector—the decision-making unit in most asset pricing models. Basing our inference on listed equities—possible due to Giglio et al. (2024)—addresses these concerns.

Third, we contribute to the literature on using equilibrium models in the spirit of Lucas to understand financial market data. Specifically, we propose an international asset pricing model with Epstein–Zin preferences as in Colacito and Croce (2013), and Colacito et al.

[4]The idea is to exploit information in portfolios of stocks with different cash-flow growth and discount rate properties. Related work studying the term-structure of the equity risk premium using firm-level measures of duration includes: Weber (2018); Jankauskas et al. (2021); Gormsen and Lazarus (2023).

[5]For a literature review, see Van Binsbergen and Koijen (2017).

(2018) and incorporate short-term, country-specific disaster risks—as in Barro (2006) and Wachter (2013)—together with a global common risk factor that drives the underlying disaster intensities.[6] We show that the interplay of the two sources of risks allows us to replicate key empirical patterns of the term structure of equity risk premia documented in this paper. In particular, we show the degree of financial market integration significantly influences both the level and the slope of the equity term structure. Moreover, our model successfully matches not only the unconditional and conditional moments of the equity term structure but also standard moments of cash flows and asset prices.

Fourth, our work contributes to the literature on financial market integration and its asset-pricing implications. We introduce the maturity dimension of equity risk premia into the study of integration. Using a 150-year UK sample, we show that the abolition of capital controls in 1979 coincides with a structural decline in short-maturity premia, which steepens the equity term structure. This complements evidence that the early 1980s mark a shift toward deeper cross-border integration, captured by legal liberalization and de facto price and quantity co-movements.[7]

The paper is organized as follows. Section 2 describes the data sources, vairable construction and estimation procedure. Section 3 reports novel facts on the term structure of equity risk premia using 150 years of UK data. Section 4 introduces the theoretical framework and interprets the empirical findings through the lens of the model. Section 5 concludes.

[6] Wu (2025) proposes a complementary rare-disaster mechanism in which time-varying post-disaster recovery generates rich unconditional and conditional equity term structures for dividend strips. In contrast, our framework emphasizes how financial integration reallocates local disaster risk and global long-run risk across maturities.

[7] Legal openness increased during the 1980s and 1990s, as codified in current- and capital-account indices (Quinn and Inclan, 1997; Quinn, 1997). De facto integration rose with faster external adjustment after the collapse of Bretton Woods (Taylor, 2002) and discrete post-1970 shifts in key moments—higher real-exchange-rate volatility, a lower uncovered-interest-parity slope, and a lower Backus–Smith correlation (Colacito and Croce, 2013). Capital-account liberalization is also associated with higher international equity-return comovement (Quinn and Voth, 2008).

2 Data and Methodology

In this section, we describe the historical data sources and the empirical methodology used to construct the term structure of equity risk premia and growth expectations in the UK from 1870 to 2019.

2.1 Data

We combine historical and recent data to construct a long-run UK dataset of equities, interest rates, and macroeconomic aggregates from 1870 to 2019. UK equity data come from Global Financial Data (GFD) and Datastream (DS), which we merge by hand.[8] Short-term interest rates and government bond yields from GFD, DS, and the Bank of England's "A Millennium of Macroeconomic Data" are used to construct both the risk-free rate and the term structure of interest rates.[9] Real GDP and real consumption growth for the UK are obtained from the Bank of England. We construct a recession index that uses real GDP growth whenever available, and a market-based proxy when they are not.[10] Dividend futures for the UK are taken from DS. For complementary analysis and robustness checks, we use U.S. equity market returns and dividend yields from CRSP, real consumption from NIPA tables, and NBER recession dates. We construct nine equity anomalies for the model estimation. These anomalies are: Dividend Yield (divp), Dividend Yield (Monthly) (divpm), Dividend Yield-Momentum (divpmom), Industry Momentum (indmom), 6-Month Momentum (mom), 12-Month Momentum (mom12), Size (Size), Beta Arbitrage (betaarb), and Composite Issuance (ciss).[11]

[8] Appendix A.1 describes in detail the construction of the cross-section and the GFD–DS merge.

[9] Appendix A.4 details the construction of the short-term risk-free rate; Appendix A.5 describes how the term-structure of interest rates is computed.

[10] Real quarterly GDP growth is available from 1955 onward; we convert this series to monthly frequency to construct the recession indicator. For the pre-1955 period, we rely on a monthly market-based signal (e.g., market dividend yield). Appendix A.3 describes the construction of the UK recession index in detail.

[11] Details on the construction of these characteristics are in Appendix A.2.

2.2 Empirical Model

We adopt the empirical model developed by Giglio et al. (2024) (henceforth GKK) to recover the equity term structure in the UK from 1870 to 2019. The model estimates the term structure of expected market returns by decomposing dividend yields into components reflecting discount rates and expected dividend growth at different maturities. Our contribution is empirical: we apply the GKK framework to a newly constructed UK dataset with long historical coverage. This allows us to examine the behavior of the equity term structure over 150 years, including its level and slope across different historical episodes. In what follows, we summarize the core features of the GKK model that are essential for interpreting our main results and refer to Appendix B for further details.

The central object of the model is the state vector F_t, which governs both the stochastic discount factor (SDF) and risk premia. It contains $k = 2p$ variables: the log returns and log dividend yields of p portfolios. Following GKK, we set $p = 4$. The state vector evolves according to a first-order VAR:

$$F_{t+1} = c + \rho F_t + u_{t+1}. \tag{1}$$

where u_{t+1} is normally distributed with conditional covariance Σ. We construct F_t using the UK market portfolio and three Principal Component (PC) portfolios obtained from a balanced panel of 18 long-short portfolios based on the 9 UK anomalies described in Section 2.1. The prices of risk λ_t are affine functions of the state vector

$$\lambda_t = \lambda + \Lambda F_t, \tag{2}$$

and the SDF takes a log-linear form:

$$m_{t+1} = -r_{f,t} - \frac{1}{2}\lambda_t' \Sigma \lambda_t - \lambda_t' u_{t+1}. \tag{3}$$

Asset prices satisfy the Euler equation

$$1 = E_t\left[\exp\left(m_{t+1}\right)\left(\frac{P_{i,t+1}+D_{i,t+1}}{P_{i,t}}\right)\right]$$
$$= E_t\left[\exp\left(m_{t+1}+\Delta p_{i,t+1}+y_{i,t+1}\right)\right], \tag{4}$$

where $\Delta p_{i,t+1} \equiv \ln(P_{i,t+1}/P_{i,t})$ is the log price change and we define $y_{i,t+1} \equiv \ln(1+D_{i,t+1}/P_{i,t+1})$ as the log dividend yield. Under log-normality, it follows that

$$0 = E_t[m_{t+1}] + E_t[\Delta p_{i,t+1}] + E_t[y_{i,t+1}] + \frac{1}{2}\mathrm{Var}_t[m_{t+1} + \Delta p_{i,t+1} + y_{i,t+1}]. \tag{5}$$

The evolution of log excess price changes is directly specified as

$$\Delta p_{i,t+1} - r_{f,t} = \gamma_{i,0} + \gamma_{i,1}F_t + \gamma_{i,2}u_{t+1} + \varepsilon^p_{i,t+1}. \tag{6}$$

where $\varepsilon^p_{i,t+1}$ are asset specific shocks. Substituting this expression into the Euler equation implies that the log dividend yield can be expressed as an affine function of the state vector and asset-specific residuals:

$$y_{i,t} = b_{i,0} + b_{i,1}F_t + \varepsilon^y_{i,t}. \tag{7}$$

As a result, excess returns (i.e., the sum of price changes and dividend yields) are also affine in factors and shocks:

$$r_{i,t+1} - r^f_t = \beta_{i,0} + \beta_{i,1}F_t + \beta_{i,2}u_{t+1} + \varepsilon^r_{i,t+1}, \tag{8}$$

where $\varepsilon^r_{i,t+1} = \varepsilon^p_{i,t+1} + \varepsilon^y_{i,t+1}$ and the coefficients $\beta_{i,0}, \beta_{i,1}, \beta_{i,2}$ are functions of the underlying parameters in the log price and dividend yield equations.

Next, we define the key quantities of interest implied by the model: the HTM risk premia $\theta^{(n)}_t$, and the HTM expected dividend growth $g^{(n)}_t$.[12] Let $P^{(n)}_t$ denote the price at time t of the

[12] While these quantities can be computed for each portfolio in the system, our focus is on the equity term

dividend strip on the market portfolio that pays D_{t+n} at time $t+n$. The total ex-dividend price of the market portfolio is then the sum of all such strips:

$$P_t = \sum_{n=1}^{\infty} P_t^{(n)}.$$

We define the HTM risk premium, $\theta_t^{(n)}$, as the annualized log expected excess return from holding the n-period dividend strip to maturity:

$$\theta_t^{(n)} \equiv \frac{1}{n} \ln \mathbb{E}_t \left[\frac{R_{t:t+n}^{(n)}}{R_{f,t:t+n}} \right].$$

We define the HTM expected dividend growth, $g_t^{(n)}$, as the annualized log expected dividend growth, in excess of the risk-free rate, over the same horizon:

$$g_t^{(n)} \equiv \frac{1}{n} \ln \mathbb{E}_t \left[\frac{D_{t+n}}{D_t} \frac{1}{R_{f,t:t+n}} \right].$$

These two quantities are related through an identity. To see this, start from the realized gross return on the dividend strip, in excess of the risk-free rate:

$$\frac{R_{t:t+n}^{(n)}}{R_{f,t:t+n}} = \frac{D_{t+n}}{P_t^{(n)}} \frac{1}{R_{f,t:t+n}} = \frac{D_t}{P_t^{(n)}} \frac{D_{t+n}}{D_t} \frac{1}{R_{f,t:t+n}}.$$

Taking conditional expectations at time t, applying logs, and dividing by n, we get:

structure of the aggregate market. Accordingly, when referring to these quantities, we omit the portfolio index i from the notation; such references should be understood as applying to the market portfolio unless otherwise noted.

$$
\begin{aligned}
\mathbb{E}_t\left[\frac{R_{t:t+n}^{(n)}}{R_{f,t:t+n}}\right] &= \frac{D_t}{P_t^{(n)}} \cdot \mathbb{E}_t\left[\frac{D_{t+n}}{D_t}\frac{1}{R_{f,t:t+n}}\right] \\
\underbrace{\frac{1}{n}\ln \mathbb{E}_t\left[\frac{R_{t:t+n}^{(n)}}{R_{f,t:t+n}}\right]}_{\theta_t^{(n)}} &= \underbrace{\frac{1}{n}\ln\left(\frac{D_t}{P_t^{(n)}}\right)}_{\equiv e_t^{(n)}} + \underbrace{\frac{1}{n}\ln \mathbb{E}_t\left[\frac{D_{t+n}}{D_t}\frac{1}{R_{f,t:t+n}}\right]}_{g_t^{(n)}}.
\end{aligned} \tag{9}
$$

Here $e_t^{(n)}$ denotes the equity yield, the main object of interest in prior work such as Van Binsbergen et al. (2013); Van Binsbergen and Koijen (2017), particularly because it can be directly inferred from traded dividend futures data. Equation (9) shows that the HTM risk premium equals the sum of the equity yield and the HTM expected dividend growth over the same horizon. In Appendix B.2, we show how each component in this decomposition maps to the model's state vector, structural shocks, and parameters.

Throughout the paper, we focus on HTM risk premia rather than holding-period returns, which have been the focus of other papers. HTM premia offer several advantages for our analysis. First, they map directly into the discount rates of dividend strips and are therefore the relevant object for valuing long-term dividend claims. Second, one-period returns on long-horizon dividend strips are not observed each period and possibly include liquidity premia and market microstructure noise (Bansal et al., 2021). Third, one-period expected returns for long-maturity strips are susceptible to short-run movements in discount rates, making them volatile and challenging to estimate. For a detailed reconciliation of one-period and HTM returns, see Kvaerner et al. (2025).

3 Empirical Results

In this section, we present new evidence on the term structure of equity risk premia using 150 years of UK data and the methodology described above. We begin in Section 3.1 by validating the empirical performance of the model, comparing model-implied forward equity yields with those extracted from traded dividend futures—providing a direct test of the model's ability to capture maturity-specific pricing of dividend claims. We then examine the time-series behavior of the model-implied dividend-growth expectations and HTM risk premia, relating their dynamics to findings in the existing literature and documenting how they evolve over the long historical sample. Section 3.2 contains the main findings reporting novel evidence on the level, slope and time variation of the equity term structure.

3.1 Validation of the GKK Framework: Implied Equity Yields and 150 Years of Term-Structure Evidence

Recall that the estimation does not incorporate data from dividend futures, so there is no mechanical link between the model's outputs and the observed forward yields. This absence of circularity ensures that the comparison offers a meaningful validation of the model. In the main text, we focus on the 5-year equity yield. The results for the 2- and 7-year yields are in Appendix F.[13] Figure 2 plots the model-implied forward equity yield alongside the equity forward yield derived from traded futures contracts. The two series exhibit tight co-movement—both in levels and, more importantly, in their responses to shocks. The similarity between the model-implied yields and those based on traded futures extends to the 2- and 7-year yields.

[Insert Figure 2 here]

This validation exercise aligns with the findings of GKK, who show that their model—estimated yields using a range of U.S. equity portfolios—closely tracks the market prices of S&P 500 dividend strips. Taken together, these results highlight the model's ability to predict equity yields without relying on data on dividend futures.

To further assess the validity of our empirical estimates, we analyze the time-series evolution of dividend growth expectations and HTM risk premia using our long-run UK dataset.

[13]The maturities are chosen to facilitate direct comparison with prior studies (e.g., Gormsen, 2021).

Figure 3 plots the 1-year and 10-year expected dividend growth rate from 1870 to 2019, together with major historical recessions, financial panics, and global conflicts.

[Insert Figure 3 here]

The estimated dividend growth expectations align closely with major macroeconomic episodes, providing economically grounded support for the validity of our framework. Expected dividends decline sharply during periods of pronounced economic stress, including the Great Depression, World Wars I and II, the 1973 oil crisis, the dot-com bust, and the Global Financial Crisis of 2008–2009. In contrast, they rise during post-crisis recoveries, such as the early 1950s and the late 1990s. Moreover, we observe that short-term growth expectations decline during recessions, while investors anticipate long-term growth recovery. This implies a steeper term-structure slope during recessions, which aligns with the results of Bansal et al. (2021), who also find a countercyclical slope of the term structure of expected dividend growth.

Several of our findings resonate with those of Campbell et al. (2013), who distinguish between stock market downturns driven by rising discount rates and those caused by deteriorating cash flow expectations. In their decomposition of returns, declines in expected cash flows are interpreted as "hard times" for long-term investors. Our model captures these episodes—most notably the Great Depression and the Global Financial Crisis—as periods marked by sharp declines in expected dividend growth. This suggests that the model successfully identifies not only cyclical fluctuations but also structural shocks that are central to understanding asset prices.

Figure 4 plots the 1-year (blue line) and 10-year (orange line) HTM equity risk premia, overlaid with shaded regions denoting major financial, economic, and geopolitical events.

[Insert Figure 4 here]

Throughout most of the sample period, long-term HTM risk premia exceed their short-term counterparts, averaging around 4.5% for 10-year dividend strips and 2.5% for 1-year strips. These levels also vary substantially over time. For example during the Great Depression, the 1973 oil crisis and the Global Financial Crisis of 2008–2009, risk premia increase. During most periods of financial distress, we observe a sharp increase in short-term HTM risk premia, while long-term premia increase to a lesser extent. This indicates that short-term equity risk premiums are more responsive to adverse economic conditions.

Taken together, the strong co-movement between model-implied variables and key macroeconomic disruptions supports the model's external validity and its ability to capture the main forces behind equity market behavior during periods of economic stress.

3.2 New Facts on the Term Structure of Equity Risk Premia

In what follows, we document four facts about the term structure of the equity risk premia. First, risk premia have declined significantly since the 1980s, particularly at short maturities, producing a steeper equity term structure in the post-1980 period. At the same time, the share of short-term yield variation attributable to discount rates has declined markedly, whereas the contributions of cash-flow and discount-rate variation to long-maturity yields remain largely unchanged across the pre- and post-1980 samples. Second, correlations in risk premia between the UK and the U.S. have increased substantially over the same period, particularly at the short end. Third, across the full 150-year sample, the equity term structure is generally upward-sloping: investors require higher compensation for bearing long-term relative to short-term equity risk. Fourth, during episodes of financial distress, short-term risk premia rise disproportionately, causing the term structure to flatten, or even invert. The first two facts are novel and provide the foundation for analyzing how rising financial market integration has shaped the term structure of equity risk premia over time. Facts 3 and 4 reconcile and extend prior empirical findings on the term structure of equity risk premia using our extended UK dataset.

We begin by documenting unconditional evidence on the term structure of equity risk premia and its variation in recessions and expansions. Figure 5 presents the estimated term structure of the HTM equity risk premium. Panel a) reports estimates based on the full sample, while Panel b) distinguishes between expansions and recessions.

[Insert Figure 5 here]

Panel a) of Figure 5 shows that the HTM equity term structure is, on average, upward sloping, indicating that investors demand higher compensation for bearing long-term equity risk. The expected return on the 1-year dividend strip is about 2.8%, while 30-year strips require a premium of roughly 4.5%. The upward sloping term-structure of HTM equity risk premia documented here is consistent with standard macro-finance models featuring long-run risk (e.g., Bansal and Yaron, 2004), habit preferences (e.g., Campbell and Cochrane, 1999) or time-varying rare disaster risk (e.g., Wachter, 2013).

Panel b) of Figure 5 shows the HTM term structure of the equity risk premium conditional on the state of the economy. During recessions, the entire term structure shifts upward, reflecting higher required compensation for risk. Conversely, during expansions, risk premia are lower across all maturities. The difference between good and bad times is most pronounced at short horizons, while the long-end is less affected by the state of the economy. Hence, in line with the results of Berg (2014) and Bansal et al. (2021) for the US, we find that the slope of the HTM term structure of equity risk premia decreases in bad times in the UK. Note that our results are robust to alternative definitions of recession periods and essentially unchanged when using a recession index based on the price-dividend ratio (see Figure 16 in Appendix F).

Figure 6 compares the term structure of equity risk premia before and after 1980.[14] We split the sample at this point to distinguish between the UK economy before and after the Thatcher government dismantled Britain's remaining capital controls in 1979 (see Cole and

[14] See Appendix D for the economic motivation for this benchmark, robustness to nearby cutoffs (1970 and 1990), and formal evidence from Chow tests around 1980 and Bai–Perron multiple-break tests that place a data-driven break in the early 1980s.

Obstfeld, 1991; Quinn, 1997; Taylor, 2002; Obstfeld and Taylor, 2004; Bekaert et al., 2005, 2006, 2007; Eichengreen, 2008). Prior to this reform, capital controls—implemented under the Exchange Control Act of 1947—effectively segmented domestic financial markets from global capital markets. Cross-border asset flows were tightly restricted, and the authorities kept the pound sterling non-convertible for capital account transactions, even as trade in goods remained largely open.

[Insert Figure 6 here]

The top row of Figure 6 displays the (unconditional) term structures for the pre-1980 period (1870–1979) and the post-1980 period (1980–2019). The term structure was significantly higher across all maturities before 1980. This suggests that investors required greater compensation for risk in the earlier period. The right panel of the top row shows the difference between the two term structures (1870–1979 minus 1980–2019). The gap is largest at short maturities and narrows with horizon. Hence, the decline in risk premia over time has been concentrated at short to intermediate maturities, while long-term equity risk premia have remained relatively stable since the 1870s.

A possible concern is that this pattern mainly reflects differences in the frequency of expansions and recessions across the two periods. To show that this is not the case, the second and third rows of Figure 6 condition on expansions and recessions, respectively. We find that conditioning on either expansions or recessions leaves the difference between the 1870–1979 and 1980–2019 samples essentially unchanged. Our result also holds under alternative definitions of recessions, see Appendix F Figure 17.

The 1980 breakpoint affects not only the level and slope of the equity term structure but potentially also the cross-country co-movement of risk premia. To examine this, we analyze correlations in risk premia between the UK and the U.S. at each maturity. Specifically, for each horizon n, we estimate the standardized regression

$$\theta_{\mathrm{UK},t}^{(n)} = \alpha_n + \beta_n \, \theta_{\mathrm{US},t}^{(n)} + \varepsilon_t^{(n)}, \tag{10}$$

where $\theta_{\mathrm{US},t}^{(n)}$ denotes the HTM risk premium in the U.S. We obtain data on $\theta_{\mathrm{US},t}^{(n)}$ from Kvaerner et al. (2025) who extend the estimates in GKK back to 1937. All regressors and regressands are standardized within each regression, so β_n measures the correlation between UK and U.S. risk premia with horizon n. We use Newey–West standard errors with 12 lags. Regressions are estimated separately in the pre-1980 (1937–1979) and post-1980 (1980–2018) samples, and the figures report β_n together with 95% confidence intervals. Due to data availability in the U.S. sample, the analysis starts only in 1937.

Figure 7 reports estimates of cross-country comovement in HTM risk premia, measured by β_n across maturities. We find a pronounced rise in correlations since 1980. Before 1980, risk premia were essentially uncorrelated between the UK and the U.S. In contrast, in the post-1980 period, correlations increased sharply, especially for the short- to medium-maturity premia (1 to 10 years).

[Insert Figure 7 here]

We end this section with a variance decomposition of equity yields in the spirit of Van Binsbergen et al. (2013). Since the dividend yield aggregates equity yields across all horizons, decomposing their variance into discount-rate variation, expected cash-flow growth variation, and their interaction provides a horizon-by-horizon view of the forces underlying movements in dividend yields. This allows us to evaluate how the contributions of discount rates and expected cash-flow growth differ across maturities and whether their relative importance differs pre- and post-1980.

Recall equity yields satisfy $e_t^{(n)} = \theta_t^{(n)} - g_t^{(n)}$, where $\theta_t^{(n)}$ is the HTM risk premium and

$g_t^{(n)}$ is expected dividend growth. Thus, the variance of $e_t^{(n)}$ admits the exact decomposition

$$\mathrm{Var}(e_t^{(n)}) = \mathrm{Var}(\theta_t^{(n)}) + \mathrm{Var}(g_t^{(n)}) - 2\ \mathrm{Cov}(\theta_t^{(n)}, g_t^{(n)}), \tag{11}$$

which isolates the contributions of risk premia, expected cash-flow growth, and their comovement towards the variance of equity yields. Table 1 reports the variance decomposition.

[Insert Table 1 here]

Panels A–C show that at the 10- and 30-year maturities, the three samples deliver the same broad message: roughly one-half of the variance of equity yields reflects the joint movement of risk premia and expected dividend growth, and the remaining variance comes from the separate contributions of risk-premium variation and variation in expected dividend growth. While in the pre-1980 sample (Panel B), risk-premium variation contributes somewhat more than expected-growth variation, overall, the relative importance of each term remains very similar across the whole sample (Panel A), the pre-1980 period, and the post-1980 period (Panel C).

In contrast, for short maturities, the relative importance of each term in explaining the variance of equity yields changes between periods. At the one-year horizon, Panels B and C show a change between the pre- and post-1980 periods. After 1980, variation in risk premia accounts for only about 4% of the variance of equity yields (versus roughly 18% before 1980), variation in expected dividend growth accounts for nearly twice as much as in the earlier period (around 80% versus 50%). The covariance term accounts for roughly half of its pre-1980 contribution (about 16% versus 34%).

4 Theoretical Analysis

To understand the mechanisms behind the evolution of the term structure of expected dividend growth and risk premia, we develop an international asset pricing model with Epstein–

Zin preference as in Colacito and Croce (2013) (henceforth CC). Our model incorporates country-specific disaster risk—building on Barro (2006) and Wachter (2013)—alongside a global common long-run risk factor. To discipline the model, we adopt parameter values that are widely used in the asset-pricing literature. The resulting predictions for the level, slope, and dynamics of the equity term structure follow entirely from this standard calibration in conjunction with the added risk-sharing mechanism embedded in our framework. In what follows, we present the structure of the model, outline its core mechanisms, and revisit the empirical evidence on the term structure of equity risk premia through the lens of the model.[15]

4.1 The Model

The economy consists of two countries, home (h) and foreign (f) with two goods. As in CC, we denote by X the home good of country h and by Y the home good of country f. Aggregate consumption in each country is defined over the two goods as:

$$C_t^h = (x_t^h)^\alpha (y_t^h)^{1-\alpha} \quad \text{and} \quad C_t^f = (x_t^f)^{1-\alpha} (y_t^f)^\alpha, \tag{12}$$

where x_t^i and y_t^i denote consumption of good X and Y in country $i \in \{h, f\}$. The parameter $\alpha \in [0, 1]$ captures the degree of home bias in consumption preferences. A value of $\alpha > 0.5$ implies that consumers in each country have a preference for domestically produced goods.

Endowments. Each country is endowed with a stochastic supply of its home good. The

[15]Our framework separates measurement from interpretation. The empirical equity term structure (from the previous section) is recovered with a reduced-form SDF that imposes no-arbitrage on the payoff span of our equity returns and yields. The theoretical analysis from this section uses a structural SDF delivered by the model to study how autarky versus integration and disaster risk shape the equity term structure. Consistency is judged by comparing implications for the term structure—its average downward slope, its cyclical variation, and the post-1980 change in its slope and volatilities.

endowments follow a cointegrated process:

$$\begin{aligned}
\log X_{t+1} &= \log X_t + \mu_x + \tau(\log Y_t - \log X_t) + J_{x,t+1} + \epsilon_{x,t+1} \\
\log Y_{t+1} &= \log Y_t + \mu_y - \tau(\log Y_t - \log X_t) + J_{y,t+1} + \epsilon_{y,t+1} \\
J_{j,t+1} &= N_{j,t+1} d, \quad j \in x, y.
\end{aligned} \tag{13}$$

Here, $\tau \in (0, 1)$ is a cointegration parameter and $\epsilon_{j,t+1}$ are normally distributed country-specific short-run shocks with volatility σ_j and correlation ρ_{xy}.[16] μ_i denotes average log endowment growth. For simplicity, we set $\mu_x = \mu_y = \mu$.

We introduce short-term, country-specific disaster risks $J_{j,t+1}$ into the economy and show that they play a central role in how financial integration shapes the term structure of equity risk premia. Let $d < 0$ denote the size of the downward jump in consumption. The number of disasters follows a country-specific Poisson counting process $N_{j,t+1}$ with a global common risk factor λ_t. The risk factor λ_t drives the disaster intensities in both countries and follows a discrete-time version of a CIR process as in Wachter (2013):

$$\lambda_{t+1} = \bar{\lambda}(1 - \rho_l) + \rho_l \lambda_t + \phi_l \sqrt{\lambda_t} \omega_{t+1},$$

with $\omega_{t+1} \sim$ i.i.d. $N(0, \sigma_z)$. The idea is that, while disasters may be country-specific, a high value of λ_t reflects a period of global financial stress in which disaster risk increases uniformly across countries. From equation (13), it follows that the cointegration term $CI_t = (\log Y_t - \log X_t)$ enters the model as a state variable with the following dynamics:

$$CI_{t+1} = (1 - 2\tau) CI_t + J_{y,t+1} - J_{x,t+1} + \epsilon_{y,t+1} - \epsilon_{x,t+1}. \tag{14}$$

[16]Note that $\tau \in (0, 1)$ is necessary to ensure the existence of an equilibrium; otherwise, the endowments of the two goods diverge (see Colacito et al., 2019). We follow Colacito and Croce (2013) and use a small positive value for τ, which ensures that cointegration does not quantitatively affect our results.

Preferences. The representative household in each country has Epstein–Zin preferences:

$$U_t^i = \left[(1-\delta)(C_t^i)^{1-\frac{1}{\psi}} + \delta E_t\left((U_{t+1}^i)^{1-\gamma}\right)^{\frac{1-\frac{1}{\psi}}{1-\gamma}}\right]^{\frac{1}{1-\frac{1}{\psi}}}, \quad i \in \{h, f\}, \tag{15}$$

where δ denotes subjective discount factor, ψ measures the intertemporal elasticity of substitution (IES) and γ is the coefficient of relative risk aversion. Preference parameters are identical across countries.

Market structure and financial risk sharing. To analyze how financial market integration and international risk sharing affect the term structure of equity premia, we consider two market environments.

Case 1: Financial autarky. We first consider an environment with free trade in goods markets but no financial market integration. The idea is to mimic key features of the UK before the Thatcher government abolished capital controls in 1979. In this environment, households can trade the two consumption goods, but cannot invest in foreign bonds or equities, nor in more general state-contingent claims that pay only in particular future states. UK residents hold only UK assets, foreign residents hold only foreign assets, and no cross-border contract can deliver extra resources exactly when a domestic shock hits a country. As a result, each country must absorb and insure against unexpected changes in domestic income and output via domestic assets only. In this environment, the consumption share of each country—the fraction of world consumption that its residents absorb in each period—is fixed and pinned down by the degree of home bias α:

$$\begin{aligned} \frac{x_t^h}{X_t} &= \alpha, & \frac{y_t^h}{Y_t} &= (1-\alpha) \\ \frac{x_t^f}{X_t} &= (1-\alpha), & \frac{y_t^f}{Y_t} &= \alpha. \end{aligned}$$

Goods trade adjusts each period to maintain these home-biased consumption shares, but, in the absence of international trade in future consumption claims, households cannot shift consumption across countries or states to hedge bad domestic outcomes.

Case 2: Complete markets. In our second environment, we allow full financial market integration, which mimics the global economy since the 1980s. Markets are complete so that households can trade a full set of contingent claims to future consumption across countries. A convenient way to describe the allocation is through a time-varying Pareto weight S_t. A global planner chooses state-contingent consumption for home and foreign households to maximise a weighted sum of their lifetime utilities, subject to global resource constraints. S_t measures how much the planner values an extra unit of home consumption relative to an extra unit of foreign consumption at time t. When $S_t > 1$, the planner places more weight on home utility, so home households receive a larger share of world consumption.

In a decentralised economy with complete markets, the same allocation arises from competitive trade in international financial assets. We can therefore interpret S_t as a reduced-form summary of the relative wealth and risk tolerance of home and foreign investors, as implied by equilibrium prices and portfolios. In this setting, the optimal international risk-sharing conditions imply that home and foreign consumption shares of the two goods satisfy:

$$\begin{aligned}
\frac{x_t^h}{X_t} =& \alpha \left[1 + \frac{(1-\alpha)(S_t - 1)}{1 - \alpha + \alpha S_t}\right], \qquad \frac{y_t^h}{Y_t} = (1-\alpha)\left[1 + \frac{\alpha(S_t - 1)}{\alpha + (1-\alpha)S_t}\right] \\
\frac{x_t^f}{X_t} =& (1-\alpha)\left[1 - \frac{\alpha(S_t - 1)}{1 - \alpha + \alpha S_t}\right], \quad \frac{y_t^f}{Y_t} = \alpha \left[1 - \frac{(1-\alpha)(S_t - 1)}{\alpha + (1-\alpha)S_t}\right].
\end{aligned}$$

These risk-sharing conditions imply a simple insurance mechanism. Before shocks hit, home and foreign households trade state-contingent claims: the home country gives up some consumption in states where its economy is strong in return for extra consumption in states where it is weak. When a negative shock strikes the home country, its marginal utility rises and the effective Pareto weight S_t tilts in its favor, so international financial markets transfer resources from foreign to home investors. In that state, the home country absorbs a larger share of both the home and the foreign good than under financial autarky, so home consumption no longer moves one-for-one with a bad domestic shock.

Consumption growth in the home country is given by:

$$\frac{C_{t+1}^h}{C_t^h} = \left(\frac{x_{t+1}^h}{x_t^h}\right)^{\alpha} \left(\frac{y_{t+1}^h}{y_t^h}\right)^{1-\alpha} = \underbrace{\left(\frac{\frac{x_{t+1}^h}{X_{t+1}}}{\frac{x_t^h}{X_t}}\right)^{\alpha} \left(\frac{\frac{y_{t+1}^h}{Y_{t+1}}}{\frac{y_t^h}{Y_t}}\right)^{1-\alpha}}_{\text{Growth due to risk sharing}} \underbrace{e^{\alpha(\Delta X_{t+1})+(1-\alpha)(\Delta Y_{t+1})}}_{\text{Autarky growth}}$$

where $\Delta X_{t+1} = \log X_{t+1} - \log X_t$. The first part describes the change in the consumption share due to risk sharing, and the second part denotes consumption growth in the financial autarky case. An analogous expression holds for the foreign country.

Equilibrium. First order optimality implies that Pareto weights S_t evolve as:

$$S_{t+1} = S_t \frac{M_{t+1}^h}{M_{t+1}^f} \frac{e^{\Delta c_{t+1}^h}}{e^{\Delta c_{t+1}^f}}, \tag{16}$$

where $\Delta c_{t+1}^i = \log \frac{C_{t+1}^h}{C_t^h}$ denotes the log growth rate of consumption in country i and the stochastic discount factor in country i, M_{t+1}^i, implied by Epstein–Zin preferences is given by:

$$M_{t+1}^i = \delta \left(\frac{C_{t+1}^i}{C_t^i}\right)^{-\frac{1}{\psi}} \left(\frac{U_{t+1}^i}{E_t\left((U_{t+1}^i)^{1-\gamma}\right)^{\frac{1}{1-\gamma}}}\right)^{\frac{1}{\psi}-\gamma}. \tag{17}$$

The relative share of world consumption of the home country, W_t^h, depends on the Pareto weight S_t via:

$$W_t^h = \frac{S_t}{1+S_t}. \tag{18}$$

Hence, the time variation in W_t^h summarizes how international risk sharing reallocates world consumption between the two countries over time.

Equity markets. Dividends represent a levered claim to consumption in the respective country as in Bansal and Yaron (2004) and Wachter (2013). As a result, log dividend growth of the market portfolio is given by:

$$\Delta d_{t+1}^i = \mu_d + k(\Delta c_{t+1}^i - \mu). \tag{19}$$

As in the data, we decompose the risk premium of the market portfolio into strips with different maturities. To compute the HTM risk premium for each strip, we use the decomposition in equation (9). Specifically, we first compute equity yields and get n-period dividend growth expectations $g_t^{(n)}$ from simulations. The HTM risk premium $\theta_t^{(n)}$ is the sum of the two.

We solve the model globally using the collocation projection method of Pohl et al. (2021). They show that a global method is necessary for models that include persistent long-run shocks, as we have. For that, we first solve for the equilibrium value functions of the two countries and optimal consumption allocations as functions of the state variables $\{W_t^h, \lambda_t, CI_t\}$. To obtain a stationary representation, we detrend all non-stationary variables by multiplying with $\frac{1}{X_t^\alpha Y_t^{1-\alpha}}$. Once we have solved for the equilibrium, we use the stochastic discount factor in equation (17) to compute the price-to-dividend ratio of the market portfolio and the equity yields for both countries. Additional details on the implementation and numerical procedure are provided in Appendix E.

4.2 Risk Sharing and the Term Structure of Equity Risk Premia

We next show how financial integration shapes equity risk premia across horizons. We first compare how local and global shocks affect the economy under different degrees of risk sharing. Throughout this Section, we use the calibration in Table 2. The left panel of Figure 8 plots the change in the fraction of total world consumption that goes to the home country, $W_{t+1}^h - W_t^h$, against the global shock ω_{t+1}. We set $W_t^h = 0.5$, so the two countries start with the same consumption share. In this symmetric case, both countries demand the same extra return per unit of global risk, because ω_{t+1} shifts local disaster intensities in the same way in both countries. As a result, there is no gain from reallocating resources across countries in response to a global shock, and consumption shares remain constant.

In contrast, a consumption disaster in the home country leads to a rise in its share of total world consumption. In the right panel of Figure 8, we show this by plotting $W_{t+1}^h - W_t^h$

against local disaster shocks $J_{x,t+1}$. Because the home country dislikes this local disaster risk most, it is willing to forgo more consumption in normal times in exchange for extra consumption during the disaster. The foreign country takes the other side of this trade: it receives additional consumption in normal times and, in return, transfers resources to the home country in disaster states. When a local shock occurs, these state-contingent transfers raise the home country's consumption share. Thus, financial risk sharing dampens the effects of local shocks on the home country's consumption.

[Insert Figure 8 here]

How does the risk-sharing mechanism affect the term structure of equity risk premia? Before we elaborate on the risk-sharing mechanism below, we note that in the absence of global long-run risks (i.e. $\lambda_t = \lambda \; \forall \; t$), risk premia on short- and long-maturity claims would coincide, resulting in a flat term structure of the equity risk premia. Long-run risks, however, raise the risk premia on long-maturity claims and therefore produce an upward-sloping equity term structure; see Gormsen (2021). Financial market integration enables the sharing of short-run risks while leaving exposure to long-run shocks unchanged. As a result, risk sharing primarily reduces risk premia on short-maturity claims, steepening the equity term structure—consistent with the post-1980 evidence reported in Section 3.2.

We illustrate this mechanism in Figure 9, which plots the 2-year and 30-year hold-to-maturity equity risk premia as a function of the home country's share of world consumption, W_t^h. When $W_t^h = 1$, the home country owns the entire world economy, so there is no foreign partner to share risk with; this corresponds to financial autarky. As W_t^h falls below one, the scope for international risk sharing begins. More risk sharing lowers required risk premia, because part of the aggregate risk can be shifted to foreign investors. This effect is powerful for short-term risk premia: with integrated markets, investors can partly hedge country-specific disaster risk by trading claims across borders, as discussed above. In contrast, long-term risk premia move little with W_t^h. In our model, far-ahead cash flows depend mainly on global, slow-moving shocks that push down world consumption in both countries simultaneously.

Since these shocks leave the distribution of world consumption across countries essentially unchanged, reallocating consumption between home and foreign investors does little to insure against them. Hence, the compensation required for long-run risk is almost independent of W_t^h.

[Insert Figure 9 here]

In what follows, we demonstrate how this mechanism can qualitatively and quantitatively account for the post-1980 changes in the equity term structure documented in Section 3.2.

4.3 Calibration and Simulated Moments

We first describe the model calibration and then show that it replicates standard moments of consumption and dividend growth, as well as key features of asset prices. The calibrated parameter values are in Table 2.

[Insert Table 2 here]

For the preference parameters, we adopt standard values from the long-run risk literature: a relative risk aversion of 10, an intertemporal elasticity of substitution (IES) of 1.5, and an annual subjective discount factor of 0.98 (Bansal and Yaron, 2004; Bansal et al., 2012). We calibrate the means of the endowment processes to match standard moments of consumption and dividend growth, and we set the volatility of the i.i.d. normal shocks to zero to isolate the impact of global, long-run shocks and short-run disaster risks on the term structure of equity risk premia. We adopt the values for the home bias and cointegration parameter from Colacito and Croce (2013). The parameters governing the disaster processes are in line with the values reported in Barro (2006) and Wachter (2013). For example, Barro (2006) reports an average disaster size of approximately 30%, while Wachter (2013) uses an average disaster probability of 3.55%, a persistence of 0.92, and a volatility of 0.067 for the disaster intensity process. In our calibration, we use a slightly lower volatility, as higher values render the discrete-time CIR process ill-defined and lead to numerical convergence issues.

Table 3 reports standard moments of consumption growth, dividend growth, and asset prices for the UK for the sample periods from 1870-1979 and 1980-2019. We refer to the UK as the home country and the US as the foreign country, and therefore rely on the UK market portfolio and the risk-free rate for asset-pricing moments, as described in Section 2. To obtain the corresponding model-implied moments, we simulate 10,000 samples, each of the same length as the empirical dataset, and report the median, along with the 5% and 95% quantiles, across the simulated samples. To reduce the influence of initial conditions in our simulations, we apply a 1000-year burn-in period for all state variables. The autarky specification corresponds to Case 1, with free trade in goods markets but no trading of future consumption claims. The financial integration specification corresponds to Case 2, with complete markets. In the simulations, we assume an initial home-country share of world consumption of 3%, which approximately matches the UK's average share of global GDP over the 1980–2019 period.

[Insert Table 3 here]

The model matches standard moments of cash flows and asset prices. It reproduces both the average equity risk premium of approximately 5% and the low risk-free rate. The model's average consumption and dividend growth are broadly consistent with empirical growth rates, which range from -0.05% to 2.16% across subperiods.[17] Despite omitting short-run i.i.d. shocks, the volatilities of consumption and dividend growth in the data still fall within the 5%–95% quantiles of the simulated moments.

4.4 Revisiting Economic History

We now use the model to understand the historical evolution of the term structure of equity risk premium and growth expectations. Figure 10 plots n-period expected dividend growth (left), equity yields (center), and equity risk premia (right) in good and bad times. Solid

[17] The negative average dividend growth rate in the pre-1980 sample is primarily driven by the effects of the First and Second World Wars.

lines depict model-implied outcomes. The dashed lines show the corresponding empirical quantities. Good and bad times in the data are defined as in Section 3. In the model, bad (good) times are identified as periods when the price-dividend ratio falls below (rises above) its 25th percentile.[18]

[Insert Figure 10 here]

Several observations merit attention. First, the term structure of expected dividend growth varies systematically with the business cycle, consistent with empirical evidence. In particular, the model generates an upward-sloping term structure of expected dividend growth during recessions and a downward-sloping term structure during expansions. In downturns, heightened global disaster intensity increases the likelihood of local disasters, thereby dampening short-term growth expectations. Still, households expect the economy to recover over time, so they believe dividend growth will improve in the long run. In contrast, during booms, households expect strong short-term growth but also recognize that this pace cannot last forever. As a result, they expect dividend growth to slow down over time and return to typical levels.

Second, equity risk premia are high during periods of greater likelihood of global disasters and low when expected growth rates are high. The term structure of equity risk premia flattens during recessions. The reason is that investors demand substantial compensation for short-term disaster risk when global disaster probabilities rise, but they expect these probabilities to revert toward their average level over longer horizons. As a result, long-term risk premia increase less than short-term risk premia, producing a flattening of the equity term structure during recessions. These model-based predictions closely align with the empirical facts presented in Section 2. As shown in Figure 18 in Appendix F, all our qualitative model predictions become more pronounced when economic conditions are classified more stringently. Specifically, the figure shows the term structure of growth expectations, equity yields

[18]Note that as in Section 4.3, we compute term structures by simulating 10,000 samples, each with the same length as the empirical dataset and assuming an initial consumption allocation of $W_t^h = 0.03$. We obtain estimates for the term structures by averaging over sample paths.

and the equity risk premia conditional on the likelihood of global disasters λ_t. Increases in λ_t increase the slope of growth expectations, while it decreases the slope of the term structure of equity risk premia.

Putting these elements together, our model implies an upward-sloping equity yield forward curve in good times and a (slightly) downward-sloping curve in bad times. In expansions, the term structure of expected dividend growth slopes upward, while risk premia are relatively flat across maturities, leading to equity yields increasing with maturity. In recessions, however, equity yields invert: short-term growth expectations fall sharply, and investors demand higher compensation for short-term risk, resulting in both low expected cash flows and high discount rates at the short end of the yield curve. The net effect is that equity yields decrease with maturity.

4.5 Financial Integration and the Term Structure of Risk Premia

As shown in Section 3.2, short-term risk premia are significantly lower in the post-1980 period compared to the pre-1980 period. This shift has flattened the term structure of equity risk premia. We now employ the model to show—both qualitatively and quantitatively—how financial integration has shaped the slope of the equity risk premium term structure.

Figure 11 plots the term structure of equity risk premia under different regimes. Dashed lines show the empirical term structures; solid lines report their model-implied counterparts. For the empirical data, term structures are plotted in blue for the period 1870–1979 and in red for 1980–2019. In the model, we compare a regime of financial autarky (blue) with a regime of full financial market integration (red).

The model successfully replicates the observed levels of equity risk premia across both sample periods. Financial market integration reduces risk premia, consistent with empirical evidence. The model also captures the corresponding change in the shape of the term structure. Under financial integration, short-maturity premia decline more sharply than long-maturity premia, thereby steepening the term structure. As shown in the right panel of Figure 11, the reduction in premia is roughly 1.1% for short-term claims but only about 0.4% for long-term claims.[19]

[Insert Figure 11 here]

Figure 12 plots the equity term structures in the financial integration and autarky regimes, conditional on recessions and expansions. In the model, risk sharing is strongest during recessions, leading to a somewhat larger reduction in short-maturity premia under financial integration. Yet, consistent with the data, the magnitudes across regimes remain similar, and the model quantitatively matches the conditional patterns observed in the data.

[Insert Figure 12 here]

As shown in Section 3.2, the post-1980 steepening of the equity term structure also affects the variance of risk premia and their contribution to overall equity-yield volatility. We revisit these patterns using our model and show that the proposed risk-sharing mechanism aligns with the observed empirical changes. To illustrate this, Figure 13 reports the variances of equity yields $e_t^{(n)}$, HTM risk premia $\theta_t^{(n)}$, and dividend growth expectations $g_t^{(n)}$ in the model under both autarky and financial integration. The left panels display the absolute variances across maturities. In contrast, the right panels show the contribution of variation in risk premia and dividend growth to the variance of equity yields, following the same decomposition as in Table (1).

Financial integration reduces the variance of short-term risk premia, whereas the variance of long-term risk premia is barely affected. This implies a shift in the relative contributions of discount-rate and cash-flow variation to the overall volatility of equity yields. Under financial autarky, fluctuations in short-maturity discount rates are the dominant source of variation in short-term equity yields. With financial integration, this effect weakens, and variation in expected dividend growth becomes the dominant driver of short-term yield fluctuations.

[19] As already noted in Section 4.4, the model does not fully match the steepness of the short end of the equity term structure.

In sum, the model not only matches the time-series properties of the equity term structure, but also introduces a novel mechanism linking international financial integration to the pricing of long- and short-maturity equity claims. We demonstrate that financial integration—and the associated risk-sharing dynamics—can quantitatively replicate the post-1980 changes of the equity term structure, offering a new unified explanation for how global financial markets shape equity risk premia across maturities.[20]

5 Conclusion

We study the term structure of equity risk premia in the U.K. using a new 150-year dataset constructed with the method of Giglio et al. (2024). We establish four new empirical facts. First, the HTM equity term structure is upward sloping on average: long-maturity dividend strips earn persistently higher premia than short-maturity strips. Second, the slope is highly cyclical, flattening in recessions as short-term premia spike disproportionately. Third, we uncover a structural break around 1980, when short-maturity premia collapse, generating a much steeper equity term structure. Fourth, cross-country co-movement in HTM premia between the U.K. and the U.S. rises sharply after 1980—particularly at short horizons—pointing to a substantial shift toward globally integrated pricing of short-run equity risks.

To interpret these facts, we develop an international asset-pricing model with country-specific disaster risk and a global long-run risk factor. The model reveals a new risk-sharing channel: greater financial integration improves insurance against short-horizon, country-specific shocks, but leaves exposure to global long-run risks essentially unchanged. As a result, financial integration lowers compensation for short-maturity risks and thereby generates the steepening in the term structure of equity risk premia observed in the data post-1980.

[20]Note that our model does not account for the post-1980 increase in the correlation of short-term risk premia. Importantly, Colacito and Croce (2011) document that short- and long-run correlations of endowments have shifted substantially since the 1980s. In contrast, in our framework, these correlations are held constant across regimes to isolate the effect of risk sharing on the equity term structure. Incorporating changes in endowment correlations could therefore help explain the changes in the correlations of equity risk premia as well. We leave this extension for future research.

The model also reproduces the level, slope, and cyclical variation of the equity term structure over our 150-year sample, while matching standard moments of consumption growth, dividend growth, and market returns. Hence, we provide a unified explanation of the evolution of the equity term structure and show how a new risk-sharing mechanism redistributes risk compensation across maturities.

Our results open several avenues for future work. The asymmetric nature of risk sharing across maturities suggests that other markets with distinct horizon exposures—such as sovereign debt, corporate credit, or volatility claims—may exhibit similar integration-driven shifts in pricing. Accounting for changes in endowment correlations and other institutional shifts may further clarify how financial integration reallocates risk across horizons. More broadly, our findings highlight the value of long-run data combined with international asset-pricing models for understanding how market structure and capital flows shape the pricing of risk across maturities.

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Tables

Table 1: Variance Decomposition of Equity Yields

Panel A. 1870-2019

	Variances				Contributions		
Maturity	$e^{(n)}$	$\theta^{(n)}$	$g^{(n)}$	-2Cov($\theta^{(n)}, g^{(n)}$)	$\theta^{(n)}$	$g^{(n)}$	-2Cov($\theta^{(n)}, g^{(n)}$)
1	0.26%	0.03%	0.16%	0.08%	12%	60%	29%
10	0.05%	0.01%	0.01%	0.02%	30%	22%	48%
30	0.01%	0.00%	0.00%	0.00%	35%	19%	47%

Panel B. 1870-1979

	Variances				Contributions		
Maturity	$e^{(n)}$	$\theta^{(n)}$	$g^{(n)}$	-2Cov($\theta^{(n)}, g^{(n)}$)	$\theta^{(n)}$	$g^{(n)}$	-2Cov($\theta^{(n)}, g^{(n)}$)
1	0.18%	0.03%	0.09%	0.06%	18%	48%	34%
10	0.03%	0.01%	0.01%	0.02%	34%	19%	47%
30	0.00%	0.00%	0.00%	0.00%	39%	16%	45%

Panel C. 1980-2019

	Variances				Contributions		
Maturity	$e^{(n)}$	$\theta^{(n)}$	$g^{(n)}$	-2Cov($\theta^{(n)}, g^{(n)}$)	$\theta^{(n)}$	$g^{(n)}$	-2Cov($\theta^{(n)}, g^{(n)}$)
1	0.35%	0.01%	0.28%	0.05%	4%	80%	16%
10	0.05%	0.01%	0.01%	0.02%	22%	29%	49%
30	0.01%	0.00%	0.00%	0.00%	26%	25%	49%

This table reports the variance decomposition of equity yields $e^{(n)} = \theta^{(n)} - g^{(n)}$ into: (i) variation in HTM risk premia $\theta^{(n)}$, (ii) variation in expected dividend growth $g^{(n)}$, and (iii) the covariance term $-2\text{Cov}(\theta^{(n)}, g^{(n)})$. For each maturity n, the table shows the unconditional variance of each component together with its percentage contribution to $(e^{(n)})$. Panel A reports results for the full sample (1870–2019), while Panels B and C report the decomposition separately for the pre-1980 (1870–1979) and post-1980 (1980–2019) periods.

Table 2: Calibration

Parameter	Value	Interpretation/impact
γ	10	risk aversion
ψ	1.5	intertemporal elasticity of substitution
δ	0.98	time discount factor
α	0.97	home bias
$\mu_x = \mu_y$	0.03	average growth rate good X/Y without disasters
$\sigma_x = \sigma_y$	0	volatility of i.i.d. normal shocks
ρ_{xy}	0	correlation of i.i.d. normal shocks
τ	0.05	cointegration of X and Y
d	-0.23	disaster size
μ_z	0.04	average disaster intensity
ρ_z	0.94	persistency of disaster intensity
σ_z	0.022	volatility of i.i.d. normal shock disaster intensity
μ_d	0.02	average growth rate of dividends without disasters
k	2	leverage parameter

Model parameters are calibrated to match annualized moments of cash flows and returns. The rightmost column describes the economic quantity each parameter primarily affects.

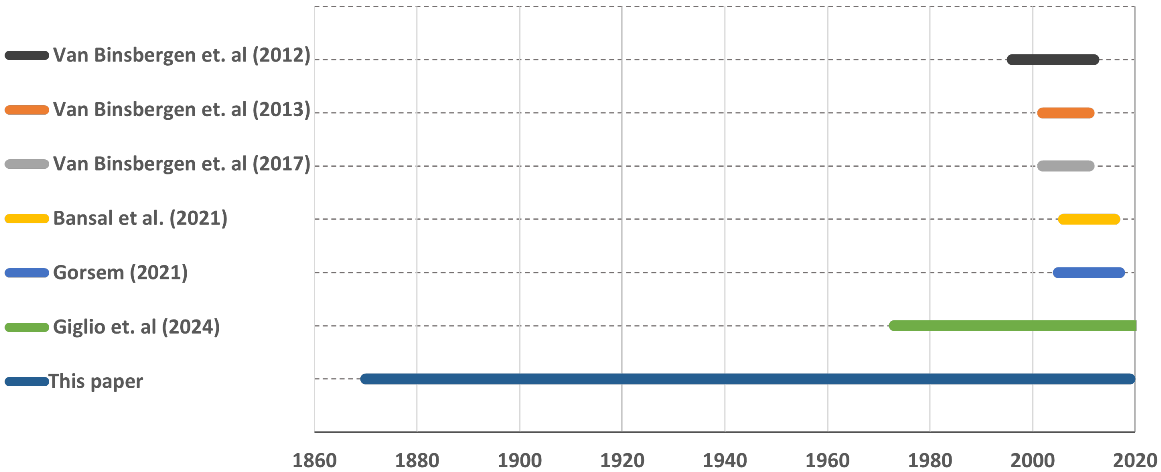
Table 3: Annualized Moments

	Data		Model	
	1870-1979	1980-2019	Autarky	Fin. integration
$E\left(\Delta c_t^h\right)$	0.0107	0.0216	0.0217	0.0224
			[0.0098; 0.0297]	[0.0101; 0.0309]
$\sigma\left(\Delta c_t^h\right)$	0.0324	0.0223	0.0362	0.0405
			[0.0024; 0.0709]	[0.0039; 0.0780]
$E\left(\Delta d_t^h\right)$	-0.0005	0.0034	0.0034	0.0049
			[-0.0205; 0.0193]	[-0.0199; 0.0217]
$\sigma\left(\Delta d_t^h\right)$	0.1427	0.1530	0.0724	0.0809
			[0.0048; 0.1428]	[0.0078; 0.1561]
$E\left(pd_t^h\right)$	3.1354	3.3310	2.9021	2.9746
			[2.7611; 3.0130]	[2.8211; 3.0973]
$\sigma\left(pd_t^h\right)$	0.1737	0.2285	0.0755	0.0884
			[0.0444; 0.1317]	[0.0487; 0.1560]
$E\left(R_{m,t}^h - R_{f,t}^h\right)$	0.0480	0.0595	0.0461	0.0391
			[0.0253; 0.0661]	[0.0186; 0.0572]
$E\left(R_{f,t}^h\right)$	0.0044	0.0291	0.0151	0.0201
			[0.0031; 0.0244]	[0.0083; 0.0294]
$\sigma\left(R_{m,t}^h\right)$	0.1273	0.1481	0.0689	0.0679
			[0.0335; 0.1083]	[0.0332; 0.1053]
$\sigma\left(R_{f,t}^h\right)$	0.0096	0.0130	0.0062	0.0066
			[0.0036; 0.0108]	[0.0036; 0.0118]

Empirical moments are computed for the UK using sample averages and standard deviations over two time periods. Correlations are reported with the US, and the pre-1980 sample begins in 1929 due to data availability for the US. In the model, we simulate 10,000 samples, each with the same length as the corresponding empirical dataset, and report the median of the simulated moments. The 5% and 95% quantiles across simulations are shown in square brackets. The autarky specification assumes no international risk sharing, while the financial integration setup assumes full financial market integration, with the home country accounting for 3% of initial world consumption.

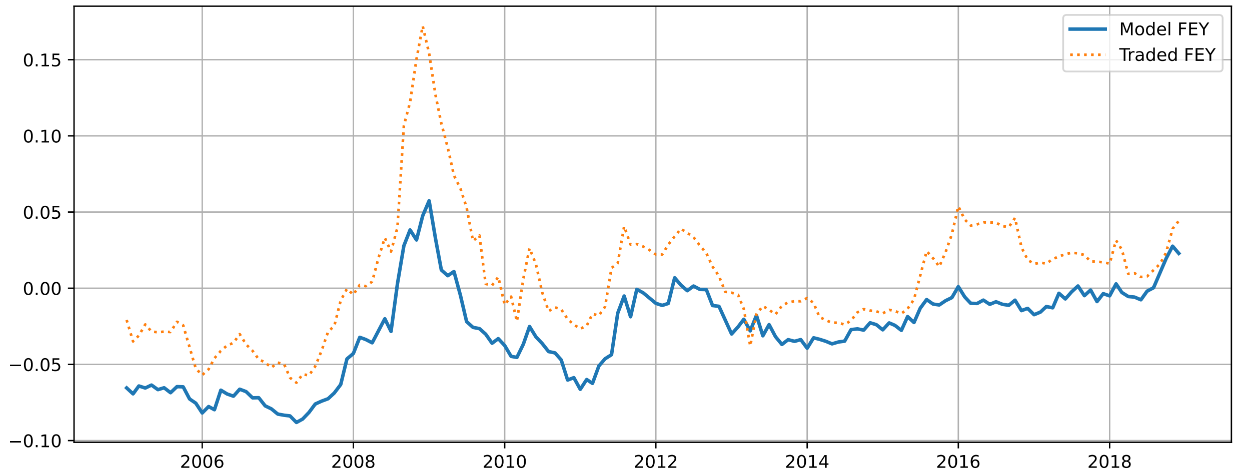
Figures

Figure 1: Time Span of Equity Yield Data in the Literature.



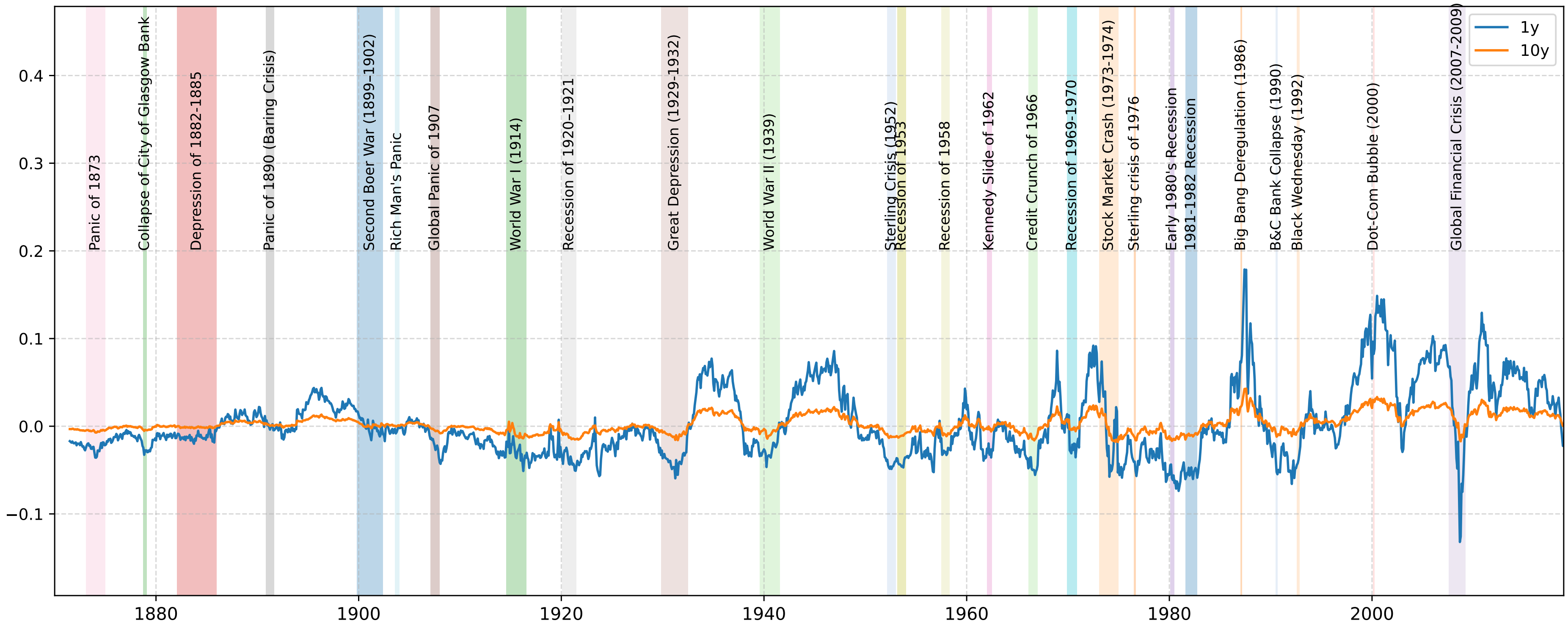
This figure shows the time spans covered by previous studies, all of which correspond to the U.S. market. The final (dark blue) bar represents the time span of our dataset, which covers the U.K. market.

Figure 2: Forward Equity Yields (5y): Model vs. Traded Dividend Futures.



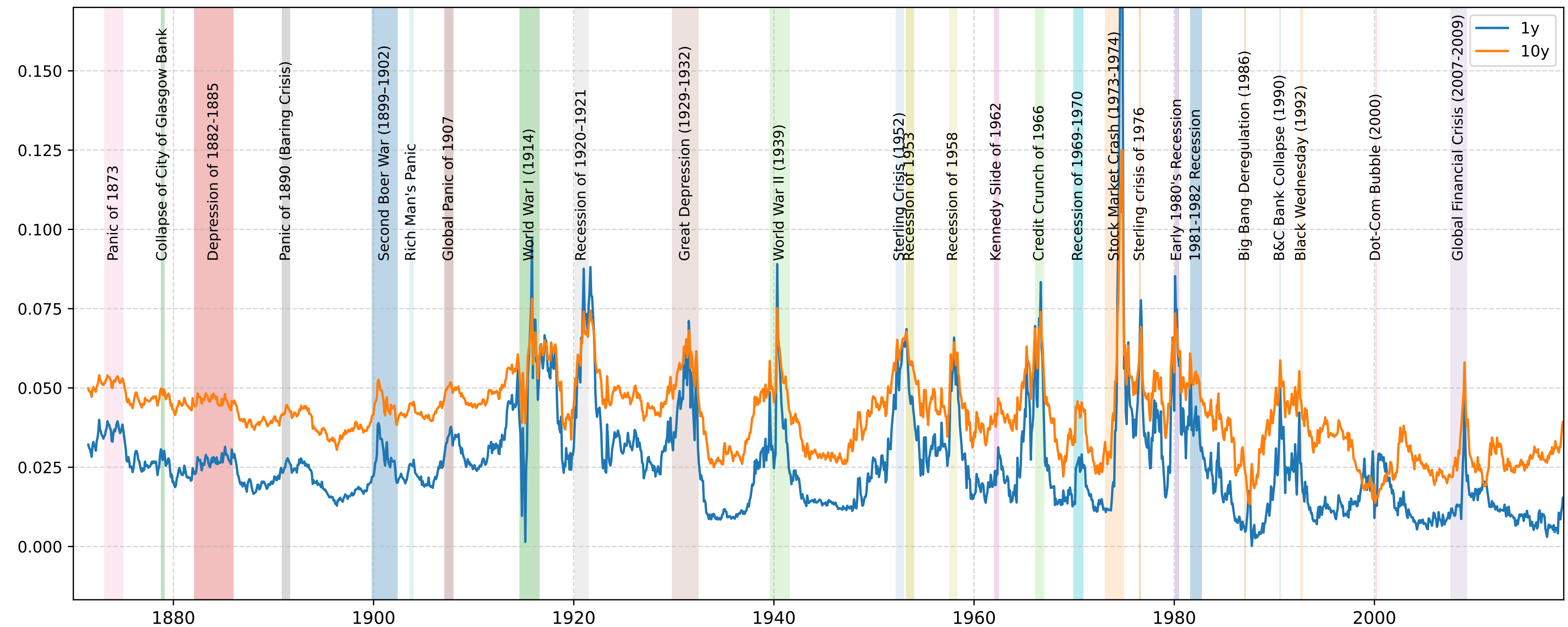
This figure compares the 5-year forward equity yield implied by our model (solid blue) to the yield extracted from traded FTSE dividend futures (dotted orange).

Figure 3: 10-Year and 1-Year HTM Expected Dividend Growth g.



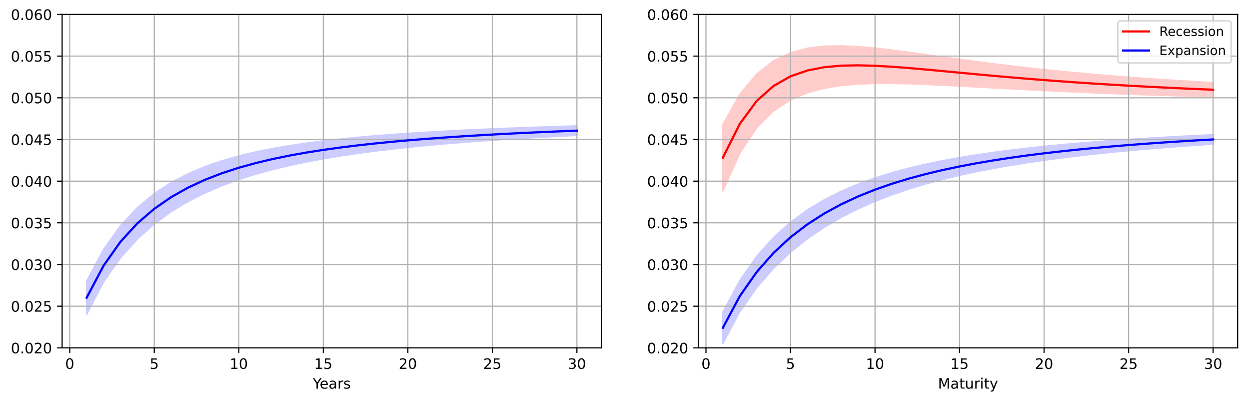
This figure displays the model-implied HTM expected dividend growth at the 1-year and 10-year maturities ($g_t^{(1)}$ and $g_t^{(10)}$, respectively). Shaded regions denote major historical episodes, including financial crises, wars, and recessions.

Figure 4: 10-Year and 1-Year HTM Risk Premia θ.



This figure displays the model-implied HTM risk premia at the 1-year and 10-year maturities ($\theta_t^{(1)}$ and $\theta_t^{(10)}$, respectively). Shaded regions denote major historical episodes, including financial crises, wars, and recessions.

Figure 5: Term Structure of HTM Risk Premia.

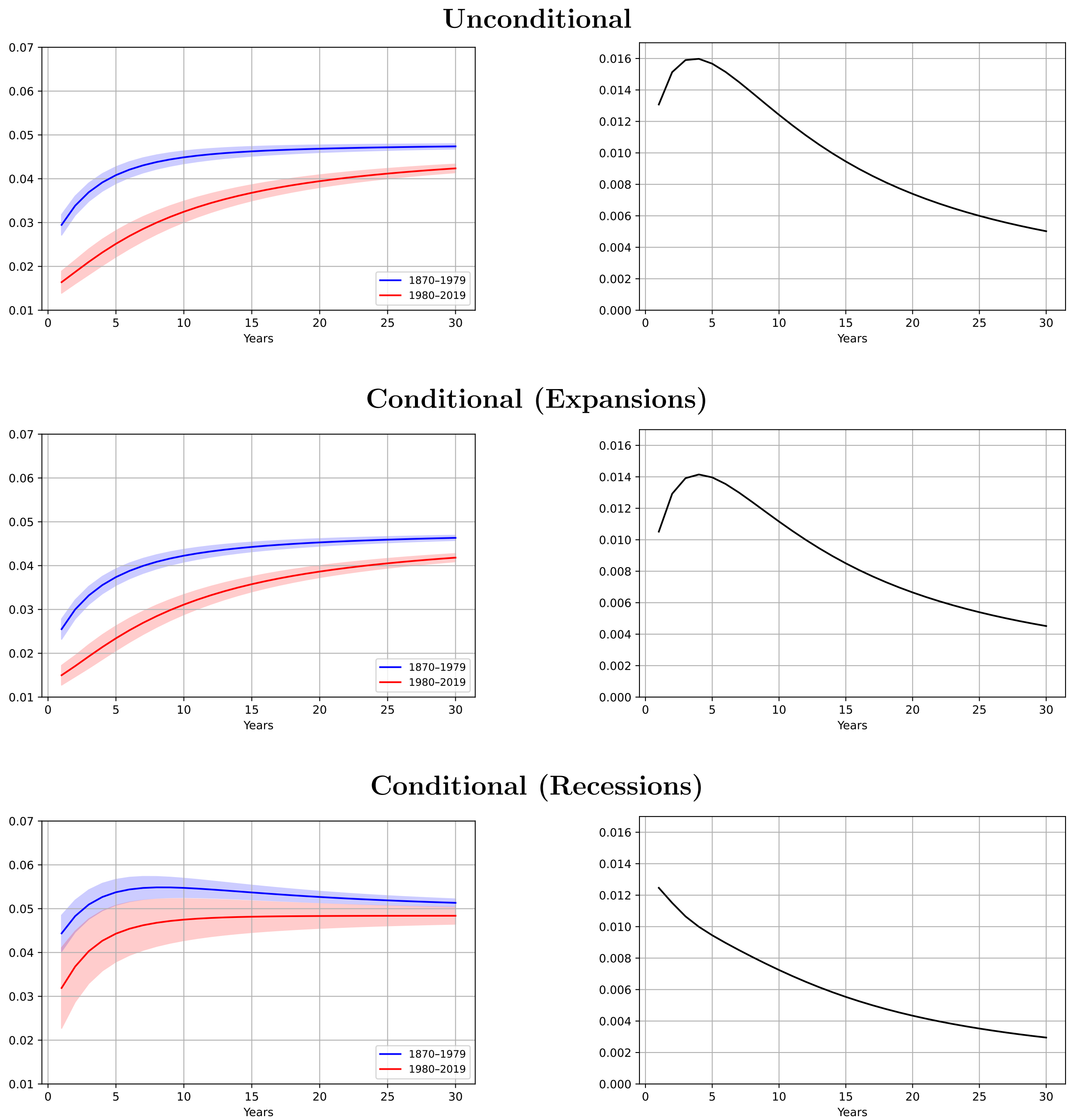


(a) Unconditional

(b) Conditional

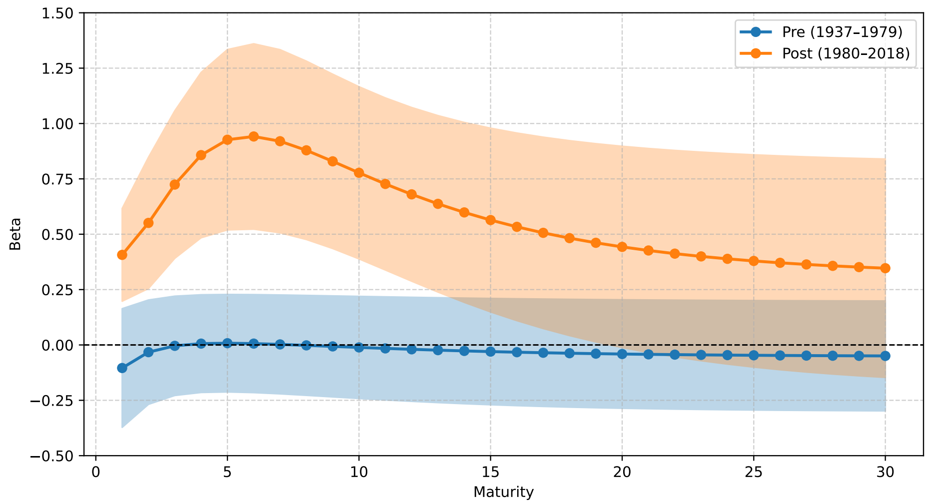
The figure plots the term structure of HTM risk premia. Panel (a) shows the unconditional premia over the full sample (1870–2019). Panel (b) shows the conditional premia during expansions and recessions, based on the recession index described in Section 2. Bands denote 95% confidence intervals based on HAC-robust standard errors.

Figure 6: Term Structure of HTM Risk Premia Before and After 1980.



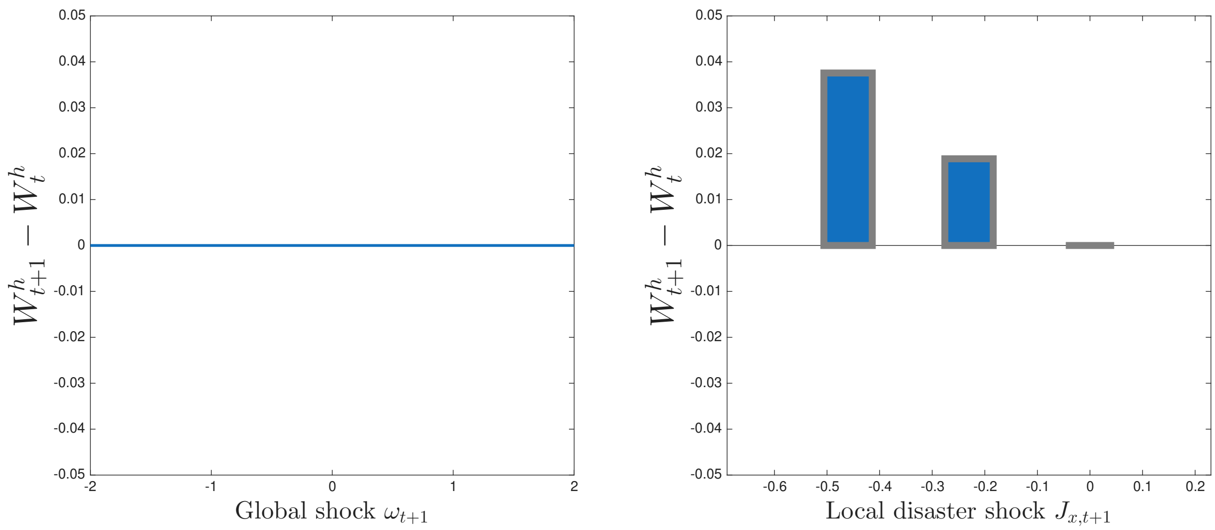
This figure plots the term structure of HTM risk premia for the 1870–1979 and 1980–2019 periods. The left column displays the estimated term structures, while the right column shows the difference (1870–1979 minus 1980–2019). The top row presents unconditional estimates for the full sample. The second and third rows condition on expansions and recessions, respectively, using the recession index described in Section 2. Shaded areas denote 95% confidence bands based on HAC-robust standard errors.

Figure 7: Cross-country betas of HTM equity risk premia by maturity.



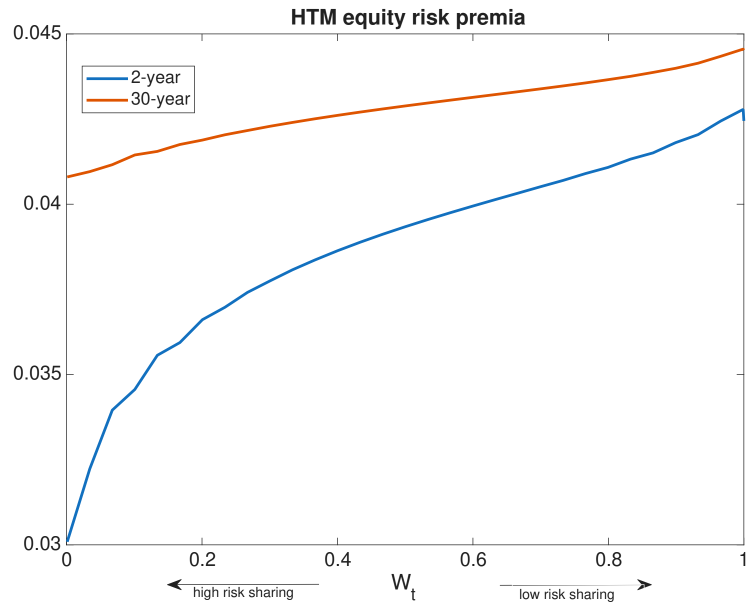
The figure reports slopes from standardized regressions (10) of U.K. HTM risk premia on U.S. HTM risk premia, for maturities from 1 to 30 years. The sample is split at 1980, with the pre-1980 period (1937–1979) shown in blue and the post-1980 period (1980–2018) in orange. Shaded areas denote 95% confidence intervals based on Newey–West standard errors with 12 lags.

Figure 8: Financial risk sharing and the impact of local and global shocks.



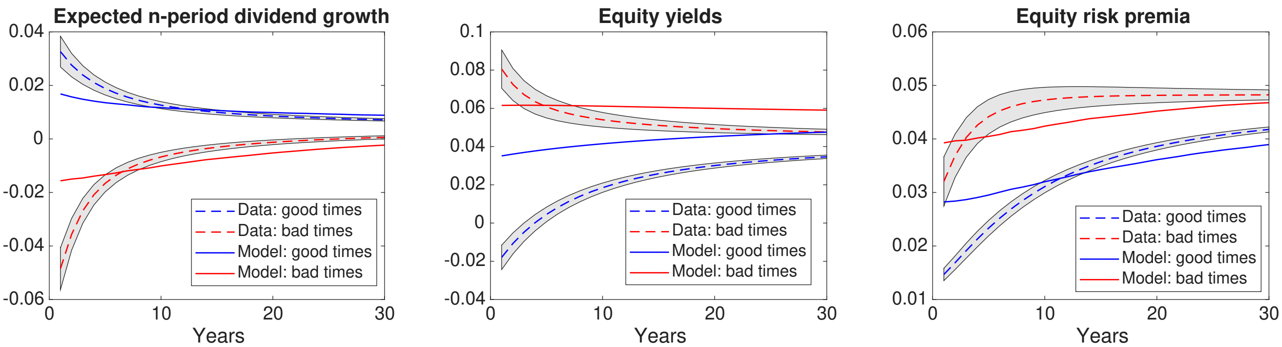
The figure plots the change in the consumption share $W_{t+1}^h - W_t^h$ as a function of the global shock $\omega_{t+1} \in (-1, 1)$ on the left and the local disaster shock $J_{x,t+1}$ with 0, 1 and 2 disaster in $t+1$ on the right. Results are shown for $W_t^h = 0.5$, $\lambda_t = \bar{\lambda}$ and $CI_t = 0$. Model parameters are provided in Table 3.

Figure 9: Financial risk sharing and pricing of long- and short-term risk premia.



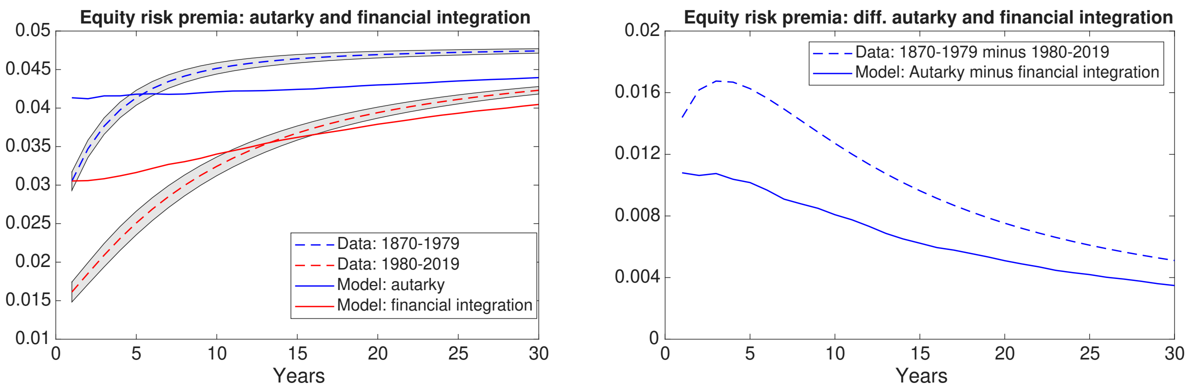
The figure plots 2-year and 30-year HTM equity risk premia as a function of the consumption share of the home country W_t. Results are shown for $\lambda_t = \bar{\lambda}$ and $CI_t = 0$. Model parameters are provided in Table 3.

Figure 10: Time variation of the equity term structure.



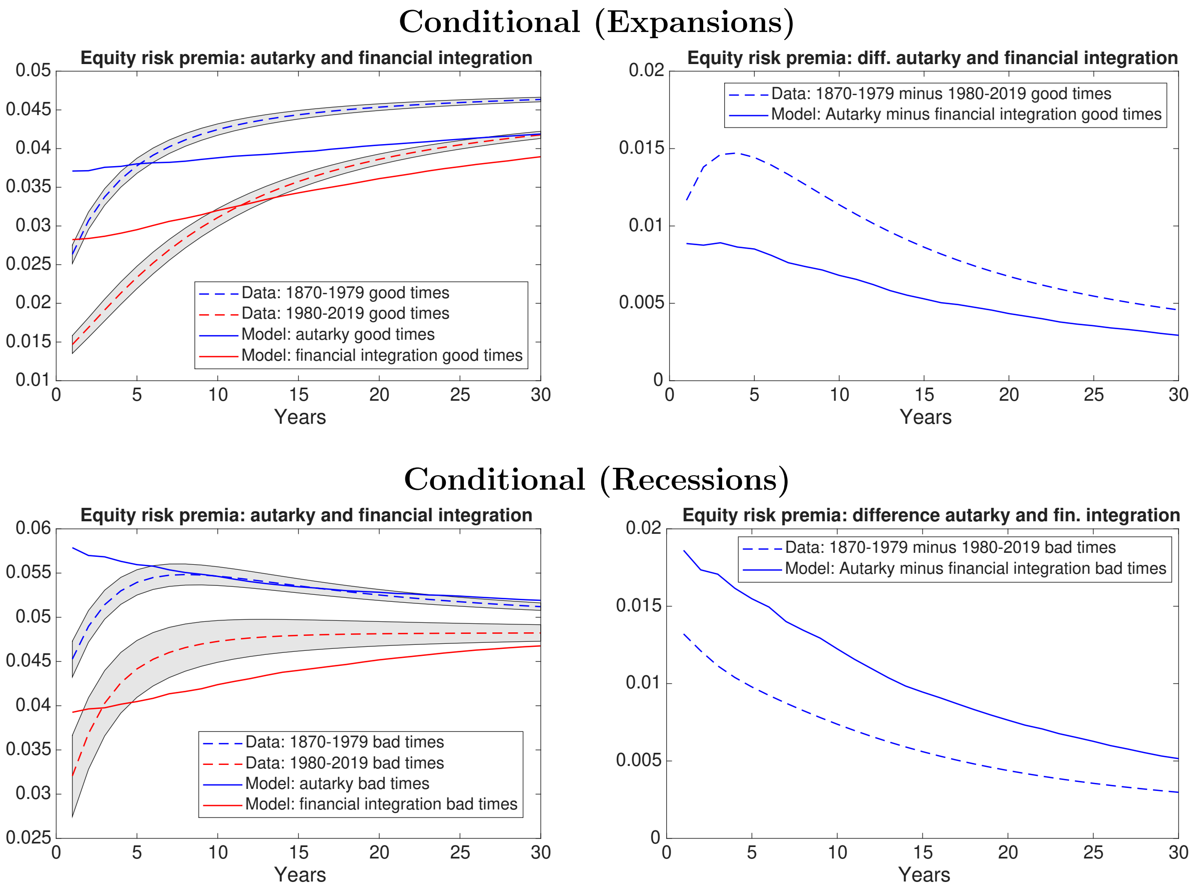
The figure displays n-period expected dividend growth (left), equity yields (center), and equity risk premia (right) in good and bad times. Solid lines represent model-implied outcomes, while dashed lines show the corresponding empirical quantities. Good and bad times in the data are defined as in Section 2. In the model, good (bad) times correspond to periods when the price-dividend ratio is above (below) its 25th percentile. Model results are obtained by simulating 10,000 samples of the same length as the empirical data and reporting the median across simulations. The calibrated parameters are reported in Table 3.

Figure 11: Financial market integration and the term structure of equity risk premia.



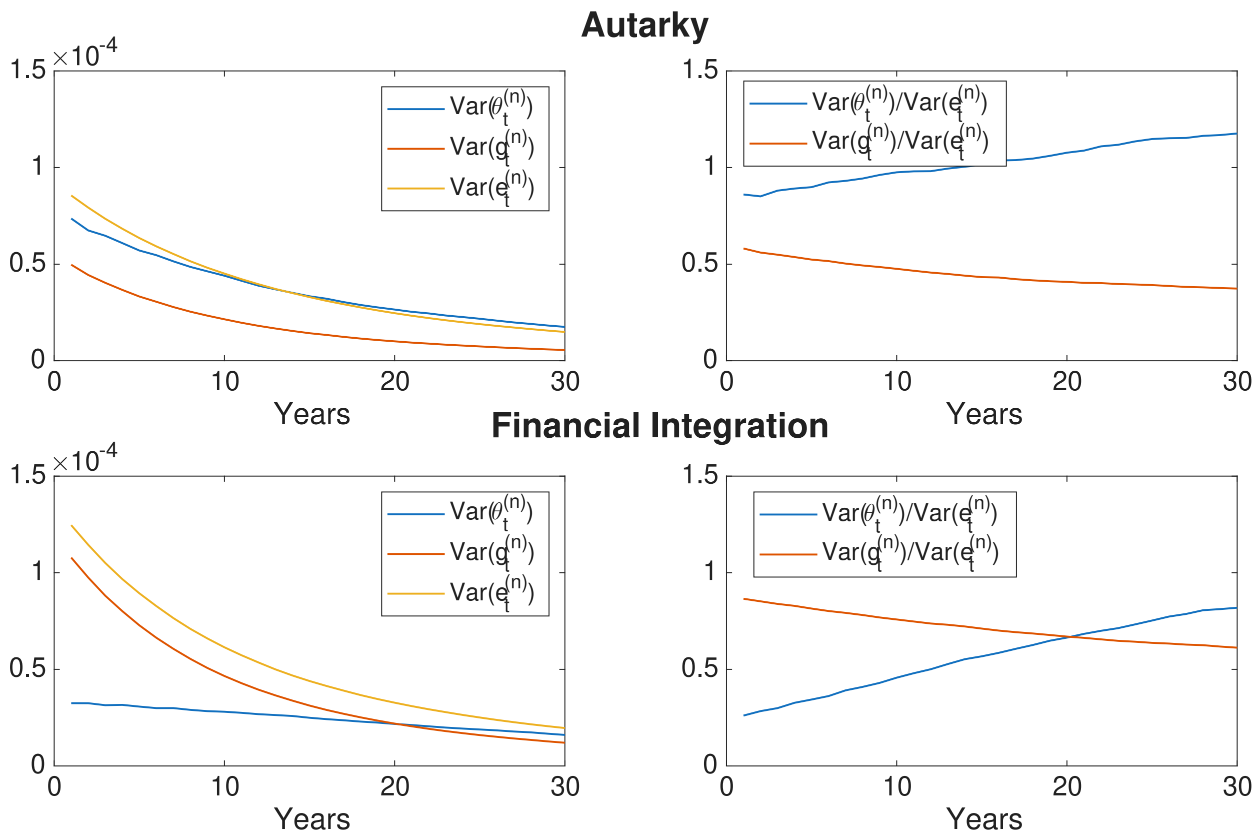
The figure plots the term structure of equity risk premia under different regimes. Solid lines show model outcomes and dashed line show the corresponding empirical outputs. For the data, term structures are plotted from 1870-1979 in blue and from 1980-2019 in red. For the model, we compare a regime of financial autarky in blue with a regime of full financial market integration in red. The right panel shows the difference between the financial autarky and the financial market integration case both for the data and model. Model results are computed simulating 10,000 samples with the length of the corresponding empirical sample and then taking medians across the samples. Model parameters are provided in Table 3.

Figure 12: Financial market integration and the term structure of equity risk premia: recessions versus expansions.



The figure plots the term structure of equity risk premia under different regimes. The first row shows results for expansions and the second row for recessions only. Solid lines show model outcomes and dashed line show the corresponding empirical outputs. For the data, term structures are plotted from 1870-1979 in blue and from 1980-2019 in red. For the model, we compare a regime of financial autarky in blue with a regime of full financial market integration in red. The right panel shows the difference between the financial autarky and the financial market integration case both for the data and model. Model results are computed simulating 10,000 samples with the length of the corresponding empirical sample and then taking medians across the samples. Model parameters are provided in Table 3.

Figure 13: Variance Decomposition of Equity Yields: Autarky versus Financial Integration.



The figure plots variances of equity yields $e_t^{(n)}$, HTM risk premia $\theta_t^{(n)}$ and dividend growth expectations $g_t^{(n)}$ in the model for the autarky and the financial integration regime. Left plots show absolute values for the variances across maturities and the right panel show the contribution of the variance of risk premia and dividend growth variation to the variance of equity yields. Model results are computed simulating 10,000 samples with the length of the corresponding empirical sample and then taking medians across the samples. Model parameters are provided in Table 3.

Appendices

A Data Construction

A.1 Equities

We construct a cross-section of UK stocks from 1870 to 2019 using Global Financial Data (GFD) and Datastream (DS). GFD provides UK stock data from 1825 to 1985, including daily prices (open, high, low, close), shares outstanding (reported quarterly), corporate actions (splits, recapitalizations, rights issues, stock distributions), and industry classifications. We use the closing price on the last trading day of each month, keep shares outstanding constant within each quarter, and construct total returns from unadjusted prices and cumulative adjustment factors. Market capitalization is the product of price and shares outstanding. Dividend yields are computed as the sum of regular dividends over the trailing 12 months (adjusted for corporate actions) divided by the end-of-month price, and are aggregated using GFD's industry taxonomy. DS provides UK stock data from 1964 to 2019, including end-of-month prices, total returns, shares outstanding, market capitalization, dividend yields, and industry classifications based on SIC codes. We use total return fields `RZ` or `RI` when available and compute returns manually otherwise; dividend yields are used as reported. For market capitalization, we use `WC08001` when available and `MV` otherwise, dropping observations with inconsistencies between reported market capitalization and the product of price and shares outstanding following Schmidt et al. (2015). We map SIC codes to the GFD industry taxonomy and apply the same filters as in Ince and Porter (2006) and Schmidt et al. (2015), dropping firms with fewer than 24 months of data, zero return variation, return jumps above 300%, or high first-order autocorrelation.

GFD and DS are not linked by a predefined mapping, so we match them manually over the overlapping period 1964–1985. In this window, GFD includes 3,119 unique stocks and DS 2,627, of which we ultimately match 1,450 using a two-step procedure. First, we estimate pairwise regressions of returns using up to 100 overlapping non-zero observations and retain candidate pairs that either (i) have similar names (Jaro–Winkler distance below 0.20) and high regression fit ($R^2 \geq 0.99$) or beta close to one ($0.99 < \beta < 1.01$), or (ii) have both moderate fit ($R^2 > 0.8$) and beta in the range 0.7 to 1.0; no match is accepted at this stage. Second, we verify each candidate manually using company records from UK Companies House and Grace's Guide, ignoring purely cosmetic differences in abbreviations or punctuation. We confirm a match only if both names appear in the filings of the same legal entity or we can clearly establish that one is the legal continuation of the other (e.g., renamings, takeovers, or reorganizations). Statistical similarity is never sufficient on its own to accept a match, although we do confirm some cases where names differ if corporate records unambiguously link the entities. Appendix A.1 provides additional implementation details.

A.2 Anomaly Characteristics

The construction of our anomaly characteristics closely follows the definitions in Giglio et al. (2024). For most variables rebalanced annually, values are updated in June or December. Jensen et al. (2023) lag accounting variables by four months. Our long sample does not allow us to construct accounting-based characteristics before 1985.[21] The closest we have to an accounting variable is dividend yield. We lag this measure by four months to reflect potential reporting delays, which is also consistent with the idea that historical investors may not have observed dividend payments immediately.

Size (`size`). Following Fama and French (1993), size is defined as market equity measured at the end of June of the previous year. Specifically, for any month t in year Y, we assign:

$$\text{size}_{i,t} = ME_{i,\text{Jun}(Y-1)}.$$

Dividend Yield (`divp`). Following Naranjo et al. (1998), dividend yield is defined using

[21]Previous work has documented significant quality issues in Datastream accounting variables prior to the mid-1980s (Ince and Porter, 2006; Schmidt et al., 2015).

the value observed in December of the prior calendar year. Specifically, for any month t in year Y, we assign:

$$\text{divp}_{i,t} = DY_{i,\text{Dec}(Y-1)}.$$

Dividend Yield (Monthly) (`divpm`). This is our own monthly analogue of the dividend yield variable in Naranjo et al. (1998). Just as `valuem` offers a monthly implementation of `value` in Giglio et al. (2024), we define:

$$\text{divpm}_{i,t} = DY_{i,t},$$

where dividend yield at month t is assumed to be known with a six-month reporting delay.

Dividend Yield-Momentum (`divpmom`). We construct a long-sample version of `value-momentum` in Giglio et al. (2024) by replacing book-to-market with dividend yield. The signal is available from 1870 onward and defined as the sum of the cross-sectional ranks of dividend yield and six-month momentum:

$$\text{divpmom}_{i,t} = \text{rank}(\text{divp}_{i,t}) + \text{rank}(\text{mom}_{i,t}).$$

Industry Momentum (`indmom`). Following Moskowitz and Grinblatt (1999), we construct an industry momentum variable by first computing the value-weighted return of each industry in every month, and then compounding those returns from $t-1$ to $t-6$:

$$\text{indmom}_{j,t} = \prod_{k=1}^{6} \left(1 + r^{\text{ind}}_{j,t-k}\right) - 1,$$

where $r^{\text{ind}}_{j,t-k}$ is the value-weighted return of industry j in month $t-k$. Industries are then ranked cross-sectionally based on their indmom, and each firm inherits the industry's rank:

$$\text{indmom}_{i,t} = \text{rank}(\text{indmom}_{j,t}) \quad \text{for firm } i \in j.$$

6-Month Momentum (`mom`). Following Jegadeesh and Titman (1993), we compute six-month momentum by compounding returns from month $t-2$ to $t-6$, skipping the most recent month:

$$\text{mom}_{i,t} = \prod_{k=2}^{6}(1 + r_{i,t-k}) - 1.$$

12-Month Momentum (`mom12`). Following Jegadeesh and Titman (1993), twelve-month momentum is defined analogously, compounding returns from $t-2$ to $t-12$:

$$\text{mom12}_{i,t} = \prod_{k=2}^{12}(1 + r_{i,t-k}) - 1.$$

Beta Arbitrage (`betaarb`). In the spirit of Cooper et al. (2008), we estimate rolling betas over a 60-month window using monthly excess returns. For each firm i, the beta at month t is computed as:

$$\beta_{i,t} = \frac{\text{Cov}(r_{i,t-k}, r_{t-k}^{\text{mkt}})}{\text{Var}(r_{t-k}^{\text{mkt}})}, \quad k = 1, \ldots, 60,$$

where $r_{i,t}$ denotes the firm's excess return and r_t^{mkt} is the market excess return computed as the value-weighted return of all firms.

Composite Issuance (`ciss`). Following Daniel and Titman (2006), composite issuance is computed as:

$$\text{ciss}_{i,t} = \log(ME_{i,t-13}) - \log(ME_{i,t-60}) - \log\left(\prod_{l=13}^{60}(1 + r_{i,t-l})\right),$$

where returns are compounded over 48 months using monthly total returns.

Our long sample allows us to construct six additional characteristics not included in the main analysis: Medium-Term Reversal (`momrev`), Short-Term Reversal (`strev`), Long-Term Reversal (`lrrev`), Price (`price`), Seasonality (`season`), and Idiosyncratic Volatility using Dividend Yield (`ivoldivp`).

We ultimately exclude these variables because their inclusion makes the estimation procedure unstable. Most of these anomalies deliver little to no CAPM alpha in the latter part of the sample, consistent with the broader finding that many anomalies have weakened or disappeared over time. However, they may have been relevant in earlier decades, which can lead them to receive large weights in the principal components. These weights, in turn, introduce noise when the same variables fail to explain returns later in the sample. The reverse also applies: characteristics that are only informative in the latter part of the sample might receive little weight if they had weak pricing power in earlier periods. This time inconsistency compromises the stability of the estimated SDF and motivates our focus on a more stable subset of characteristics.

For completeness, we define these six excluded characteristics below:

Medium-Term Reversal (`momrev`). Following Jegadeesh and Titman (1993), medium-term reversal is computed using returns from months $t-14$ to $t-19$:

$$\text{momrev}_{i,t} = \prod_{k=14}^{19} (1 + r_{i,t-k}) - 1.$$

Short-Term Reversal (`strev`). Following Jegadeesh (1990), short-term reversal is defined as the return in month $t-1$:

$$\text{strev}_{i,t} = r_{i,t-1}.$$

Long-Term Reversal (`lrrev`). Following De Bondt and Thaler (1985), long-term reversal is computed as the compounded return from month $t-13$ to $t-60$:

$$\text{lrrev}_{i,t} = \prod_{k=13}^{60} (1 + r_{i,t-k}) - 1.$$

Price (`price`). Following Blume and Husic (1973), price is computed as the log of the ratio of market equity to shares outstanding, lagged one month:

$$\text{price}_{i,t} = \log\left(\frac{ME_{i,t-1}}{SHROUT_{i,t-1}}\right).$$

Seasonality (`season`). Following Heston and Sadka (2008), seasonality is defined as the average return the firm earned in the same calendar month over the previous five years. For a firm observed in month m of year Y, we assign:

$$\text{season}_{i,t} = \frac{1}{5}\sum_{k=1}^{5} r_{i,m(Y-k)},$$

where $r_{i,m(Y-k)}$ is the return in the same calendar month m, k years earlier.

Idiosyncratic Volatility using Dividend Yield (`ivoldivp`). We construct a long-sample version of idiosyncratic volatility, `ivol` in Giglio et al. (2024), by replacing the value characteristic with dividend yield. Rolling regressions are estimated over 60 months (requiring at least 30 observations), using:

$$r_{i,\tau} - r_{f,\tau} = \alpha_i + \beta_i^{\text{mkt}}(r_{m,\tau} - r_{f,\tau}) + \beta_i^{\text{size}} \cdot \text{size}_{i,\tau} + \beta_i^{\text{divp}} \cdot \text{divp}_{i,\tau} + \varepsilon_{i,\tau}, \quad \tau = t-59, \dots, t,$$

and defining:

$$\text{ivoldivp}_{i,t} = \text{SD}(\varepsilon_{i,t-59:t}).$$

A.3 UK Recession Index

We construct a monthly recession indicator for the UK by combining GDP-based signals with a market-based proxy.[22] For the period starting in July 1955, we identify recessions using official quarterly real GDP data from the Office for National Statistics (ONS).[23] Following standard practice (e.g. Thomas et al., 2010), we define a recession as any period with

[22] Unlike the US, the UK does not have an official recession chronology maintained by an institution like the NBER.

[23] Specifically, we use the series Gross Domestic Product: chained volume measures, seasonally adjusted (code `ABMI`), reported in millions of pounds sterling.

two consecutive quarters of negative real GDP growth, and classify all months within such quarters as recession months.

Before 1955, quarterly GDP data are unavailable and only annual real GDP estimates exist, which are too coarse to identify cyclical turning points at monthly or quarterly frequency. To overcome this limitation, we construct a market-based recession proxy using the price-dividend (P/D) ratio of the equity market. The P/D ratio is computed as the inverse of the market-level dividend yield. Dividend yields are calculated at the stock level as the sum of dividends paid over the trailing 12 months divided by the stock price, and are then aggregated into a market-level yield using value weights based on market capitalization. A month is classified as a recession if the P/D ratio is below the 25th percentile of its full-sample distribution; otherwise, it is classified as an expansion. This classification reflects the empirical observation that the P/D ratio tends to decrease during economic downturns.

A.4 Risk-Free Rate

We construct a consistent measure of the monthly return on the safest available short-term asset in the UK, which we refer to as the risk-free rate. This rate reflects the most liquid and creditworthy short-term instrument available to investors at a given point in time. From 2008 onward, we use the 1-month UK Treasury bill yield; from 1900 to 2007, we use the 3-month UK Treasury bill yield. Both series are from GFD and represent standard choices in the literature for government-backed short-term borrowing costs (Meyer et al., 2022). For the period before 1900, we follow Campbell et al. (2021) and use market-based discount rates on high-quality commercial paper (prime bills), which better reflect actual short-term funding conditions. We extend the series back to 1870 using data compiled from Parliamentary Papers (1857), Nishimura (1971), and Capie and Webber (1985), all of which are publicly available through the Bank of England's "A Millennium of Macroeconomic Data."[24]

[24] Available at https://www.bankofengland.co.uk/statistics/research-datasets

A.5 Term Structure of Interest Rates

We use UK government bond yields to construct forward equity yields. Yield data comes from GFD and DS, which complement each other.[25] We collect all available monthly yield series with maturities of one year or more, some of which begin as early as 1870. When DS yields are missing, we use GFD data instead.[26] We interpolate the available yields to construct a complete term structure of interest rates with annual steps from 1 to 30 years. Specifically, we fit a Nelson-Siegel-Svensson model month by month, using yields at the following maturities: 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 15, 20, and 30 years. We estimate the term structure of yields only when they are available for at least four of these maturities.

B Empirical Model

B.1 Model Structure

State vector and its dynamics. We begin with a linear factor model. The state vector is

$$\underset{(k\times 1)}{F_t} = \begin{pmatrix} \underset{(p\times 1)}{F_{r,t}} \\ \underset{(p\times 1)}{F_{y,t}} \end{pmatrix}, \quad k = 2p, \tag{20}$$

where $F_{r,t}$ are excess returns and $F_{y,t}$ log dividend yields of the same p portfolios.

[25]For example, 1-year yields start in 1970 in DS but only in 1979 in GFD.

[26]Tickers are available upon request.

The state evolves according to a VAR with homoskedastic Gaussian errors $\underset{(k\times 1)}{u_{t+1}} \sim \mathcal{N}(0, \underset{(k\times k)}{\Sigma})$:

$$\underset{(k\times 1)}{F_{t+1}} = \underset{(k\times 1)}{c} + \underset{(k\times k)}{\rho}\underset{(k\times 1)}{F_t} + \underset{(k\times 1)}{u_{t+1}} \tag{21}$$

$$\begin{pmatrix} \underset{(p\times 1)}{F_{r,t+1}} \\ \underset{(p\times 1)}{F_{y,t+1}} \end{pmatrix} = \begin{pmatrix} \underset{(p\times 1)}{c_r} \\ \underset{(p\times 1)}{c_y} \end{pmatrix} + \begin{pmatrix} \underset{(p\times p)}{\mathbf{0}} & \underset{(p\times p)}{\rho_{r,y}} \\ \underset{(p\times p)}{\mathbf{0}} & \underset{(p\times p)}{\rho_{y,y}} \end{pmatrix} \begin{pmatrix} \underset{(p\times 1)}{F_{r,t}} \\ \underset{(p\times 1)}{F_{y,t}} \end{pmatrix} + \begin{pmatrix} \underset{(p\times 1)}{u_{r,t+1}} \\ \underset{(p\times 1)}{u_{y,t+1}} \end{pmatrix}. \tag{22}$$

The log stochastic discount factor (SDF) is

$$\underset{(1\times 1)}{m_{t+1}} = -\underset{(1\times 1)}{r_{f,t}} - \frac{1}{2}\underset{(1\times k)}{\lambda_t'}\underset{(k\times k)}{\Sigma}\underset{(k\times 1)}{\lambda_t} - \underset{(1\times k)}{\lambda_t'}\underset{(k\times 1)}{u_{t+1}} \tag{23}$$

with affine prices of risk

$$\underset{(k\times 1)}{\lambda_t} = \underset{(k\times 1)}{\lambda} + \underset{(k\times k)}{\Lambda}\underset{(k\times 1)}{F_t} \tag{24}$$

$$\begin{pmatrix} \underset{(p\times 1)}{\lambda_{r,t}} \\ \underset{(p\times 1)}{\lambda_{y,t}} \end{pmatrix} = \begin{pmatrix} \underset{(p\times 1)}{\lambda_r} \\ \underset{(p\times 1)}{\mathbf{0}} \end{pmatrix} + \begin{pmatrix} \underset{(p\times p)}{\mathbf{0}} & \underset{(p\times p)}{\Lambda_{ry}} \\ \underset{(p\times p)}{\mathbf{0}} & \underset{(p\times p)}{\mathbf{0}} \end{pmatrix} \begin{pmatrix} \underset{(p\times 1)}{F_{r,t}} \\ \underset{(p\times 1)}{F_{y,t}} \end{pmatrix}. \tag{25}$$

Where (25) reflects the restrictions motivated in the main text.

Dynamics of test assets. The dynamics of log prices, dividend yields, and total returns for the n test assets in the economy are:

$$\underset{(n\times 1)}{\Delta p_{t+1}} - r_{f,t}\mathbf{1}_n = \underset{(n\times 1)}{\gamma_0} + \underset{(n\times k)}{\gamma_1}\underset{(k\times 1)}{F_t} + \underset{(n\times k)}{\gamma_2}\underset{(k\times 1)}{u_{t+1}} + \underset{(n\times 1)}{\varepsilon^p_{t+1}}, \tag{26}$$

$$\underset{(n\times 1)}{y_t} = \underset{(n\times 1)}{b_0} + \underset{(n\times k)}{b_1}\underset{(k\times 1)}{F_t} + \underset{(n\times 1)}{\varepsilon^y_t}, \tag{27}$$

$$\underset{(n\times 1)}{r_{t+1}} - r_{f,t}\mathbf{1}_n = \underset{(n\times 1)}{\beta_0} + \underset{(n\times k)}{\beta_1}\underset{(k\times 1)}{F_t} + \underset{(n\times k)}{\beta_2}\underset{(k\times 1)}{u_{t+1}} + \underset{(n\times 1)}{\varepsilon^r_{t+1}}. \tag{28}$$

where n denotes the number of portfolios in the cross-section, $\mathbf{1}_n$ denotes the $n \times 1$ vector of ones, and $\varepsilon^p_{t+1}, \varepsilon^y_t, \varepsilon^r_{t+1} \in \mathbb{R}^{n\times 1}$ are asset-specific shocks.

The intercept, slopes, and shocks of the total return can be expressed as a function of the other parameters and their shocks:

$$\beta_0 = \underset{(n\times 1)}{\gamma_0} + \underset{(n\times 1)}{b_0} + \underset{(n\times k)}{b_1}\underset{(k\times 1)}{c} \tag{29}$$

$$\beta_1 = \underset{(n\times k)}{\gamma_1} + \underset{(n\times k)}{b_1}\underset{(k\times k)}{\rho} \tag{30}$$

$$\beta_2 = \underset{(n\times k)}{\gamma_2} + \underset{(n\times k)}{b_1} \tag{31}$$

$$\varepsilon_{t+1} = \underset{(n\times 1)}{\epsilon_{t+1}} + \underset{(n\times 1)}{\nu_{t+1}} \tag{32}$$

Euler equation. The Euler Equation implies restrictions on the parameter in the system of the form:

$$\mathcal{A}_{n,t} + \frac{1}{2}\,\mathrm{diag}(\mathcal{B}_{n,t}) = \mathbf{0}$$

Which, we derive in detail.

The Euler Equation for a particular asset i is:[27]

$$\mathbb{E}_t\left[e^{m_{t+1}+\Delta p^i_{t+1}+y^i_{t+1}}\right] = 1. \tag{33}$$

which since $m_{t+1}, \Delta p^i_{t+1}, y^i_{t+1}$ are normally distributed, is equivalent to:

$$\mathbb{E}_t\left[m_{t+1} + \Delta p^i_{t+1} + y^i_{t+1}\right] + \frac{1}{2}\,\mathbb{V}_t\left[m_{t+1} + \Delta p^i_{t+1} + y^i_{t+1}\right] = 0. \tag{34}$$

[27] $\Delta p_{t+1}, y_{t+1}, r_{t+1}$ contain n assets:

$$\Delta p_{t+1} \equiv (\Delta p^1_{t+1}, \ldots, \Delta p^n_{t+1})', \quad y_{t+1} \equiv (y^1_{t+1}, \ldots, y^n_{t+1})', \quad r_{t+1} \equiv (r^1_{t+1}, \ldots, r^n_{t+1})'.$$

Stacking across $i = 1, \ldots, n$ yields vector restrictions. Define:

$$\mathcal{A}_{n,t} \equiv \mathbb{E}_t \left[\underset{(n\times1)}{m_{t+1}\mathbf{1}_n + \Delta p_{t+1} + y_{t+1}} \right]. \tag{35}$$

Substitute the SDF (23), the log–price equation (26), the next–period yield equation $y_{t+1} = b_0 + b_1 F_{t+1} + \varepsilon^y_{t+1}$, the VAR process (21), and use $\mathbb{E}_t[u_{t+1}] = 0$, $\mathbb{E}_t[\varepsilon^p_{t+1}] = \mathbb{E}_t[\varepsilon^y_{t+1}] = 0$. Then:

$$\begin{aligned}
\mathcal{A}_{n,t} &= \mathbb{E}_t\Big[\big(-r_{f,t} - \tfrac{1}{2}\,\lambda_t'\Sigma\lambda_t - \lambda_t' u_{t+1}\big)\mathbf{1}_n + \big(r_{f,t}\mathbf{1}_n + \gamma_0 + \gamma_1 F_t + \gamma_2 u_{t+1} + \varepsilon^p_{t+1}\big) \\
&\qquad + \big(b_0 + b_1(c + \rho F_t + u_{t+1}) + \varepsilon^y_{t+1}\big)\Big] &(36)\\
&= \mathbb{E}_t\Big[-\tfrac{1}{2}\,(\lambda_t'\Sigma\lambda_t)\mathbf{1}_n + (\gamma_0 + \gamma_1 F_t) + \big(b_0 + b_1(c + \rho F_t)\big) \\
&\qquad + \underbrace{\big(-\mathbf{1}_n\lambda_t' + \gamma_2 + b_1\big)}_{(n\times k)} u_{t+1} + \varepsilon^p_{t+1} + \varepsilon^y_{t+1}\Big]. &(37)
\end{aligned}$$

Then, take conditional expectations at time t to arrive at the $n \times 1$ vector:

$$\mathcal{A}_{n,t} = -\tfrac{1}{2}\,\mathbf{1}_n(\lambda_t'\Sigma\lambda_t) + (\gamma_0 + \gamma_1 F_t) + \big(b_0 + b_1(c + \rho F_t)\big), \tag{38}$$

For the variance part, define:

$$\mathcal{B}_{n,t} \equiv \mathbb{V}_t \left[\underset{(n\times1)}{m_{t+1}\mathbf{1}_n + \Delta p_{t+1} + y_{t+1}} \right]. \tag{39}$$

Identify t-dated terms in:

$$\begin{aligned}
m_{t+1}\mathbf{1}_n &= -r_{f,t}\mathbf{1}_n - \tfrac{1}{2}\,\lambda_t'\Sigma\lambda_t - \lambda_t' u_{t+1}, \\
\Delta p_{t+1} &= r_{f,t}\mathbf{1}_n + \gamma_0 + \gamma_1 F_t + \gamma_2 u_{t+1} + \varepsilon^p_{t+1}, \\
y_{t+1} &= b_0 + b_1(c + \rho F_t) + b_1 u_{t+1} + \varepsilon^y_{t+1},
\end{aligned}$$

That is:

$$m_{t+1}\mathbf{1}_n = (\text{terms at } t) \;-\; \underbrace{\mathbf{1}_n\lambda_t'}_{(n\times k)} \underset{(k\times1)}{u_{t+1}}, \tag{40}$$

$$\Delta p_{t+1} = (\text{terms at } t) \;+\; \underset{(n\times k)}{\gamma_2} \; \underset{(k\times1)}{u_{t+1}} \;+\; \underset{(n\times1)}{\varepsilon^p_{t+1}}, \tag{41}$$

$$y_{t+1} = (\text{terms at } t) \;+\; \underset{(n\times k)}{b_1} \; \underset{(k\times1)}{u_{t+1}} \;+\; \underset{(n\times1)}{\varepsilon^y_{t+1}}. \tag{42}$$

Define the following terms that don't contain t-dated terms:[28]

$$\underset{(n\times1)}{\eta_{t+1}} \equiv \Big(\underbrace{-\mathbf{1}_n\lambda_t' + \gamma_2 + b_1}_{\underset{(n\times k)}{A_t}}\Big)\underset{(k\times1)}{u_{t+1}} \;+\; \underset{(n\times1)}{\varepsilon^p_{t+1}} \;+\; \underset{(n\times1)}{\varepsilon^y_{t+1}}, \tag{43}$$

$$A_t \equiv -\mathbf{1}_n\lambda_t' + \gamma_2 + b_1. \tag{44}$$

Then

$$\mathcal{B}_{n,t} = \mathbb{V}_t[\eta_{t+1}] = \mathbb{E}_t[\eta_{t+1}\eta_{t+1}'] \qquad (\text{zero conditional mean}) \tag{45}$$

$$= \mathbb{E}_t\Big[\big(A_t u_{t+1} + \varepsilon^p_{t+1} + \varepsilon^y_{t+1}\big)\big(A_t u_{t+1} + \varepsilon^p_{t+1} + \varepsilon^y_{t+1}\big)'\Big] \tag{46}$$

$$= \underbrace{\mathbb{E}_t[A_t u_{t+1}u_{t+1}'A_t']}_{(1)} + \underbrace{\mathbb{E}_t[A_t u_{t+1}\varepsilon^{p\prime}_{t+1}]}_{(2)} + \underbrace{\mathbb{E}_t[A_t u_{t+1}\varepsilon^{y\prime}_{t+1}]}_{(3)} + \underbrace{\mathbb{E}_t[\varepsilon^p_{t+1}u_{t+1}'A_t']}_{(4)}$$

$$+ \underbrace{\mathbb{E}_t[\varepsilon^p_{t+1}\varepsilon^{p\prime}_{t+1}]}_{(5)} + \underbrace{\mathbb{E}_t[\varepsilon^p_{t+1}\varepsilon^{y\prime}_{t+1}]}_{(6)} + \underbrace{\mathbb{E}_t[\varepsilon^y_{t+1}u_{t+1}'A_t']}_{(7)} + \underbrace{\mathbb{E}_t[\varepsilon^y_{t+1}\varepsilon^{p\prime}_{t+1}]}_{(8)} + \underbrace{\mathbb{E}_t[\varepsilon^y_{t+1}\varepsilon^{y\prime}_{t+1}]}_{(9)}. \tag{47}$$

Use the orthogonality conditions

$$\mathbb{E}_t[u_{t+1}u_{t+1}'] = \Sigma, \quad \mathbb{E}_t[u_{t+1}\varepsilon^{p\prime}_{t+1}] = \mathbf{0}, \quad \mathbb{E}_t[u_{t+1}\varepsilon^{y\prime}_{t+1}] = \mathbf{0}, \quad \mathbb{E}_t[\varepsilon^p_{t+1}\varepsilon^{y\prime}_{t+1}] = \mathbf{0},$$

[28] Please note that $\mathcal{A}_{n,t}$ is a different object than A_t.

and define

$$\Sigma_p \equiv \mathbb{E}_t[\varepsilon^p_{t+1}\varepsilon^{p\prime}_{t+1}], \qquad \Sigma_y \equiv \mathbb{E}_t[\varepsilon^y_{t+1}\varepsilon^{y\prime}_{t+1}].$$

Then (2) = (3) = (4) = (6) = (7) = (8) = $\mathbf{0}$, and

$$\mathcal{B}_{n,t} = A_t \Sigma A_t' + \Sigma_p + \Sigma_y. \tag{48}$$

The vector of variance contributions that enters the Euler equations is the diagonal of $\mathcal{B}_{n,t}$:

$$\operatorname{diag}\big(\mathcal{B}_{n,t}\big) = \operatorname{diag}\big(A_t \Sigma A_t'\big) + \operatorname{diag}\big(\Sigma_p\big) + \operatorname{diag}\big(\Sigma_y\big). \tag{49}$$

Recall

$$\mathcal{A}_{n,t} = -\tfrac{1}{2}\,\mathbf{1}_n(\lambda_t'\Sigma\lambda_t) + (\gamma_0 + \gamma_1 F_t) + \big(b_0 + b_1(c + \rho F_t)\big),$$

and

$$\operatorname{diag}\big(\mathcal{B}_{n,t}\big) = \operatorname{diag}\big(A_t \Sigma A_t'\big) + \operatorname{diag}\big(\Sigma_p\big) + \operatorname{diag}\big(\Sigma_y\big), \quad A_t \equiv -\mathbf{1}_n\lambda_t' + \Gamma, \quad \Gamma \equiv \gamma_2 + b_1.$$

Expand the diagonal term explicitly. For any $n \times k$ matrix Γ,

$$\operatorname{diag}\big((-\mathbf{1}_n\lambda_t' + \Gamma)\Sigma(-\mathbf{1}_n\lambda_t' + \Gamma)'\big) = \mathbf{1}_n(\lambda_t'\Sigma\lambda_t) - 2\,\Gamma\Sigma\lambda_t + \operatorname{diag}(\Gamma\Sigma\Gamma').$$

Hence

$$\tfrac{1}{2}\operatorname{diag}\big(\mathcal{B}_{n,t}\big) = \tfrac{1}{2}\,\mathbf{1}_n(\lambda_t'\Sigma\lambda_t) - \Gamma\Sigma\lambda_t + \tfrac{1}{2}\operatorname{diag}(\Gamma\Sigma\Gamma') + \tfrac{1}{2}\operatorname{diag}(\Sigma_p) + \tfrac{1}{2}\operatorname{diag}(\Sigma_y). \tag{50}$$

Add the mean and variance contributions:

$$\mathcal{A}_{n,t} + \tfrac{1}{2}\,\mathrm{diag}\big(\mathcal{B}_{n,t}\big) = \Big[-\tfrac{1}{2}\,\mathbf{1}_n(\lambda_t'\Sigma\lambda_t)\Big] + \Big[\tfrac{1}{2}\,\mathbf{1}_n(\lambda_t'\Sigma\lambda_t)\Big] \\ + \underbrace{\big(\gamma_0 + b_0 + b_1 c\big)}_{\text{constant}} + \underbrace{\big(\gamma_1 + b_1\rho\big)F_t}_{\text{loads on } F_t} - \underbrace{\Gamma\Sigma\lambda_t}_{\text{depends on } \lambda_t} + \tfrac{1}{2}\,\mathrm{diag}(\Gamma\Sigma\Gamma') + \tfrac{1}{2}\,\mathrm{diag}(\Sigma_p + \Sigma_y). \tag{51}$$

The quadratic term $\mathbf{1}_n(\lambda_t'\Sigma\lambda_t)$ cancels exactly. Using the affine price-of-risk $\lambda_t = \lambda + \Lambda F_t$, we obtain:

$$\mathcal{A}_{n,t} + \tfrac{1}{2}\,\mathrm{diag}\big(\mathcal{B}_{n,t}\big) = \underbrace{\Big[\gamma_0 + b_0 + b_1 c - \Gamma\Sigma\lambda + \tfrac{1}{2}\,\mathrm{diag}(\Gamma\Sigma\Gamma' + \Sigma_p + \Sigma_y)\Big]}_{\mathcal{C}_0 \in \mathbb{R}^{n\times 1}} + \underbrace{\Big[\gamma_1 + b_1\rho - \Gamma\Sigma\Lambda\Big]}_{\mathcal{C}_1 \in \mathbb{R}^{n\times k}} F_t. \tag{52}$$

Therefore, the Euler restrictions $\mathcal{A}_{n,t} + \frac{1}{2}\,\mathrm{diag}(\mathcal{B}_{n,t}) = \mathbf{0}$ are equivalent to the two sets of parameter restrictions:

$$\mathcal{C}_0 \;=\; \gamma_0 + b_0 + b_1 c \;-\; (\gamma_2 + b_1)\Sigma\lambda \;+\; \tfrac{1}{2}\,\mathrm{diag}\big((\gamma_2 + b_1)\Sigma(\gamma_2 + b_1)' + \Sigma_p + \Sigma_y\big) \;=\; \mathbf{0}_{n\times 1} \tag{53}$$

$$\mathcal{C}_1 \;=\; \gamma_1 + b_1\rho \;-\; (\gamma_2 + b_1)\Sigma\Lambda \;=\; \mathbf{0}_{n\times k} \tag{54}$$

B.2 Model Output

As explained in Section 2.2, the model delivers HTM risk premia $\theta_t^{(n)}$, HTM expected dividend growth $g_t^{(n)}$ and equity yields $e_t^{(n)}$. In this section, we explain how these quantities relate to the state vector F_t and its shocks, as well as parameters Ξ.

$e_t^{(n)}$ and $\theta_t^{(n)}$ can be derived as closed-form functions of the state vector F_t and the model parameters.[29] Equity yields $e_t^{(n)}$ can be written as:

$$e_t^{(n)} = \frac{1}{n} \ln \left(\frac{D_t}{P_t^{(n)}} \right) = \frac{1}{n} \left[\ln(\exp(y_t) - 1) - \ln \left(\frac{P_t^{(n)}}{P_t} \right) \right],$$

where $y_t \equiv \ln \left(1 + \frac{D_t}{P_t}\right)$ is the log dividend-price ratio.

The price of the dividend strip relative to the price of the portfolio is derived as follows. Consider the price for any portfolio n-periods from now: $P_{t+n} = P_t \exp(\sum_{i=1}^{n} \Delta p_{t+i})$. The dividend yield n-periods from now for such portfolio is $\frac{D_{t+n}}{P_{t+1}} = [\exp(y_{t+n}) - 1]$. Both terms follow directly from the definitions of the model. Combining them gives:

$$\frac{D_{t+n}}{P_t} = \frac{D_{t+n}}{P_{t+1}} \frac{P_{t+1}}{P_t} = [\exp(y_{t+n}) - 1] \exp(\sum_{i=1}^{n} \Delta p_{t+i}). \tag{55}$$

Thus, the price of a claim to D_{t+n}, $P_t^{(n)}$, as a fraction of the price of the portfolio, P_t, is:

$$\begin{aligned}
\frac{P_t^{(n)}}{P_t} \equiv w_t^n &= \mathbb{E}_t^Q \left(\frac{D_{t+n}}{P_t} \exp(-\sum_{i=1}^{n} r_{f,t+i-1}) \right) \\
&= \mathbb{E}_t^Q \left([\exp(y_{t+n}) - 1] \exp(\sum_{i=1}^{n} \Delta p_{t+i}) \exp(-\sum_{i=1}^{n} r_{f,t+i-1}) \right) \\
&= \mathbb{E}_t^Q \left(\exp \left(y_{t+n} + \sum_{i=1}^{n} (\Delta p_{t+i} - r_{f,t+i-1}) \right) \right) - \mathbb{E}_t^Q \left(\exp \left(\sum_{i=1}^{n} (\Delta p_{t+i} - r_{f,t+i-1} \right) \right) \\
&= \exp(a_{n,1} + d_{n,1} F_t) - \exp(a_{n,2} + d_{n,2} F_t)
\end{aligned} \tag{56}$$

where $\Delta p_{t+1} \equiv \ln(P_{t+1}/P_t)$ is the capital gain from t to $t+1$, $\mathbb{E}_t^Q$ denotes expectation under the risk-neutral measure. The parameters a_n and d_n satisfy the following recursions:

$$a_n = a_{n-1} + \gamma_0^\star + d_{n-1} c^\star + \frac{1}{2}(d_{n-1} + \gamma_2)\Sigma(d_{n-1} + \gamma_2)' + \frac{1}{2}\sigma_\nu \tag{57}$$

$$d_n = \gamma_1^\star + d_{n-1}\rho^\star. \tag{58}$$

[29]The HTM expected dividend growth $g_t^{(n)}$ can thus be obtained as a residual.

With initial values $a_{0,1} = b_0 + \frac{1}{2}(\sigma_r^2 - \sigma_\nu^2), d_{0,1} = b_1, a_{0,2} = 0, d_{0,2} = 0, \sigma_r^2 = var(\epsilon_t), \sigma_\nu^2 = var(\nu_t)$. In these formulas, stars indicate risk-neutral parameters.[30] The risk premium $\theta_t^{(n)}$, is computed as:

$$\begin{aligned}
\ln \mathbb{E}_t \left[\frac{R_{t:t+n}^{(n)}}{R_{f,t:t+n}} \right] &= \ln \mathbb{E}_t \left[\frac{D_{t+n}}{P_t} \frac{1}{R_{f,t:t+n}} \right] - \ln \left[\frac{P_t^{(n)}}{P_t} \right] \\
&= \ln \left[\exp(\widehat{a}_{n,1} + \widehat{d}_{n,1} F_t) - \exp(\widehat{a}_{n,2} + \widehat{d}_{n,2} F_t) \right] \\
&\quad - \ln \left[\exp(a_{n,1} + d_{n,1} F_t) - \exp(a_{n,2} + d_{n,2} F_t) \right],
\end{aligned}$$

where the notation $\widehat{x}$ refers to the physical counterpart to the risk-neutral parameter x.

C Validation of GKK Model on UK Data

In Section 3.1, we validated the model by comparing the model-implied 5-year forward equity yield to the yield implied by traded FTSE 100 dividend futures. In this section, we extend the comparison to the 2-year and 7-year forward yields.

[Insert Figure 14 here]

Figures 14 and 15 plot the model-implied yields (solid blue) alongside their market-implied counterparts (dotted orange). The results confirm that the alignment documented in the 5-year case extends to other maturities. These findings further support the validity of the GKK model in capturing the maturity-specific pricing of dividend strips using only UK equity data.

[Insert Figure 15 here]

[30] $\gamma_0^* = \gamma_0 - \gamma_2 \Sigma\lambda, \quad \gamma_1^* = \gamma_1 - \gamma_2 \Sigma\Lambda, \quad c^* = c - \Sigma\lambda, \quad \rho^* = \rho - \Sigma\Lambda$

D Financial Integration and the 1980 Breakpoint

We use 1980 as a reference point because it follows a discrete policy break and aligns with a sustained rise in cross-border financial integration. To justify this choice, we proceed in three steps. We first review economic evidence from the literature on financial liberalization and integration. We then show that our findings are robust to alternative cutoff years. Finally, we present formal structural break tests, both at pre-specified and data-driven breakpoints, to validate the 1980 benchmark and to place it in historical perspective.

In the United Kingdom, capital controls were abolished in 1979; the literature documents a subsequent and persistent increase in international financial integration. On the de jure, indices that codify current- and capital-account regulations for OECD countries and, later, broader cross-country samples show a pronounced shift toward legal openness during the 1980s and 1990s (Quinn and Inclan, 1997; Quinn, 1997).[31] On the de facto side, time-series evidence on current-account dynamics indicates very limited international capital mobility under Bretton Woods and a return to much greater flexibility in the floating-rate era after 1973, approaching late-nineteenth-century levels (Taylor, 2002).[32]

Prices and quantities move in ways characteristic of a higher financial-integration regime. For the United States and the United Kingdom, and more broadly across the G7, Colacito and Croce (2013) document a discrete post-1970 shift in key moments: real exchange-rate volatility is roughly twice as large, the uncovered interest parity slope falls sharply, and the Backus–Smith correlation declines substantially. These price–quantity patterns are consistent with deeper international risk sharing and greater financial integration. Historical panel evidence also links capital-account liberalization to higher international equity-return

[31] Quinn and Inclan (1997) construct legal openness indexes for current- and capital-account transactions for 20 OECD countries over 1950–1988 using systematic codings of the IMF *AREAER*, documenting substantial liberalization beginning in the late 1970s and accelerating through the 1980s. Quinn (1997) extend the coding to a 64-country sample through 1994 and show similar global liberalization patterns, together with the political and economic correlates of these regulatory changes.

[32] Taylor (2002) estimate regime-dependent current-account adjustment in a long-run panel and show slower adjustment under Bretton Woods, with a reversion toward the faster pre-1914 speeds in the post-1973 float, consistent with higher effective financial integration.

correlations, reinforcing the interpretation that markets became more financially integrated as legal barriers fell (Quinn and Voth, 2008).[33]

Taken together, the 1979 policy break, the subsequent legal liberalization in the 1980s and 1990s, the de facto rise in international adjustment and price–quantity co-movements after 1970, and the panel evidence on cross-market comovement provide a coherent economic and empirical basis for using 1980 as the benchmark year for analyzing changes in the equity term structure associated with financial integration.

Turning to robustness, we show that our main result is not sensitive to the exact choice of the cutoff year. Figure 19 compares the term structure of hold-to-maturity (HTM) risk premia before and after three alternative cutoff years—1970, 1980, and 1990—both unconditionally and conditional on business cycle phases (expansions and recessions). Across all cutoff choices, the post-cutoff term structure is consistently steeper than its pre-cutoff counterpart, with the difference concentrated at short maturities. This pattern holds whether we consider the full sample, expansions only, or recessions only. These results indicate that our main finding—a steeper term structure in the financially integrated regime—is not an artifact of selecting 1980 as the breakpoint. Rather, it reflects a persistent change in pricing dynamics that is robust to reasonable variation in the cutoff year and is consistent with the timeline of financial integration documented in the literature cited above.

[Insert Figure 19 here]

To complement this robustness check, we subject our choice of 1980 to formal statistical evaluation. Specifically, we test for the presence of a structural break in the mean of the equity term structure slope around that year. We implement a series of Chow tests using six candidate break dates spanning from October 1979 to March 1980. October 1979 is of particular economic relevance, as it coincides with the formal abolition of U.K. capital

[33] Quinn and Voth (2008) assemble a multi-country panel for 1890–2001 using a de jure capital-account openness index and show that bilateral equity-return correlations are significantly higher in more open regimes and that the entire correlation distribution shifts rightward under openness.

controls. For each candidate date, we estimate the model separately for the pre- and post-break subsamples, as well as for the pooled sample, under the null hypothesis of parameter stability. Rejection of the null indicates a statistically significant change in the mean slope at the specified breakpoint. Table 4 reports the results:

[Insert Table 4 here]

As shown in the table, all six specifications yield highly significant F-statistics, with p-values effectively equal to zero. While the precise magnitude of the test statistic varies slightly across breakpoints, the overall pattern is stable and consistently rejects the null of no change. These results provide statistical support for a structural break in the slope around 1980, complementing the historical narrative of financial integration and reinforcing our interpretation of the post-1980 period as an economically distinct regime characterized by deeper international integration.

Because the Chow test relies on a pre-specified breakpoint, one may still question whether the analysis is robust to the choice of date. To address this concern, we implement the multiple structural break tests proposed by Bai and Perron (1998, 2003), which identify breaks in the mean of the equity term structure slope without requiring prior assumptions about their location. Table 5 reports the results.

[Insert Table 5 here]

Panel A presents the global *supF* test of the null hypothesis of no structural breaks against an unspecified alternative. The test rejects strongly ($supF = 203.0$, $p < 0.01$), indicating that the mean slope changes over the sample. Panel B reports sequential $supF(\ell + 1 \mid \ell)$ statistics for models with up to five breaks. All values are highly significant, and the decision rule consistently favors the inclusion of an additional break at each step. This pattern provides further evidence of structural changes in the mean slope. To determine the number of breaks, Panel C reports the residual sum of squares (RSS) and Bayesian Information

Criterion (BIC) for models with zero to five breaks. The BIC reaches its minimum at four breaks, which we adopt as the preferred specification. Panel D lists the estimated break dates and their associated 95% confidence intervals: July 1910, August 1932, August 1955, and February 1983. Notably, one of the estimated breaks—February 1983—falls squarely within the window considered in our Chow analysis, providing strong data-driven validation that 1980 is indeed a reasonable breakpoint for our study.

The earlier break dates—1910, 1932, and 1955—likely correspond to major historical disruptions: the run-up to World War I, the Great Depression of the 1930s, and the aftermath of World War II, respectively. These periods were characterized by severe geopolitical and macroeconomic instability. The first break, dated to July 1910, appears to align with the tail end of the so-called "first era of globalization" at the turn of the 20th century. While this timing is intriguing and may reflect meaningful shifts in international financial conditions, the impending outbreak of World War I and the limited development of global capital markets during that era complicate interpretation. Moreover, we emphasize that asset trade during the early 20th century was likely less liquid and subject to greater frictions than in modern markets. Theoretical models of financial integration typically assume that investors can arbitrage freely across countries—an assumption that is plausible for the post-1980 environment but less so for 1910. We therefore view the earlier break dates as historically important but analytically more difficult to connect to the mechanisms emphasized in our framework.

In summary, we provide both economic motivation and empirical validation for focusing on 1980 as the reference point for our analysis. Historically, this date marks the beginning of a well-documented period of financial liberalization and increasing international capital mobility. Empirically, our results are robust to alternative cutoff years, and both the Chow and Bai–Perron tests support the presence of a structural change in the slope of the equity term structure around this period. While earlier break dates may capture meaningful shifts in the global financial landscape, they are less directly linked to the mechanisms of modern financial integration emphasized in our framework. Overall, the evidence supports our interpretation of the post-1980 period as an economically distinct regime shaped by deeper financial integration. The revised version of the paper incorporates this extended discussion and the accompanying new results.

E Theoretical Model

We use the following transformation to obtain a stationary model which we solve using projection methods (see Pohl et al. (2021)). Define $V^h = \frac{U_t^h}{X_t^\alpha Y_t^{1-\alpha}}$ so that

$$V_t^h = \left[(1-\delta) \left(\frac{C_t^h}{X_t^\alpha Y_t^{1-\alpha}} \right)^{1-\frac{1}{\psi}} + \delta E_t \left(\left(V_{t+1}^h \left(\frac{X_{t+1}}{X_t} \right)^\alpha \left(\frac{Y_{t+1}}{Y_t} \right)^{1-\alpha} \right)^{1-\gamma} \right)^{\frac{1-\frac{1}{\psi}}{1-\gamma}} \right]^{\frac{1}{1-\frac{1}{\psi}}}.$$

A similar expression can be derived for $V^f = \frac{U_t^f}{Y_t^\alpha X_t^{1-\alpha}}$. From equation (12) it follows that

$$\frac{C_t^h}{X_t^\alpha Y_t^{1-\alpha}} = \left(\frac{x_t^h}{X_t} \right)^\alpha \left(\frac{y_t^h}{Y_t} \right)^{1-\alpha}. \tag{59}$$

The stochastic discount factor (17) in terms of V_t^h is given by

$$M_{t+1}^i = \delta \left(\frac{C_{t+1}^i}{C_t^i} \right)^{-\frac{1}{\psi}} \left(\frac{V_{t+1}^i \left(\frac{X_{t+1}}{X_t} \right)^\alpha \left(\frac{Y_{t+1}}{Y_t} \right)^{1-\alpha}}{E_t \left(\left(V_{t+1}^i \left(\frac{X_{t+1}}{X_t} \right)^\alpha \left(\frac{Y_{t+1}}{Y_t} \right)^{1-\alpha} \right)^{1-\gamma} \right)^{\frac{1}{1-\gamma}}} \right)^{\frac{1}{\psi}-\gamma}.$$

Consumption growth in country h can be written as

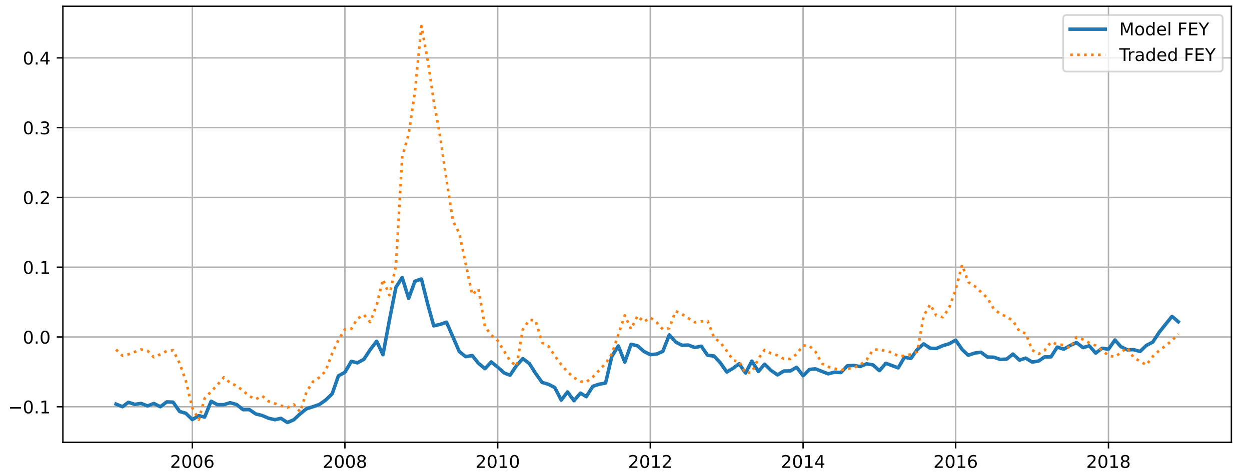
$$\frac{C_{t+1}^h}{C_t^h} = \frac{\frac{C_{t+1}^h}{X_{t+1}^\alpha Y_{t+1}^{1-\alpha}}}{\frac{C_t^h}{X_t^\alpha Y_t^{1-\alpha}}} \left(\frac{X_{t+1}}{X_t} \right)^\alpha \left(\frac{Y_{t+1}}{Y_t} \right)^{1-\alpha} \tag{60}$$

where the first part captures the change in the consumption share and the second part consumption growth of the aggregate economy.

We use the collocation projection method presented in Pohl et al. (2021) to solve for the equibrium. For that, we first solve for value functions and asset prices in the representative agent cases. We use those solutions as initial guesses for the two country economy. The state space is 3-dimensional and consists of $\{W_t^h, \lambda_t, CI_t\}$. We determine the approximation interval of λ_t and CI_t by simulating long time series of the processes (1,000,000 periods). For the approximation interval of W_t^h we use $[0.001, 0.999]$ to adequately capture dynamics close to the boundary cases. We use a uniform grid for all three state variables with 31 nodes for W_t^h, 7 nodes for z_t and 5 nodes for CI_t. We use Gauss-Hermite quadrature with 5 nodes to compute expectations over normally distributed variables. To approximate expectations over the Poisson disasters, we assume a maximum number of disaster per period of 2. We verify that non of the parameter choices of our computational procedure has a significant impact on model outcomes. For that, we increased the number of collocation nodes, approximation interval and approximation nodes in the quadrature methods and found that it does not materially affect our models outcomes.

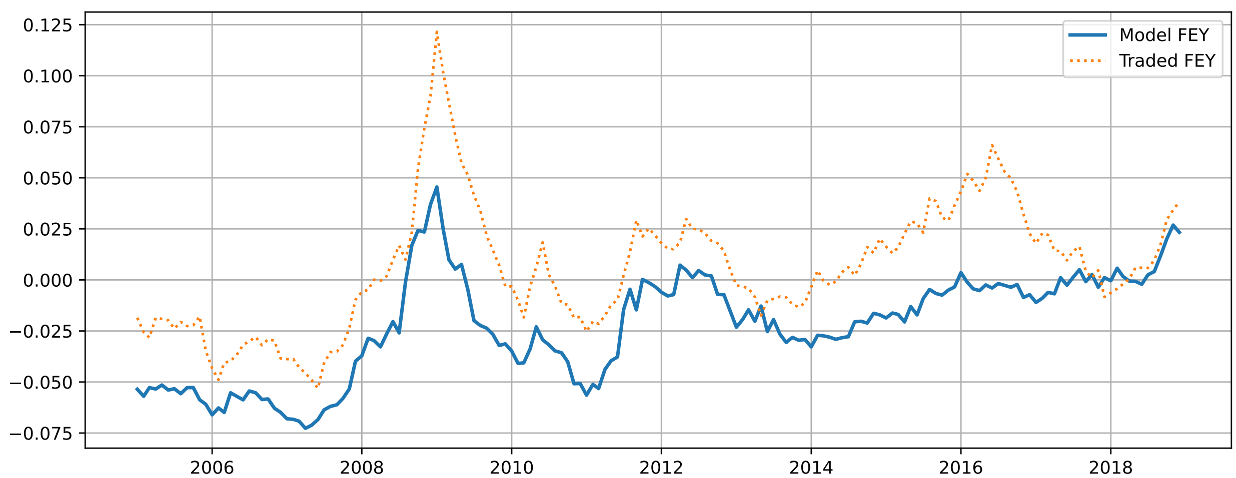
F Additional Figures

Figure 14: Forward Equity Yields (2y): Model vs. Traded Dividend Futures.



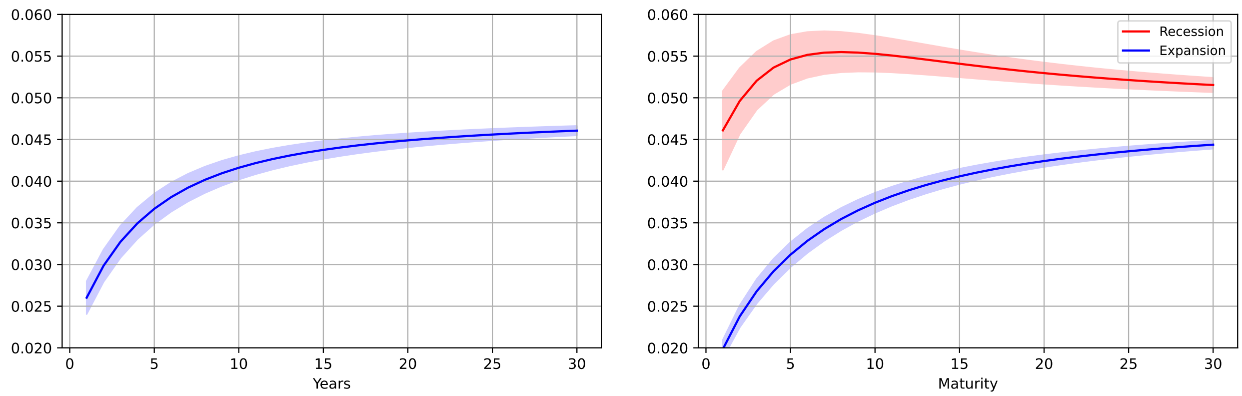
This figure compares the 2-year forward equity yield implied by our model (solid blue) with the yield extracted from traded dividend futures (dotted orange).

Figure 15: Forward Equity Yields (7y): Model vs. Traded Dividend Futures.



This figure compares the 7-year forward equity yield implied by our model (solid blue) with the yield extracted from traded dividend futures (dotted orange).

Figure 16: Term Structure of HTM Risk Premia

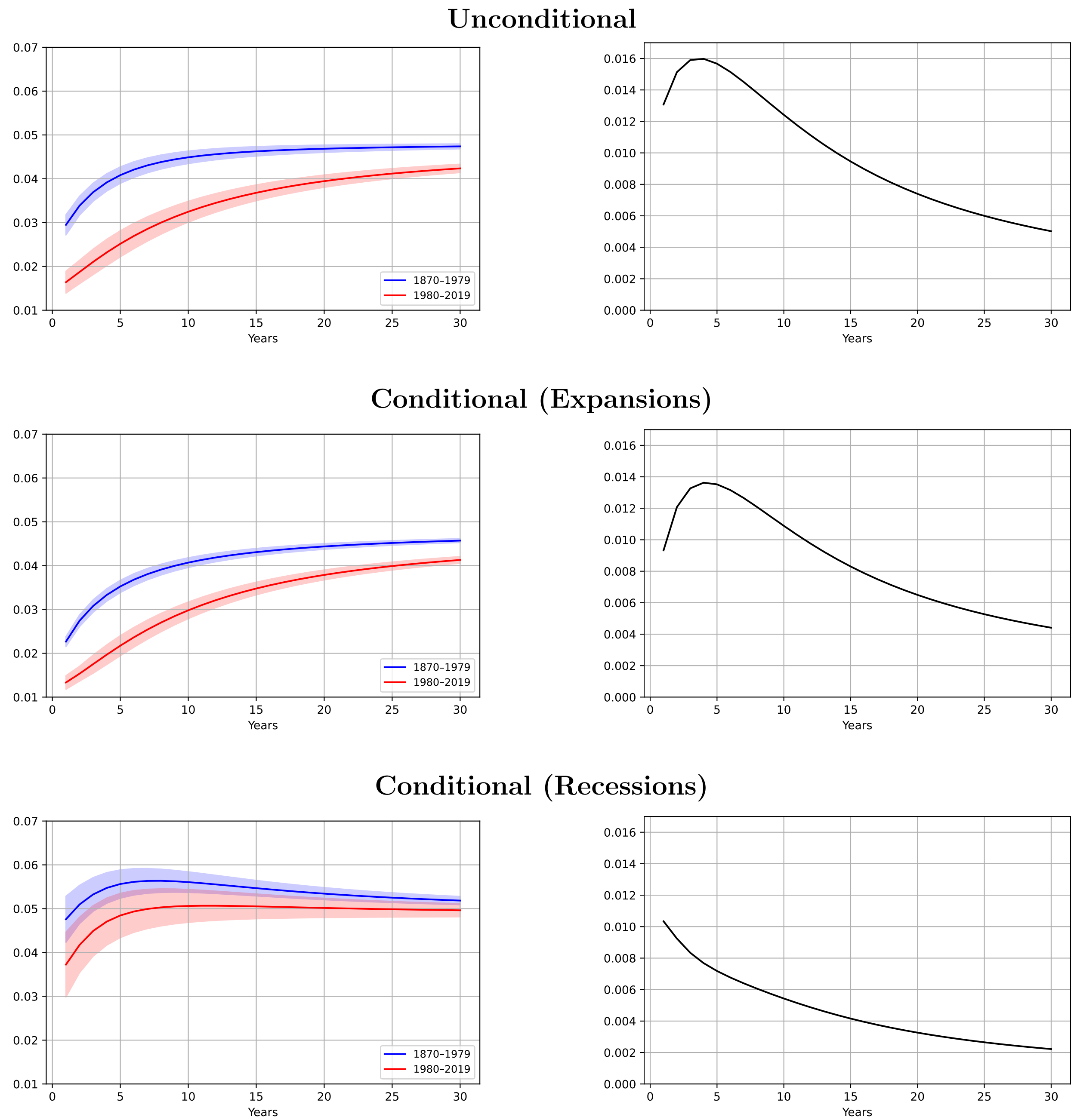


(a) Unconditional

(b) Conditional

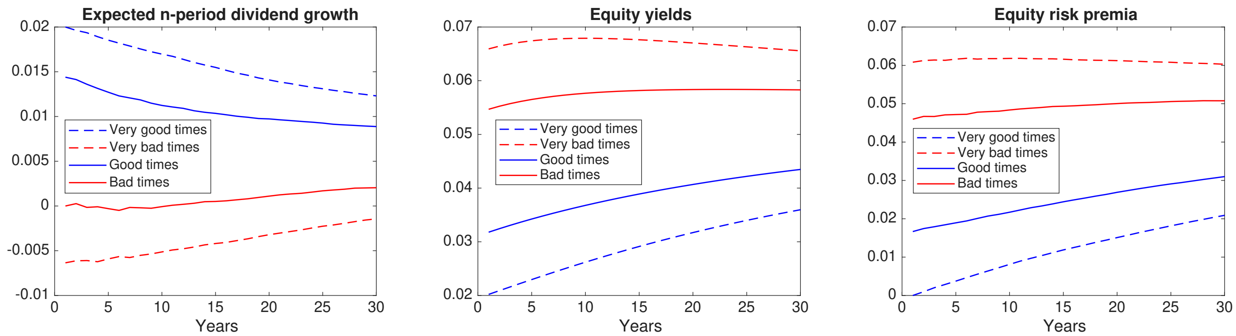
The figure plots the term structure of HTM risk premia. Panel (a) shows the unconditional premia over the full sample (1870–2019). Panel (b) shows the conditional premia during expansions and recessions, based on a recession index based exclusively on the price-dividend ratio, classifying months with price-dividend ratio below the 25th percentile as recessions. Bands denote 95% confidence intervals based on HAC-robust standard errors.

Figure 17: Term Structure of HTM Risk Premia Before and After 1980.



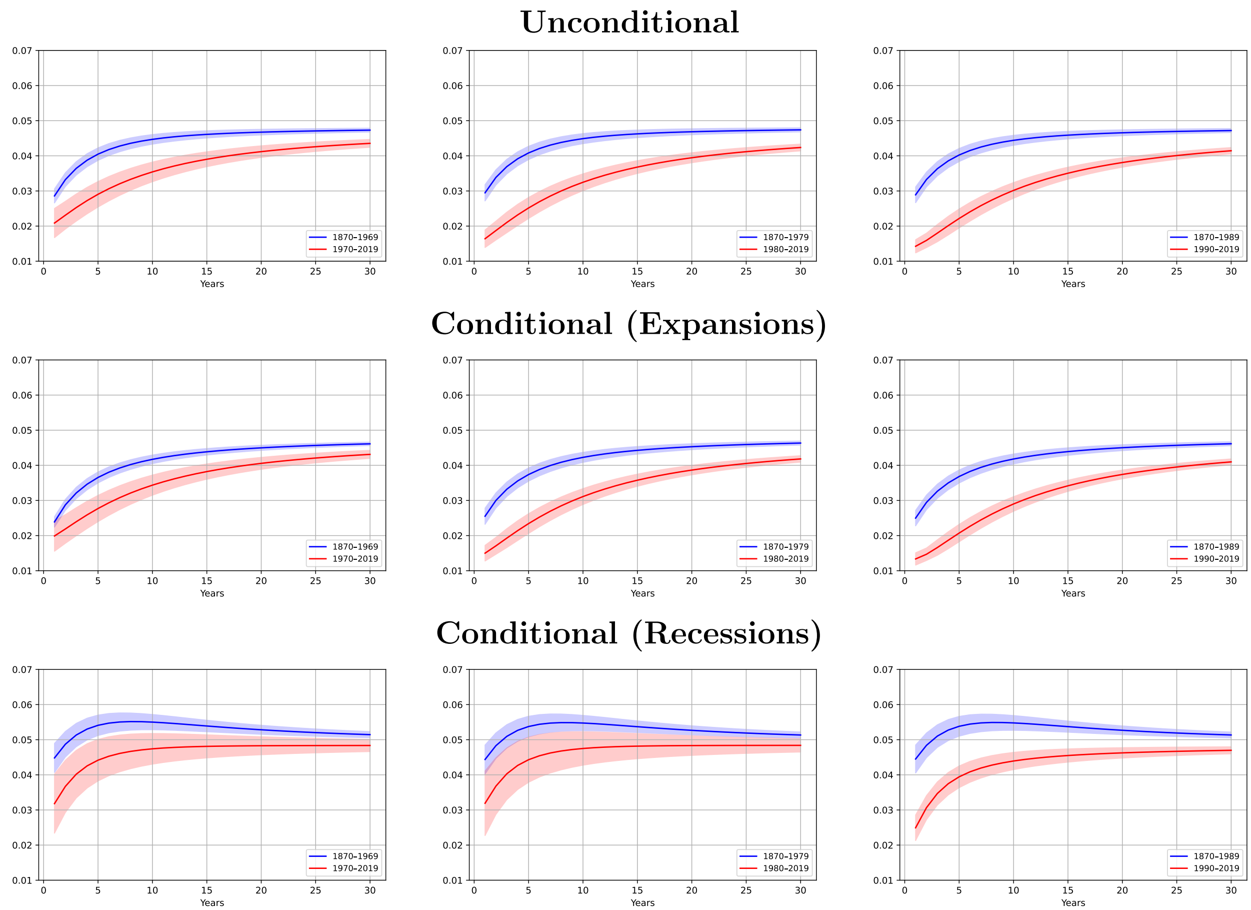
This figure plots the term structure of HTM risk premia for the 1870–1979 and 1980–2019 periods. The left column displays the estimated term structures, while the right column shows the difference (1870–1979 minus 1980–2019). The top row presents unconditional estimates for the full sample. The second and third rows condition on expansions and recessions, respectively, using a recession index based exclusively on the price-dividend ratio, classifying months with price-dividend ratio below the 25th percentile as recessions. Shaded areas denote 95% confidence bands based on HAC-robust standard errors.

Figure 18: Time variation of the equity term structure



The figure plots model implied n-period expected dividend growth (left), equity yields (center) and equity risk premia (right). Results are shown for very good times ($\lambda_t = 0.0$), good times ($\lambda_t = 0.02$), bad times ($\lambda_t = 0.06$) and very bad times ($\lambda_t = 0.08$). Consumption allocations are set to $W_t = 0.03$ Model parameters are provided in Table 3.

Figure 19: Term Structure of HTM Risk Premia Before and After Alternative Cutoff Years.



This figure compares the term structure of hold-to-maturity (HTM) risk premia before and after three alternative cutoff years: 1970 (left column), 1980 (middle), and 1990 (right). Each panel contrasts the pre- and post-cutoff term structures. The first row reports unconditional estimates. The second and third rows condition on expansions and recessions, respectively, based on the recession index described in the paper. Shaded areas denote 95% confidence bands constructed using HAC-robust standard errors.

G Additional Tables

Table 4: Chow Test for a Break in the Slope of the Equity Term Structure

Break	N_{pre}	N_{post}	RSS_{pooled}	RSS_{pre}	RSS_{post}	$F-stat$	$p-val$
October 1979	1299	471	0.31	0.25	0.04	127.50	<0.001
November 1979	1300	470	0.31	0.25	0.04	129.94	<0.001
December 1979	1301	469	0.31	0.25	0.04	133.25	<0.001
January 1980	1302	468	0.31	0.25	0.04	135.25	<0.001
February 1980	1303	467	0.31	0.25	0.04	136.67	<0.001
March 1980	1304	466	0.31	0.26	0.03	141.85	<0.001

This table reports the results of Chow (1960) tests for a structural break in the mean slope of the equity term structure, defined as $\theta_t^{(30)} - \theta_t^{(1)}$, following the notation used in the paper. Six candidate break dates between October 1979 and March 1980 are considered. For each date, the table shows the number of observations before and after the breakpoint, the residual sum of squares (RSS) for the pooled model and for each subsample, the resulting F-statistic, and the associated p-value.

Table 5: Multiple Structural Break Tests in the Slope of the Equity Term Structure

Panel A: Global supF Test

Statistic	p-value
203.0	<0.01

Panel B: Sequential supF($\ell+1|\ell$) Tests

ℓ	Test Statistic	p-value	Decision
1	78.6	<0.01	Add Break
2	153.0	<0.01	Add Break
3	159.0	<0.01	Add Break
4	172.0	<0.01	Add Break
5	397.0	<0.01	Add Break

Panel C: Model Selection Based on BIC

m	RSS	BIC
0	0.314	−15,279
1	0.282	−15,464
2	0.267	−15,550
3	0.256	−15,623
4	**0.246**	**-15,682**
5	0.264	−15,550

Panel D: Break Dates with 95% Confidence Intervals

Break #	Estimated Break Date	95% CI Low	95% CI High
1	Jul 1910	Jul 1909	Aug 1910
2	Aug 1932	Dec 1931	Aug 1933
3	Aug 1955	Sep 1948	Oct 1956
4	Feb 1983	Dec 1982	Aug 1985

This table reports the results of the structural break tests of Bai and Perron (1998, 2003) applied to the slope of the equity term structure, defined as $\theta_t^{(30)} - \theta_t^{(1)}$, over the period July 1871 to December 2018. We use a trimming parameter of $\varepsilon = 0.15$. Panel A presents the global *supF* test for the null hypothesis of no structural breaks. Panel B shows the sequential $supF(\ell + 1 \mid \ell)$ test statistics for models with up to five breaks. Panel C reports the residual sum of squares (RSS) and Bayesian Information Criterion (BIC) for model selection. Panel D provides the estimated break dates and their 95% confidence intervals.