

A Factor Model Approach to Pricing Municipal Bond Portfolios

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Abstract

We adapt factor-pricing methods to state-level municipal bond portfolios and decompose risk premia into persistent, aggregate, and relative components. In normal times a national market factor dominates time-series variation, while style exposures influence the return cross-section. Treasury factors add no pricing power, consistent with segmentation from asymmetric federal taxation, and we detect a convenience premium that suppresses returns in high-tax states. During COVID the baseline structure breaks down: returns are priced on the cross-state distribution of pandemic outcomes in place of baseline factors, indicating a shift in risk transmission. Municipal risk structures are thus tractable but regime dependent.

1 Introduction

This paper develops an asset-pricing framework for state municipal portfolios by adapting a factor model widely used in pricing equities, sovereign debt and corporate credit. This approach is well suited for answering two related asset-pricing questions: whether the returns on municipal portfolios can be understood with a small number of common risk factors and how to quantify different types of risk in municipal bond portfolios.

The paper makes two departures from the traditional approach in the municipal bond literature that typically focuses on pricing individual issues and emphasizes tax treatment of interest income. The first departure is pricing a state’s entire municipal bond portfolio rather than individual bonds. State-level aggregation makes return co-movement more transparent to observe and analyze. The second departure is to use portfolio-based factors, familiar from the literature on other asset classes, which facilitates cross-asset comparisons. The model includes the return on a broad national municipal bond index, which plays the role of the market portfolio, as well as style factors – value, carry, low-risk, and momentum – constructed by sorting state portfolios on observable characteristics. The value factor is associated with cheap assets outperforming expensive ones; the carry factor captures returns from holding high-yielding assets absent adverse price changes; the low-risk factor is related to volatility; and the momentum factor may be related to inertia in transactions or information flows. Additionally, Treasury slope and convexity factors control for term-structure risks that may provide information beyond municipal returns. The model then decomposes risk premia into persistent, aggregate time-varying, and relative time-varying components.

Our first set of results derives from the risk premium decomposition. In normal times, state municipal portfolios are dominated by a common national municipal-market component. The state betas are concentrated between approximately 0.8 and 1.1, so that individual state portfolios behave similarly to the aggregate market. The high mean and low variance of market betas is the reason why the aggregate risk premium is the dominant source of variance in expected returns. Thus state municipal bond returns mostly move in

concert over time in response to national interest rates, macroeconomic conditions, federal policy, and other aggregate shocks.

By contrast, the cross-sectional return variance is mostly due to smaller but substantially more heterogeneous exposures to style factors. Average style betas range from approximately 0.02 to 0.12 in absolute value while individual exposures are roughly in the -1.5 to 1.5 range. Risk premia on style factors are significant, indicating that other states' returns are useful basis assets for constructing municipal risk-mimicking portfolios.

Persistent risk premia, however, account for only slightly more than 1 percent of the cross-sectional return variance. Persistent yield spreads across states thus do not translate into persistent differences in realized holding returns. High-yield state portfolios may experience larger adverse price adjustments or call-related effects, while low-yield portfolios experience more moderate repricing. This finding is consistent with competitive, no-arbitrage pricing and shows how yield-based interpretations of municipal risk premia may be potentially misleading.

Our second set of results concerns taxation and market segmentation. Tax clientele effects have been a well known feature of municipal bond markets and our model detects them in two related ways. In the current tax regime, municipal bond income is exempt from federal tax but Treasury income is not, making the tax treatment asymmetric. Treasury slope and convexity factors have little incremental pricing power for municipal bonds, consistent with market segmentation. However, under an alternative tax policy where the federal tax treatment of Treasuries and municipal bonds becomes symmetric, the model re-allocates expected returns towards the Treasury factors and away from the municipal market factor. This suggests that asymmetric taxation helps sustain segmented markets. Beyond this, we detect a statistically significant convenience premium – a general equilibrium effect that suppresses returns on municipal bonds in states where after-tax returns on taxable investments are low.

Our third set of results derives from applying the framework to the COVID period and asking whether the baseline factor structure is invariant to the nature of aggregate

shocks. The pandemic is uniquely suited to this question because it combined a large national shock with heterogeneous effects by state. We find that the structure is not invariant: the composition of risk premia shifts, investors price the cross-section on a state variable absent from conventional factors, and the risk transmission mechanism changes. We also show that the model accommodates the regime change once factors built on the new state variable are added.

The baseline model structure breaks down during the spring and summer of 2020. As yields and state portfolio returns reshuffled repeatedly, the carry and value factors that account for relative returns in normal times become collinear with the market factor, and the market factor's risk premium becomes unstable. The one baseline factor that appears to gain pricing power during the spring 2020 crash, the Treasury slope, does so only because the model omitted a state variable.

That state variable, as it turns out, is the cross-state distribution of pandemic outcomes. A state's own outcomes do not explain pricing errors, but long-short portfolios sorted on the distribution of cases and intensive-care occupancy across states carry significant premia and account for a third of the cross-sectional return variance in June-August 2020. Once these COVID factors are included, the apparent pricing power of the Treasury slope disappears, as the cross-sectional variation it had absorbed is attributed to the differential impact of the pandemic. Municipal returns thus depended on the entire distribution of pandemic scenarios across states rather than on in-state conditions, even though the pandemic and the policy response were national events. By late 2020 the model reverts toward a one-factor structure with relative risk premia still above normal levels.

The pandemic also altered the composition of risk, and it did so in two stages. The March crash shows greater aggregate dominance: relative risk premia fell to 1-1.5 percent of predicted-return variance, below their pre-COVID share of about 2.5 percent, so that states moved almost entirely together. The summer recovery had the opposite effect, raising the relative risk premium share to 13 percent as states repriced at different speeds; in the cross-section, style and COVID factors together explained over 70 percent of return variance,

compared with 55 percent for style factors alone before the pandemic.

In sum, municipal portfolios share a strong national factor structure in normal times, but the composition of risk premia, the state variables investors price on, and the transmission of risk are regime dependent, and the factor model provides a tractable way to detect and accommodate such regime changes.

Related Literature. Our work bridges the literature on pricing municipal bonds and common factor approaches to pricing other asset classes.

The dominance of the municipal market factor is consistent with evidence that fixed-income returns contain a large common component. Litterman and Scheinkman (1991) show that a small number of common factors account for most Treasury return variation, while Driessen, Melenberg, and Nijman (2003), Diebold, Li, and Yue (2008), and Jotikasthira, Le, and Lundblad (2015) document substantial commonality across national government-bond markets and yield curves. In sovereign credit markets, Longstaff et al. (2011) find that global financial conditions account for a large portion of sovereign spread variation relative to country-specific fundamentals. Related evidence appears in corporate credit: Collin-Dufresne, Goldstein, and Martin (2001) identify a large common component in credit-spread changes, while Gebhardt, Hvidkjaer, and Swaminathan (2005) show that systematic bond-market exposures help organize corporate-bond returns. Our contribution is to document an analogous structure for state municipal portfolios: a near-unit exposure to the aggregate municipal market dominates time-series return variation, whereas more heterogeneous carry, value, and low-risk exposures primarily organize relative returns across states.

Our selection of factors is based on models from the finance literature which have been used to price the risk premia of government and corporate bonds. Examples include Israel, Palhares, and Richardson (2018), Bektic, Hachenberg, and Schiereck (2020), Gava, Lefebvre, and Turc (2020), and Dang, Hollstein, and Prokopczuk (2023). The only work we are aware of applying such factor models to municipal bonds are Sibears (2017) and Wang (2022). Sibears (2017) constructs the factors from 6 indices related to bond characteristics. Our analysis, by contrast, uses sorted portfolios of state-specific indices to enable direct

comparison with factor models for other asset classes. Wang (2022) uses the Fama-MacBeth approach with a 1-factor model from panel data on individual issues to examine the link between current yields and future returns. Our paper instead focuses on decompositions of risk both in normal market conditions and during the financial turmoil of COVID.

By introducing a factor model for municipal bonds, this paper contributes to the overall literature on municipal bond pricing. One classic strand of this literature studies idiosyncratic risk and focuses on the effect of local economic and policy environments on municipal bond yields; for example, Painter (2020) examines the effect of climate change exposure on munis. Yang (2019) and Gao, Lee, and Murphy (2019) consider the effects of state bankruptcy policies on municipal bonds, while Marks and Raman (1987) and Benson and Marks (2007) evaluate fiscal performance more broadly. Broadus and Cook (1988) provides an analysis of the effects of a large set of bond and issuer characteristics.

Another strand of the literature seeks to understand bond-specific financial characteristics, particularly the effect of credit quality on individual bonds. Schwert (2017) finds that ratings influence spreads more than liquidity characteristics, while Ang, Bhansali, and Xing (2014) find the opposite. Adelino, Cunha, and Ferreira (2017), Cornaggia, Cornaggia, and Israelson (2018), and Benson and Marks (2022) exploit an exogenous rating recalibration to identify the effect of ratings on yields. Other papers that examine credit quality include Capecci (1991), Rivers and Yates (1997), Chalmers (1998), Peng, Kriz, and Wang (2014), and Cornaggia, Hund, and Nguyen (2022).

Other papers study the effects of market structure, opaqueness, and transaction costs in the muni market; Harris and Piwowar (2006), Green, Hollifield, and Schürhoff (2007), and Griffin, Hirschey, and Kruger (2023) are key examples. Additionally, Hildreth and Zorn (2005) and Griffin, Hirschey, and Kruger (2025) provide helpful surveys highlighting structural issues in muni markets.

Due to the tax-free nature of municipal bonds, much of the literature has focused on the effect of income tax exemptions, tax clientele effects and market segmentation. Some papers that have explored this channel include Livingston (1982), Kidwell, Koch, and Stock

(1984), Poterba (1989), Lovely and Wasylenko (1992), Longstaff (2011), and Fulkerson and Haskell (2021). In this paper we consider the tax sensitivity of municipal bond returns before and after the Tax Cuts and Jobs Act (TCJA). Garrett et al. (2023) look at the overall effect of TCJA on muni yields. Fleckenstein and Longstaff (2025) find a tax-related “convenience premium” in municipal bond price and measure its change due to the limitation of SALT deduction under TCJA. Pevzner et al. (2025) show that county-level muni yields became less sensitive to local property taxes after the SALT caps were introduced. Our results are in general agreement in this literature. Our estimates present suggestive indirect evidence of tax-related market segmentation, a convenience premium and the influence of local tax rates on pricing.

In studying the initial phase of COVID’s effect on municipal bond yields, we interact with recent papers studying the early days of the COVID crisis and the movements of muni yields. Bi et al. (2020) and Li, O’Hara, and Zhou (2024) investigate financial characteristics that influenced municipal bond responses in early 2020. Multiple papers have studied the effects of the Federal Reserve’s Municipal Liquidity Facility (MLF) on the municipal bond market, finding that MLF increased liquidity, decreased yields, alleviated credit risk, and encouraged debt issuance; these papers include Haughwout, Hyman, and Shachar (2026), Bi and Marsh (2022), Fritsch, Bagley, and Nee (2021), Bordo and Duca (2023), and Li and Lu (2020). A number of recent papers investigate the relationship between local COVID situations and municipal bond yields, including Odusami and Mansur (2022), Tran and Uzmanoglu (2022), Li and Lu (2020). Cusatis and Hoxha (2022), Wu and Ostroy (2021), and Wu and Ostroy (2022) analyze COVID-related national market disruptions. Compared to these strands of COVID-period literature, our approach offers a complementary and broader perspective by examining how types of risk changed from normal market conditions to the COVID episode.

Finally, the paper complements the wider literature on the relationship between pandemic risk and asset prices. Baker et al. (2020) support this paper’s finding that COVID was a unique event: they show that the stock market volatility in response to COVID far exceeded the stock market reaction to any other infectious disease outbreak over the previous

120 years. Another recent paper by Park, Kang, and Eom (2024) estimate significant risk premia for COVID factors constructed from firm-level and sector-level exposures to pandemic risk. Compared to studies of stock returns, the advantage of our research design is in exploiting the variation in pandemic dynamics across states. The surprising finding for our analysis of the COVID period is that own-state economic conditions do not have much effect on municipal debt prices whereas the cross-state distribution of pandemic intensity emerges as a relevant state variable.

2 A Factor Model of Municipal Bond Prices Applied to the Pre-COVID Period

The linear factor model has a long history in finance literature and is rooted in utility maximization under the assumption that the stochastic discount factor takes a linear form. Our selection of factors draws on insights from both the corporate and municipal bond literatures. In choosing the factors we try to balance model parsimony with its ability to capture essential categories of risks identified and studied in prior research. The national municipal bond market factor plays the same role in the model as the market portfolio in CAPM-type models. Four “style” factors—value, carry, momentum, and low-risk—are adopted from the corporate bond pricing models and are thought to capture broad asset pricing regularities. We provide a more detailed interpretation of these factors in the context of municipal bond pricing in Section 2.2.2. Finally, we control for possible interactions between the Treasury and the municipal bond markets by adding the slope and convexity factors describing the Treasury yield curve.

2.1 Fama-MacBeth Two-Step Estimation

The model relates the rate of return on an index of municipal bonds in state i on day t to N aggregate factors, defined by the row vector $\mathbf{F}_t = [F_t^1, \dots, F_t^N]$. Our estimation is based

on the familiar two-step procedure introduced in Fama and MacBeth (1973). For each state i municipal bond index we run the time series regression

$$r_{it}^e = \gamma_i + \mathbf{F}_t \beta_i + u_{it}, \quad (1)$$

where $r_{it}^e = r_{it} - r_t^F$ is the excess return of state i 's return r_{it} over the risk-free return in the economy r_t^F .¹ The column vector $\beta_i = [\beta_i^1, \dots, \beta_i^N]^\top$ contains the exposures of each of the aggregate factors on excess returns in state i . Each estimated $\hat{\beta}_i^n$ corresponds to the estimated exposure of muni index i to factor n . Each vector $[\hat{\gamma}_i; \hat{\beta}_i]$ captures the estimated pricing model for excess returns in a state.

After estimating the individual state models, we perform a second step regression to recover the aggregate risk premia corresponding to each of the factors. We do this by running a series of cross-sectional regressions on each day t :

$$r_{it}^e = \lambda_t \hat{\beta}_i + \alpha_{it}, \quad (2)$$

where $\lambda_t = [\lambda_t^1, \dots, \lambda_t^N]$ is the row vector of time- t risk premia for each of the factors. We can then estimate the average risk premium for factor n along with its variance, given by

$$\bar{\lambda}^n = \frac{1}{T} \sum_t \hat{\lambda}_t^n; \quad (3)$$

$$\sigma^2(\bar{\lambda}^n) = \frac{1}{T^2} \sum_t (\hat{\lambda}_t^n - \bar{\lambda}^n)^2. \quad (4)$$

We further apply the Shanken correction to the variances to account for the fact that the β s are estimated regressors.

¹The rate of return can be either pre-tax or after-tax. See Section 2.3.1 for an explanation of the tax adjustment.

2.2 Risk Factors

2.2.1 Data and Samples

Our municipal bond data are at the state-day level. For each state, we use the total return index of municipal bonds available from S&P Global.² These indices are constructed to represent a group of municipal bonds from an array of government entities, including state governments, local governments, and other entities such as utilities boards.³ Data on daily treasury yields are from the Federal Reserve Bank of St. Louis. Daily risk-free returns are computed from the total return index of 1-3 month Treasury bills, also obtained from the S&P 500 fixed-income indices.

Start dates for the municipal bond indices vary across states; we begin our sample in January 2014 so that all states have equal numbers of observations in their pricing models. Our pre-pandemic model is estimated through the end of 2019.

We consider two broad classes of factors to be included in the vector $\mathbf{F}_t$. In this section, all factors are conventional, i.e., factors that have been used in other pricing models. Section 3.2.4 adds unconventional factors to capture the possible regime change during the COVID pandemic.

2.2.2 Style Factors

Style factors are based on four potential determinants of excess returns familiar from asset pricing literature and thought to capture asset pricing regularities. These factors take a characteristic of an asset, sort the assets on that characteristic, and compute the hypothetical

²We download the main series directly from S&P, while the extended samples are obtained from Bloomberg.

³More specifically, S&P constructs these indices as strict subsets of its overall municipal bond index. The index is largely a rules-based index, including all securities meeting the inclusion criteria. The subindices are simply those bonds from the overall index having the characteristic defined by the subset. For example, the index for Michigan includes all bonds from the overall index that were issued by Michigan state and local governments. For more information, see the methodology at <https://www.spglobal.com/spdji/en/documents/methodologies/methodology-sp-fixed-income-index-math.pdf>.

return of a zero net wealth portfolio that is long on the top quintile of the distribution and short on the bottom quintile.⁴ The return of this portfolio is intended to capture the influence of the characteristic on average excess returns, and that return is what we call a factor. The risk premium associated with the factor estimated in (2) indicates whether the factor plays a significant role in pricing the asset.

We consider four such portfolio-based factors constructed by sorting on style characteristics. The *carry* factor sorts on yield-to-worst, or lowest possible yield from buying and holding a callable bond.⁵ The *value* factor sorts on real yield-to-maturity, or YTM minus the 10-year inflation expectation.⁶ The *momentum* factor sorts on the trailing six-month return. Finally, *low-risk* is sorted on the lowest to highest volatility of the last 18 months of returns, and then weights the returns by the inverse of volatility.

Figure 1 illustrates the performance of style factors over the period 2012-2021. With the exception of three muni selloff events that correspond to “taper tantrum” in 2013, post-2016 election tax policy fears and the onset of the COVID pandemic, carry and value factors generally earn positive returns. Returns on the low risk factor rise during the selloff events and fall as the muni market recovers.

2.2.3 Municipal Bond and Treasury Market Factors

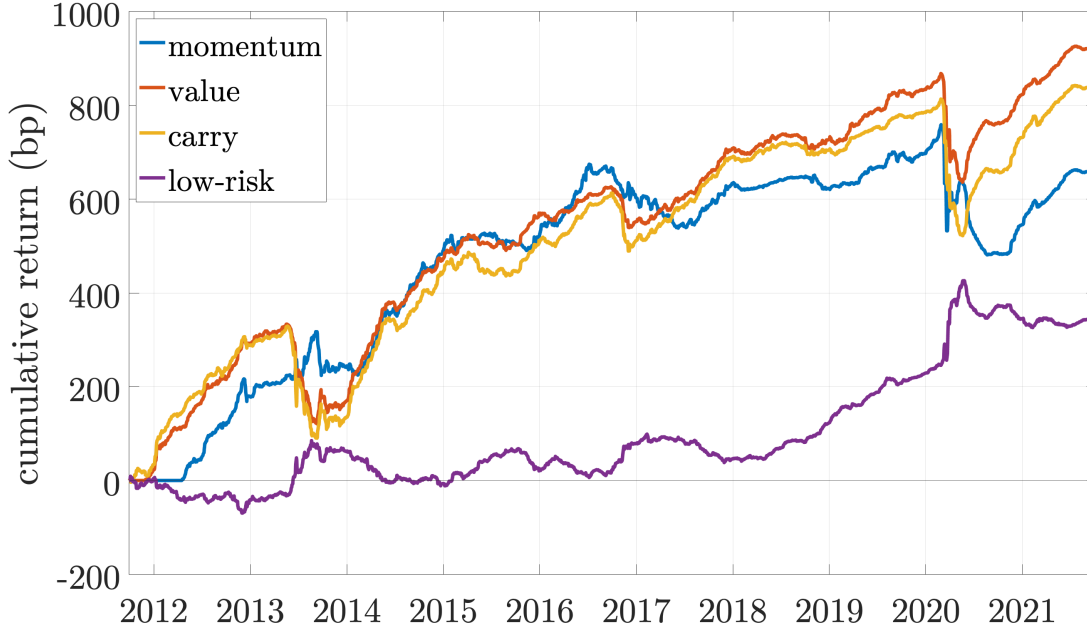
Following standard practice in the asset pricing literature, our model includes a market factor. Specifically, we include a municipal bond market factor, which captures the component of returns that is common to all municipal bonds. We additionally include two Treasury market factors aimed at capturing the role of macroeconomic and financial risks, as reflected in the term structure of the Treasury yield curve. The first factor is the *slope* of the yield curve,

⁴While there is no existing vehicle to short these S&P indices, the returns on these indices tend to be highly correlated with their analogous state-level ETFs; these ETFs, in contrast, can be easily shorted. We prefer to use the S&P indices for our analysis due to their uniformity across states, allowing for easy comparison.

⁵Specifically, according to the S&P methodology, the yield-to-worst is computed by generating a yield for each potential redemption date, given the bond’s potential stated calls prior to maturity, and then taking the lowest out of all these yields. The yield-to-worst data we use are provided directly by S&P.

⁶This expectation is the median 10-year ahead inflation from the Survey of Professional Forecasters.

Figure 1: Cumulative Return on Style Factors, 2012-2021



Note: Lines indicate the historical performance of long-short portfolios based on the indicated style factors for state municipal bond indices from S&P. A portfolio for each factor is long on the top quintile of the distribution and short on the bottom quintile.

represented by the spread between the 3-year and 3-month treasury rates. The second factor is *convexity*, or the spread between the 5-year and the average of the 2-year and 10-year Treasury rates.

2.3 The Factor Model in 2014-2019

This section presents the factor model estimates over the pre-COVID sample. Section 2.3.1 describes the tax adjustment that puts all factors on an after-tax basis. Section 2.3.2 reports the baseline risk premia and decomposes predicted returns into persistent, aggregate, and relative components, showing that the national market factor dominates time-series variation while heterogeneous style exposures drive returns across states. Section 2.3.3 turns to taxation, relating the weak pricing power of Treasury factors to market segmentation and using the model's pricing errors to detect a convenience premium before and after the TCJA.

2.3.1 Tax Adjustment in the Factor Model

We formulate the baseline model in terms of after-tax returns. Income from U.S. Treasury securities is taxed at the ordinary income rate at the federal level and is generally exempt from state and local taxes. Therefore, the tax adjustment for Treasury-derived returns is straightforward: multiply every Treasury-derived return used in construction of r_{it}^e and $\mathbf{F}_t$ by the federal after-tax coefficient, $1 - \tau_{F,t}$, in year t . Returns on municipal bonds are exempt from federal taxes and do not require adjustment for federal tax.

The situation with state taxes is more nuanced. Municipal bond income is exempt from home state taxes in almost every state, though not universally. As a result, our muni-derived factors and outcome variables are effectively equal to after-tax returns, so we make no further tax adjustment.

The implicit assumption we are making with this adjustment is that the factor model describes the behavior of a representative national investor consisting of tax residents of all states but operating across segmented markets. The investor’s portfolio consists solely of home state municipal bonds, which confer a unique tax advantage and reinforce market segmentation.⁷ Under this assumption, all factors correspond to after-tax returns.

2.3.2 Results from the Baseline Model

Table 1 reports time series average risk premia $\bar{\lambda}^n$ (unadjusted for exposure) measured in annual basis points. Appendix Table A1 column 1 shows the corresponding exposure-adjusted risk premia $\bar{\beta}\bar{\lambda}$ by factor group.

Dominance of market factor and aggregate risk premium. One of the main results concerns the similarities between the states’ bond portfolios and the national index. The market factor has the largest contribution to predicted returns (Table A1, column 1). States have large and fairly similar exposures to the market factor: exposures (market betas,

⁷Pirinsky and Wang (2011), Schultz (2012), and Babina et al. (2021) note market segmentation by state.

not reported in the table) to the market factor across states are in a fairly tight range $[0.8, 1.1]$ with the average of 0.95. In other words, state municipal bond portfolios behave somewhat similarly to levered or delevered versions of the national municipal bond index. In fact, 97.5 percent of the variance in the model’s predicted returns are due to the time-varying aggregate risk premia that capture market-wide shocks that hit all states simultaneously. The last result follows from the fact that market betas have high mean and low variance.⁸

Table 1: Baseline Model and its Variants: Average Risk Premia ($\bar{\lambda}^n$, annual basis points)

Factor	Baseline	Drop Treasury	Drop bond market	Style factors only	Symmetric tax
Style Factors					
Momentum	161** (0.030)	173** (0.018)	152** (0.022)	136** (0.037)	152* (0.052)
Value	180*** (0.000)	180*** (0.000)	180*** (0.000)	184*** (0.000)	179*** (0.000)
Carry	163*** (0.000)	168*** (0.000)	160*** (0.000)	156*** (0.000)	154*** (0.000)
Low-risk	-41 (0.170)	-27 (0.353)	-49* (0.085)	-50* (0.072)	-65** (0.046)
Bond market factor	583*** (0.000)	576*** (0.000)	-	-	526*** (0.001)
Treasury Market Factors					
Slope	2050 (0.224)	-	2622 (0.191)	-	7542** (0.037)
Convexity	434 (0.101)	-	525* (0.089)	-	1016** (0.034)
Avg. Adj. R^2	0.29	0.26	0.26	0.22	0.29

Note: Values displayed are the average risk premia, in annualized basis points, for a given factor (row) in a given model (column), where models include different combinations of the factors. Average risk premia for the factors are computed as described in Equation 3 of the text. Data are daily state municipal bond indices from S&P from 2014 to 2019. p -values for these estimates are in parentheses. *: $p \leq 0.1$; **: $p \leq 0.05$; ***: $p \leq 0.01$.

Cross-sectional return variance. Despite the similarities in municipal bond portfolios across states that the model detects through the market factor, there are important cross-state differences. Other states’ returns are, in fact, useful basis assets for constructing municipal risk-mimicking portfolios. This is reflected in significant risk premia on the carry, value and momentum factors. In contrast to the market factor, states’ exposures to these factors are an order of magnitude lower but have much higher variance. Style betas are in a broad range $[-1.5, 1.5]$, and average exposures are between 0.02 and 0.12 in absolute value. Treasury betas are smaller still, around 0.001, and the risk premia on Treasury factors are insignificant, even though Treasury factors contribute 12 percent to the cross-sectional

⁸In a single-factor model with just the market factor M , the ratio of predicted return variances due to aggregate versus relative risk premia would be equal to $\frac{(\mathbb{E}_i(\beta_i^M))^2}{\text{Var}_i(\beta_i^M)}$.

return variance (Table A1, columns 5-6).

Relatedly, the risk premium in the cross-section is overwhelmingly explained by time-varying relative returns, that is, by differential responses in more/less exposed states to movements in the aggregate risk premia. Further decomposing the cross-sectional variance by factor group reveals that relative risk premia associated with the market factor, style factors and Treasury factors account for 33, 55 and 12 percent of the cross-sectional return variance respectively (Table A1, lines 2-4, columns 5-6), leaving just over 1 percent of cross sectional variance accounted for by persistent risk premia (Table A1, line 5, column 5).

Small role of persistent risk premia. Since persistent risk premia explain almost none of the cross-sectional return variance, there are almost no persistent differences in realized holding returns. This is despite the persistent yield spreads across states. This would be the case if the high-yield states experienced more dramatic adverse shocks to prices of their bonds and low-yield states experienced more moderate repricing. Generally, such an outcome is consistent with no-arbitrage pricing.

Interpretation of style factors. In an integrated market, an asset’s rate of return should match that of a similar asset. If states with similar portfolio compositions share return behavior, the model can pick up these commonalities through the state-sorted factors. The carry, value and low risk sorted portfolios plausibly serve as “asset-matching” tools and proxy for unobservable state portfolio characteristics. In other words, the carry, value and low risk sorted portfolios may add information about a state’s portfolio absent in its own index by statistically linking groups of states with similar and dissimilar municipal bond indices. It is harder to interpret the momentum factor this way: why would a state’s municipal bond returns depend on recent high and low performers in other states? The importance of momentum may just capture trading frictions or similar inertial behaviors of transactions, funds flows or information.

Style factors illustrate the advantage of model parsimony. Detailed, time-consistent information on each state portfolio’s composition may be difficult to obtain. Instead, portfolios sorted on observable characteristics correlated with these variables can serve as useful

basis assets.

Robustness check. Because the pricing implications of style factors are more subtle than those of the market factor, Columns 2-4 of Table 1 include robustness checks by removing other factor groups from the model. The results indicate that risk premia on style factors remain significant and not much affected regardless of which other factor group is present.

2.3.3 Tax Analysis

Evidence of market segmentation. In Table 1 column 1, Treasury risk premia are insignificant. In addition, exposures to Treasury factors are 3 orders of magnitude smaller than those on the market factor. Thus, as explained above, Treasury term structure does not seem to add useful information for pricing municipal bonds. This result is consistent with market segmentation between municipal bonds and Treasuries. We investigate whether this outcome is related to the differential federal tax treatment of the two assets. Under the existing tax regime, the federal tax treatment of municipal bonds and Treasuries is asymmetric: municipal bond income is exempt from federal tax while Treasury income is not.

In Table 1 column 5, we re-estimate the model in a symmetric tax regime that mechanically imposes federal tax on municipal bond income.⁹ As a result, risk premia on Treasury factors become significant, and the risk premium on the market factors declines slightly.

Under symmetric tax, the betas on Treasury factors are still 2-3 orders of magnitude smaller than those on the market factor. The low exposures rationalize the estimated risk premia on those factors that appear enormous (e.g. 7500 annual basis points in Table 1 column 5). In fact, the product of the two, $\bar{\beta}^n \bar{\lambda}^n$, that measures the contribution of factor n to fitted return is -1 and 6 annual basis points for the slope and convexity factors,

⁹Since the after-tax coefficient, $1 - \tau_{F,t}$, is applied to every return in r_{it}^e and $\mathbf{F}_t$, the coefficient cancels out and leaves the specification equivalent to a tax-unadjusted model.

respectively.

The comparisons of Treasury and municipal bond market risk premia between the baseline and symmetric tax models indeed suggests that market segmentation in the current regime might have to do with asymmetric tax treatment. However, these results should not be interpreted as a policy experiment measuring the market effects of the federal income tax exemption repeal, because the actual policy change would cause movements in returns and repricing not captured by data in our sample.

Market segmentation may also provide some clues as to why style factors are relevant for pricing across different asset classes. Theoretically, shared drivers of risk premia could arise from financial arbitrage, i.e. municipal bonds being good substitutes for corporate bonds in terms of asset characteristics. This force would be most relevant in integrated markets. However, in segmented markets moved primarily by aggregate risk premia, common, national level shocks become a more plausible reason for cross-asset pricing factors.

Indirect effect of state taxes on municipal bond returns. The tax-adjusted model already incorporates the effects of state taxes on municipal bond risk premia through several muni-derived factors. However, municipal bonds also compete with taxable investments such as corporate bonds, whose after-tax returns depend on both federal and state tax rates. As a result, there may be indirect effects of state taxes on municipal bond returns due to arbitrage across asset classes. Because our factor model does not include returns on taxable investments, these indirect effects may be partially reflected in the model’s pricing errors.

To test for this, we regress the pricing errors from (2), α_{it} , on the deviation of the state’s current top marginal income tax rate from the national average, $\bar{\tau}_{it} = \tau_{it} - \tau_t$. We also analyze the provision of the Tax Cuts and Jobs Act (TCJA) that capped the state and local tax (SALT) deduction using an empirical specification

$$\alpha_{it} = \delta + \theta^{pre} \bar{\tau}_{it} S_t + \theta^{post} \bar{\tau}_{it} (1 - S_t) + v_{it}, \quad (5)$$

where S_t is the dummy variable for the presence of the SALT deduction at date t . In the above equation, the estimate of δ corresponds to the average (over time and across states) pricing error in (2) and θ registers the change in municipal bond return for each percentage point deviation of the state tax rate from the national average, expressed in annual basis points. A tax-related convenience premium would imply a negative value of θ , that is lower municipal bond returns in states with lower after-tax returns on taxable investments.

Our empirical strategy can be thought of as a detection rather than measurement tool for the convenience premium. By its nature, the convenience premium is a general equilibrium phenomenon, and as such it should be already partially priced into factors. Since we cannot isolate the effect of state tax dispersion on the factors themselves, we cannot reliably detect whether the convenience premium became larger or smaller over time based on θ coefficients alone. However, we can detect the general equilibrium effect itself if θ moves opposite to the direction implied by the tax incentive.

The capped SALT deduction post-TCJA effectively *increased* the variation in after-tax returns on taxable investments across states. Accordingly, in the absence of post-TCJA factor repricing, one would expect amplified cross-state differences in the relative appeal of home-state municipal bonds to translate into a more pronounced convenience premium post-TCJA, that is, $|\theta^{post}| > |\theta^{pre}|$. In line with this, Fleckenstein and Longstaff (2025) find that the convenience premium indeed increased.

Table 2 presents estimates by sub-period pre- and post-Tax Cuts and Jobs Act. The convenience premium is present and significant in both sub-periods but we cannot assess its total magnitude, because it is spread between factors and pricing errors. Table 2 shows $|\theta^{post}| < |\theta^{pre}|$. At first glance, one would conclude that convenience premium shrunk post-TCJA, contrary to incentives. The smaller post-TCJA θ coefficient could reflect a general equilibrium effect where tax changes are incorporated into factors. However, this regression does not measure the general equilibrium effect and cannot establish how the total convenience premium has changed.

Table 2 methodology and results are consistent with Pevzner et al. (2025) who

Table 2: Evidence of Convenience Premium Pre- and Post-TCJA

Coefficient	Estimate
θ^{pre}	-5.030*** (0.000)
θ^{post}	-2.027*** (0.006)
δ	1.618*** (0.000)
R^2	0.0004
N obs.	78,100
Wald test $\theta^{pre} = \theta^{post}$: $\chi^2(1) = 4.965$ $p = 0.0259$	

Note: Values displayed are the estimated coefficients from Equation 5 of the text, where the dependent variable is the pricing error from Equation 2 in the baseline model and $\bar{\tau}_{it}$ is the deviation of the state's top marginal income tax rate from the national average. Coefficients are expressed in annual basis points per percentage point of tax-rate deviation. p -values for these estimates are in parentheses. *: $p \leq 0.1$; **: $p \leq 0.05$; ***: $p \leq 0.01$.

estimate the analog of our θ and find that tax sensitivity of municipal bond yields fell after the SALT deduction was capped. The methodology differs from Fleckenstein and Longstaff (2025) who measure the convenience premium more directly by adjusting for credit quality and liquidity first.

3 The Factor Model in the COVID-19 Period

In this section we illustrate regime dependence that may introduce risk transmission mechanisms with heterogeneous impacts across states. We expect shocks of this nature to affect both aggregate and relative risk premia. In addition, a different risk transmission mechanism might temporarily emerge and be reflected in new state variables that conventional factors do not capture.

3.1 COVID-19: A Large Aggregate Shock with Differential Impact across States

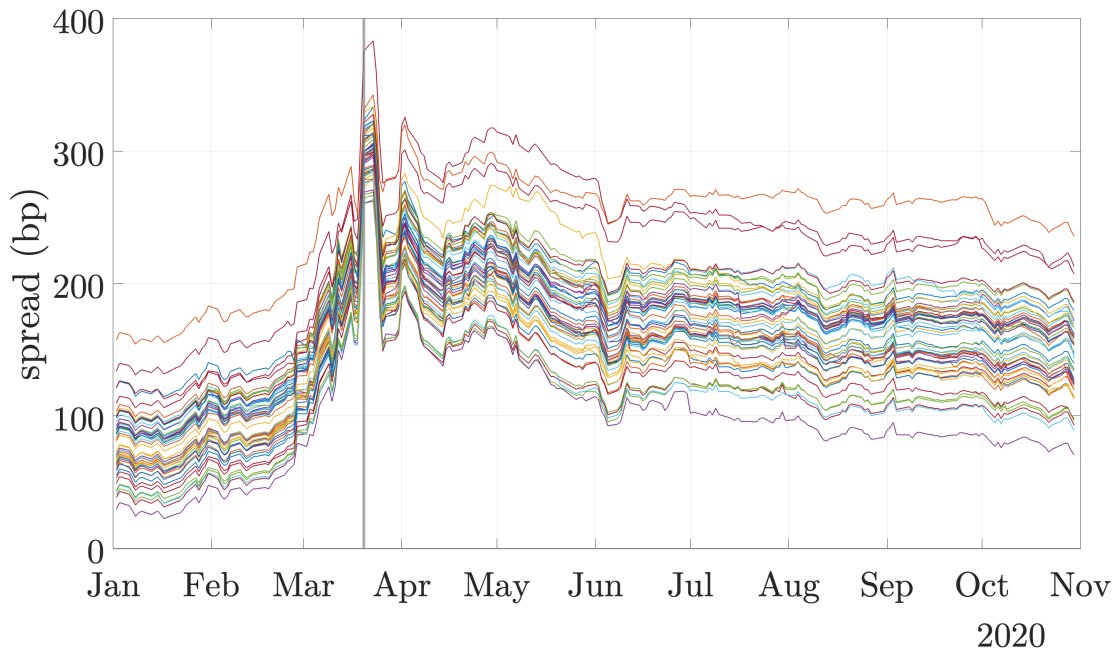
The onset of the coronavirus pandemic in early 2020 created a disruption in bond markets due to a “dash for cash,” and even traditionally safe assets – including U.S. Treasuries – experienced a rapid selloff.¹⁰ The municipal bond market also experienced this severe selloff. On March 6, 2020, the S&P Municipal Bond index yielded a rate of 2.33%; by March 20, this rate had jumped 162 basis points to 3.95%. Liquidity concerns were rampant, and uncertainty around the course of the virus and the future of municipal finances drove investors to cash safety. Municipal bond funds, owning almost \$900 billion in munis at the end of 2019, raced to sell and the primary market for new municipal debt ground to a standstill.

The distress continued even though the Federal Reserve took swift action to calm markets. On March 18 it announced a lending facility to backstop money market mutual funds, and on March 20 it announced that it would accept certain low-risk municipal bonds as collateral for these loans. On March 23, it expanded the range of municipal debt it would accept. On April 9, the Fed unveiled the Municipal Liquidity Facility (MLF) to purchase debt directly from U.S. states and large local governments; the program expanded twice over the following two months to include smaller government borrowers. Congress also stepped in during this time, providing \$340 billion in direct aid to state and local governments through the CARES act, and by providing general economic relief through other provisions such as the Paycheck Protection Program.

To illustrate the magnitude and duration of the disruption to the municipal bond market, Figure 2 plots the state-level S&P municipal index yield spreads with the 10-year U.S. Treasury rate. Spreads began to spike in late February 2020, continuing their upward trajectory until the peak on March 20 (vertical line), the first day of signaling from the Fed that it was willing to support the muni market. While spreads came down from their peak in

¹⁰Barone et al. (2023) documents that selling pressures on sovereign bonds were greater in March 2020 than during the global financial crisis of 2007-2008, and pressure on U.S. Treasuries were greater than for other major economies. Haddad, Moreira, and Muir (2021) show that investment-grade bonds moved more closely with high-yield bonds and equities than in previous crises, indicating an atypical event.

Figure 2: Impact of COVID-19 on State-Level Municipal Bond Spreads



Note: Each line represents the evolution of the spread between the constant-maturity yield on a 10-year Treasury and the yield-to-maturity (YTM) for an individual U.S. state's S&P Municipal Bond index during Jan 1-Oct 30, 2020. These indices include both state and local municipal bonds. The vertical line indicates the peak of YTM on the aggregate municipal bond index on March 20, 2020.

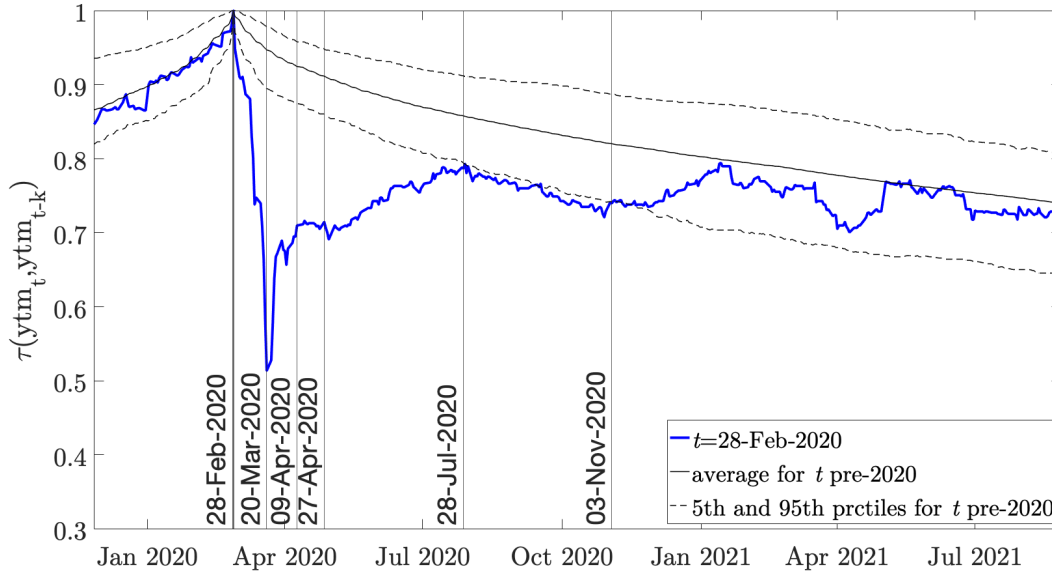
March, they remained volatile through May; volatility then decreased and spreads stabilized throughout the rest of the year, albeit at higher levels than before the pandemic.

Most interesting for our purposes is the differential impact of the pandemic on municipal risk premia across states. We observe a marked shift in *relative* bond yields: states did not remain in the same segment of the cross-sectional distribution of yields throughout this period. To be more precise about the shift in the ranking of bond yields, Figure 3 shows the evolution of the Kendall rank correlation coefficient, the blue line in the figure, commonly referred to as Kendall's τ coefficient.¹¹ Kendall's $\tau \in [-1, 1]$ measures the rank correlation of yields between the current date and a reference date. If the ranking is identical to the reference period, $\tau = 1$, and if the ranking is perfectly reversed from the reference period, $\tau = -1$.

¹¹ Kendall's τ is a measure of the agreement between the relative rankings of elements in a pair of vectors. If x and y are two vectors of size n , Kendall's τ is defined as $\tau(x, y) = \frac{2}{n(n-1)} \sum_{i < j} \text{sgn}(x_i - x_j) \text{sgn}(y_i - y_j)$.

The blue line on Figure 3 plots this τ measure in 2020 and 2021, and summarizes the relationship between the rankings of state-level municipal bond yields on a given date with the ranking as it stood on the reference date of February 28, 2020. On the reference date, τ is defined to be 1. As the time interval k widens, rank orderings shift and τ begins to fall. The solid black line shows the historical average τ for a particular k based on data from 2014 to 2019, and the dotted black lines show the 5th and 95th percentiles of the distribution of τ as k increases.¹²

Figure 3: Rank Correlation of Municipal Yields in 2020



Note: The blue line represents the Kendall's τ rank correlation measure for state municipal bond yields between a given date and the reference date, February 28, 2020. τ is equal to 1 if the ranking of states by yield is exactly the same as it was on the reference date, and equal to -1 if the ranking is perfectly opposite. The solid black line indicates the average evolution of τ from any given reference period, estimated on data from 2014 to 2019. The dashed black lines indicate the 5th and 95th percentiles. State yields are the YTM of the S&P state municipal bond indices. Kendall's τ is defined in Footnote 11. Vertical lines indicate significant dates in government policy related to the muni market in 2020.

Before Feb 28, 2020, the rank correlation pattern is not far from its pre-pandemic average – the blue line closely follows the solid black line. Right after the COVID news, τ

¹²For each date s , let $T_s(k) = \tau(y_s, y_{s+k})$ be the Kendall rank correlation function between YTM vectors y_s and y_{s+k} , $k = 0, \pm 1, \pm 2, \dots$, and let $\bar{\tau}_k = (1/N) \sum_n T_n(k)$ be its average over time. By construction, $T_s(0) = \bar{\tau}_0 = 1$, all s . The blue line on Figure 3 shows $T_t(k)$ with t fixed at Feb 28, 2020. The black line superimposes $\bar{\tau}_k$ over the pre-COVID time period 2014-2019, centered around the same reference date Feb 28, 2020. The dashed lines show the 5th and the 95th percentiles among the values of $T_s(k)$ during 2014-2019.

plummets to 0.5, well outside the 90 percent error band based on pre-2020 data. This indicates a sudden, substantial re-shuffling of yield rankings far outside the typical distribution. This anomaly persists until late July 2020; while τ gradually rises over the course of the summer of 2020, it remains outside the 90 percent interval until November 2020, when the ranking finally approaches what one would have predicted based on pre-COVID data. The figure also notes significant policy dates affecting municipal bonds.¹³

The evidence above suggests that risk mimicking portfolios sorted on yields may have become less informative about municipal bond returns. To illustrate this, we estimate the factor model during the pandemic period. We also show that regime dependence can be captured by new factors that depend on an additional state variable – the distribution of pandemic outcomes across states.

3.2 The Factor Model during the COVID Period

This section proceeds in four steps. Section 3.2.1 estimates the baseline model on the early pandemic subsample and shows that most conventional factors become uninformative, while the apparent pricing power of the Treasury slope factor points to a missing state variable. Section 3.2.2 introduces COVID factors constructed from long–short portfolios sorted on the cross-state distribution of pandemic outcomes, and Section 3.2.3 compares their dynamics with those of the style factors over the three phases of 2020 sample period. Section 3.2.4 re-estimates the model with COVID factors and contrasts the resulting composition of risk premia with the pre-COVID structure of Section 2.3.

¹³On March 20, the Federal Reserve added short-term municipal securities to its Money Market Mutual Fund Liquidity Facility, easing acute stress in tax-exempt money funds. On April 9, it announced the Municipal Liquidity Facility, a \$500 billion backstop for lending directly to state and local governments. On April 27, it broadened eligibility to smaller cities and counties and lengthened maturities. On July 28, it extended its emergency facilities, including the MMLF, through year-end. Finally, the November 3 election shaped expectations for federal aid and tax policy.

3.2.1 Baseline Model

The baseline factor structure largely breaks down during the early pandemic period. Table 3, column 1 reports risk premia for the baseline model on the March 1-May 15, 2020 subsample (Phase 1). These estimates are driven by the market crash of late March 2020 (Figure 2): the predicted average return plummets to negative 17.5 percent annual (Table A2, column 1, line 5), average exposure to the market factor rises to 1.07, and the cross-state variance of exposures rises as well. The risk premia on the market and low-risk factors become so volatile that their estimates lose significance, carry and value become collinear with the market factor and therefore uninformative for pricing, and momentum largely ceases to matter because of the dramatic re-ranking of yields. Because the entire market crashed, the market factor dominates both the total and cross-sectional return variance (Table A2, columns 3 and 6, line 2), and the larger variance of exposures amplifies this effect.

The one baseline factor that appears to gain pricing power is the Treasury slope. While the national bond market factor earns a sharply negative return, the low-risk and Treasury factors earn modest positive exposure-weighted returns (Table A2, column 1, lines 3–4), and average exposures to Treasury factors rise substantially relative to the pre-COVID sample. The higher exposure makes the information in the slope factor more visible: its risk premium is larger than in the pre-COVID period and significant, and Treasury factors account for 23 percent of the cross-sectional return variance (Table A2, columns 5–6). One could be tempted to read this as evidence that market participants focused on the expected policy response to the pandemic; after all, the Federal Reserve deployed a range of unconventional tools, including a large-scale asset purchase program that affected the longer end of the yield curve and thus the slope. We show in Section 3.2.4, however, that this attribution is misleading. The baseline model omits a state variable, and in its absence the cross-sectional variation associated with that variable is loaded onto the Treasury slope factor. Once the missing variable is included, the slope premium loses significance and the Treasury share of the cross-sectional variance falls sharply.

Somewhat surprisingly, the missing state variable is not the state’s own pandemic

situation. Pricing errors from the Phase 1 model are not statistically related to state-level mobility, cases per capita, ICU beds per capita, or deaths per capita: regressing the pricing errors on these four variables yields a joint F -statistic of 0.35 (p -value 0.85). Since neither the baseline factors nor own-state COVID characteristics explain municipal bond returns, we extend the model with COVID factors that capture a richer state variable, the entire distribution of the pandemic situation across states. We show that these factors both carry relevant information for pricing and clarify the nature of risks during the pandemic period.

Table 3: Factor Model During the COVID Pandemic

Factor	Baseline factors		Baseline + COVID factors		
	pre-COVID	Mar 1–May 15	Mar 1–May 15	Jun 1–Aug 15	Sep 1–Dec 15
Style Factors					
Momentum	161** (0.030)	56 (0.969)	1462* (0.094)	-753*** (0.002)	32 (0.585)
Value	180*** (0.000)	-	-	-	-
Carry	163*** (0.000)	-	-	-	-
Low-risk	-41 (0.170)	1069 (0.172)	784 (0.157)	-375** (0.034)	13 (0.882)
Bond market	583*** (0.000)	-1842 (0.799)	-1835 (0.752)	1714*** (0.004)	795** (0.043)
Treasury Market Factors					
Slope	2050 (0.224)	3221** (0.032)	571 (0.215)	165 (0.461)	-9 (0.966)
Convexity	434 (0.101)	241 (0.437)	139 (0.456)	114 (0.313)	19 (0.693)
COVID Factors					
Mobility	-	-	580 (0.254)	-147 (0.294)	-108 (0.132)
Cases / pop	-	-	903 (0.119)	-572*** (0.005)	-13 (0.786)
ICU beds / pop	-	-	991* (0.083)	-576*** (0.002)	43 (0.321)
Deaths / pop	-	-	999 (0.181)	-363* (0.069)	3 (0.957)
Avg. Adj. R^2	0.29	0.28	0.37	0.37	0.33
Variance decomposition	Table A1	Table A2	Table A3	Table A4	Table A5

Note: Values displayed are the average risk premia, in annualized basis points, for a given factor (row) in a given model and sample period (column). Average risk premia for the factors are computed as described in Equation 3 of the text. The value and carry factors are dropped from the 2020 subsamples because of collinearity with the bond market factor. Data are daily state municipal bond indices from S&P; COVID factors are from IHME. The pre-COVID column reproduces column 1 of Table 1. p -values for these estimates are in parentheses. *: $p \leq 0.1$; **: $p \leq 0.05$; ***: $p \leq 0.01$.

3.2.2 COVID Factors

The unique circumstances of the pandemic period suggest that unconventional, pandemic-related factors may have influenced municipal bond pricing. For instance, states experiencing severe COVID-19 outbreaks likely saw increased government spending and sharp declines in economic activity, both affecting tax revenues and bond prices. To capture these effects, we introduce four COVID-related characteristics for each state and construct long-short sorted portfolios based on these characteristics, going long on states with favorable pandemic conditions and short on states with unfavorable ones. This parallels style sorted portfolio factor construction and allows the model to capture the COVID impact on the risk premia. COVID factors capture a potential new state variable that baseline factors may miss.

The four pandemic-related variables are *mobility*, *cases*, *deaths*, and *ICU*. The mobility variable is an index of mobility within a state that captures the extent of shutdowns and voluntary social distancing. Cases and deaths record the number of estimated COVID cases and deaths at the state level. ICU records occupied ICU beds per capita in a state. We define COVID *characteristics* by this set of variables with cases, deaths, and ICU in per capita terms, and we define COVID *factors* as the return of a long-short portfolio sorted on each variable. The COVID characteristics and population at the state level are estimates from the Institute for Health Metrics and Evaluation (IHME).¹⁴

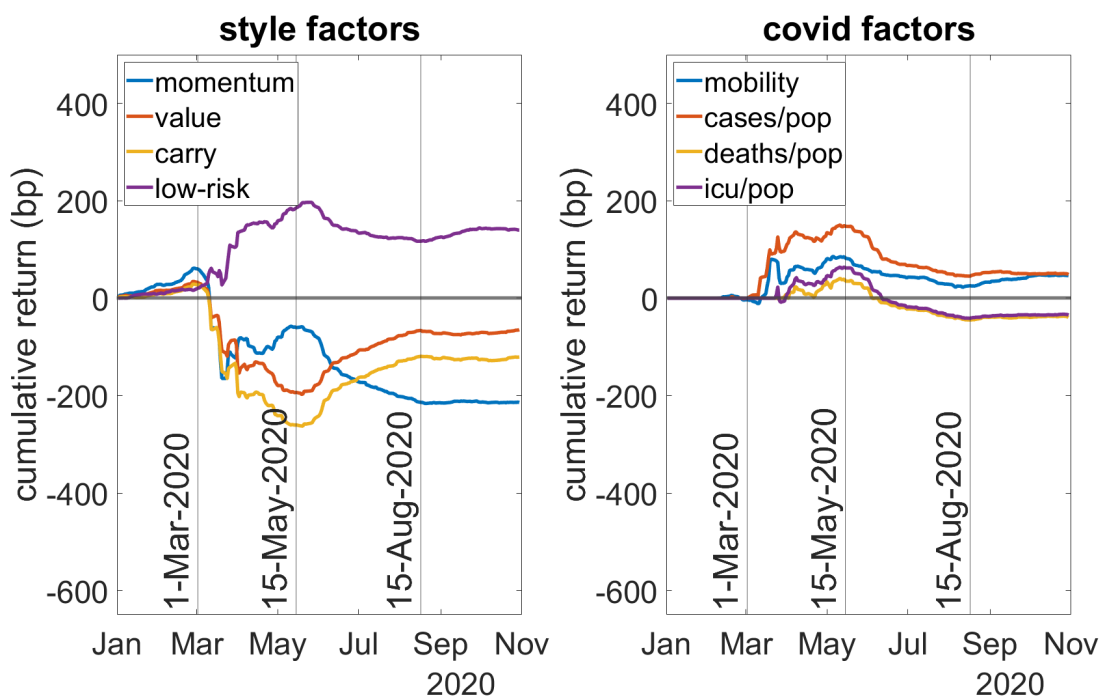
Our pandemic sample begins on March 1, 2020, in step with data availability for the COVID characteristics. The pandemic sample is split into three subsamples, analyzed separately: the early period March-May 2020 (Phase 1) and the later periods June-August 2020 (Phase 2) and September-December 2020 (Phase 3).

¹⁴IHME estimates were updated regularly through December 16, 2022. The data in this paper reflect the vintage from December 2020, which would not contain later revisions and presumably more accurately reflect the information available to investors in 2020.

3.2.3 Comparative Dynamics of Style and COVID Factors

The impact of the COVID pandemic on the municipal bond market unfolded in three distinct phases. During the early pandemic period, aggregate yields spiked and relative yields reshuffled. This substantial reordering of yields persisted until at least late summer 2020, as shown in Figure 3. Figure 4 presents cumulative returns by factor; the slope of each line at a given date indicates the contemporaneous holding return. During the early period, high-yield states began earning lower returns, as reflected in the persistent losses on the carry and value factors. Low-volatility assets were advantaged, as indicated by their positive returns, especially relative to the value and carry factors.

Figure 4: Cumulative Returns on Style and COVID-Based Factors



Note: Lines indicate the historical performance of portfolios based on the indicated factors for state municipal bond indices from S&P, beginning in January 2020. The left panel includes four style factors, while the right panel shows factors based on state-level pandemic characteristics. A portfolio for each factor is long on the top quintile of the distribution and short on the bottom quintile. Vertical lines indicate the beginnings of the Phase 1, 2, and 3 subperiods.

Return dynamics reversed rapidly between mid-May and mid-September, a period we interpret as a return to normal. Previously distressed bonds recovered more rapidly than the average bond, generating high returns on the carry and value factors. The same recovery

moved the momentum factor in the opposite direction: shorting previous low performers became a money-losing strategy. Returns on all four style factors moderated substantially after mid-September as pandemic conditions became more stable.

Over the same periods, the COVID factors generally moved in the opposite direction from the style factors, although with somewhat different timing. During the early pandemic period, going long states with favorable pandemic conditions and short states with unfavorable conditions produced positive returns. As the return to normal unfolded and distressed states experienced a more rapid recovery in asset prices, the COVID factors began losing money. After mid-September, the COVID factors became largely irrelevant, as indicated by the flat portions on the right panel of Figure 4.

3.2.4 Model with COVID Factors

Adding COVID factors changes how the model interprets the reshuffling of relative yields in Figure 3: it separates movements in returns correlated with the pandemic situation from those uncorrelated with it. COVID factors now form a group associated with low risk and national policy. Accordingly, contributions to predicted returns and to the cross-sectional variance shift away from the low-risk and Treasury slope factors toward COVID factors (compare lines 3–5 of Tables A3 and A2): the COVID group accounts for 12 percent of the cross-sectional variance, while the Treasury share falls from 23 percent to under 10 percent, so that the apparent Phase 1 pricing power of the slope factor is absorbed. Because the cross-section is now interpreted differently, momentum re-emerges with a borderline significant risk premium (Table 3, column 3). Unsurprisingly, the market factor still dominates both aggregate risk premia and the cross-sectional variance (Table A3, columns 3 and 6).

Phase 2 brings a dramatic change in the size, signs and composition of risk premia (Table 3, column 4; Table A4). As the market recovers, the risk premium on the market factor turns large, positive and significant. As the most distressed states recover fastest, the momentum, low-risk and COVID factor premia turn large and negative, with signs consistent with the dynamics in Figure 4. Exposures to style and COVID factors have low means and

high variances, whereas market betas remain tightly distributed around their mean of 0.90. Consequently, relative risk premia now play an outsized role: they account for 13 percent of total fitted-return variance versus 1–1.5 percent in Phase 1. In the cross-section, style and COVID factors account for 37 and 33 percent of variance, respectively, and the market factor’s share is much diminished, at 12 percent (Table A4, columns 5 and 6).

In Phase 3, COVID factors lose relevance and the bond market factor is the only one priced; with carry and value still collinear with the market factor, the model is essentially a one-factor model in which each state’s return is determined by its exposure to the broad municipal bond market. All other factors have dispersed betas and volatile premia. Accordingly, relative risk premia still contribute more to return variance than in normal times (Table A5, column 4, line 6), and the market, style, Treasury and COVID groups contribute roughly equal shares – 22, 20, 26 and 29 percent – of the cross-sectional variance (Table A5, columns 5–6). We interpret this phase as a partial return to normal, with volatility still too high for the baseline factor structure to fully re-emerge.

Comparison with normal times. Three features of the pre-COVID factor structure change during the pandemic. First, the market factor continues to dominate the variance of predicted returns, but its risk becomes unstable: average market betas rise from 0.95 to 1.06 and disperse in Phase 1, the aggregate premium becomes volatile enough to lose significance, and carry and value, which in normal times account for relative returns, become collinear with the market factor. Second, the balance between aggregate and relative risk premia shifts. The March crash is an almost purely aggregate event, with relative premia accounting for only 1–1.5 percent of predicted-return variance (versus about 2.5 percent in 2014–2019); by contrast, the summer recovery shifts 13 percent of the variance to relative premia, and style and COVID factors together account for over 70 percent of the cross-sectional variance in Phase 2, compared with 55 percent for style factors alone before the pandemic. Third, own-state pandemic outcomes do not explain pricing errors, whereas portfolios sorted on the cross-state distribution of cases and ICU occupancy do, and their inclusion absorbs the Phase 1 pricing power of the Treasury slope factor. By Phase 3 the model collapses toward a one-factor structure with relative premia still above normal levels, a partial return

to the pre-COVID regime.

4 Conclusion

This paper applies the tools developed for pricing equities, sovereign debt, and corporate credit to pricing municipal bonds. By using state-level portfolios as a unit of analysis, we were able to make the common factor structure of municipal returns more transparent and comparable with that of other asset classes. The resulting risk premia decomposition clarifies the role of systematic risk, whether tied to the national municipal market, to style-based characteristics, or, during periods of stress, to state-specific exposures that temporarily dominate pricing.

The normal-times results indicate that a single aggregate factor accounts for the bulk of time-series variation in state returns, while heterogeneous style exposures account for most of the cross-sectional variation. This split between aggregate and relative risk premia has practical implications for portfolio construction and risk management: it suggests that diversification across states offers limited protection against national shocks but meaningful protection against style-related risks.

The paper’s tax analysis points to market segmentation related to asymmetric federal taxation of municipal bonds and Treasuries. Our analysis of state taxes confirms the presence of a convenience premium tied to after-tax investment opportunities that suppress municipal returns in high-tax states. Moreover, the evidence suggests that this convenience premium is partially priced in to the model factors.

The COVID-period analysis shows that the model’s factor structure is regime dependent. A shock that was national in origin but heterogeneous across states made some conventional style factors collinear with the market, shifted variance from aggregate toward relative risk premia as the market recovered, and introduced a state variable, the cross-state distribution of pandemic outcomes, that priced the cross-section of returns even though

states' own outcomes did not. The COVID episode also carries a methodological lesson: when a state variable is omitted, its pricing power can be misattributed to an included factor, as it happened with the Treasury slope in the early crash. A factor that appears to gain relevance in a crisis can therefore point to potential missing variables. The model accommodated the new regime once factors sorted on the missing variable were added, which suggests a general approach for applying the framework to other severe aggregate shocks. Whether similar shifts and reversals in risk transmission occur around such episodes is a natural direction for future research.

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A Variance Decomposition in the Factor Model

State-time decomposition of returns. There are S states observed for T days. All variables below are estimated quantities, so we use γ instead of $\hat{\gamma}$, β instead of $\hat{\beta}$ and λ instead of $\hat{\lambda}$ for more compact notation. To compare the contribution of first-stage state intercepts γ_i with those of the factor components $\lambda_t\beta_i$, we define the augmented predicted return for state i in day t as

$$\mu_{it} = \gamma_i + \lambda_t\beta_i,$$

where γ_i is a state fixed effect from (1), β_i ($N \times 1$) is the vector of state i 's factor exposures, λ_t ($1 \times N$) is the vector of time- t factor risk premia.

Define cross-state and time averages:

$$\bar{\gamma} = \frac{1}{S} \sum_{i=1}^S \gamma_i, \quad \bar{\beta} = \frac{1}{S} \sum_{i=1}^S \beta_i \ (N \times 1), \quad \bar{\lambda} = \frac{1}{T} \sum_{t=1}^T \lambda_t \ (1 \times N).$$

Also define demeaned quantities:

$$\tilde{\gamma}_i = \gamma_i - \bar{\gamma}, \quad \tilde{\beta}_i = \beta_i - \bar{\beta} \ (N \times 1), \quad \tilde{\lambda}_t = \lambda_t - \bar{\lambda} \ (1 \times N).$$

The panel average predicted return is

$$\bar{\mu} = \frac{1}{ST} \sum_{i=1}^S \sum_{t=1}^T \mu_{it} = \bar{\gamma} + \bar{\lambda}\bar{\beta}.$$

Decompose the demeaned predicted return into the state component, time component and the state-time interaction term:

$$\begin{aligned} \mu_{it} - \bar{\mu} &= \mu_{it} - \bar{\gamma} - \bar{\lambda}\bar{\beta} = \gamma_i - \bar{\gamma} + \lambda_t\beta_i - \bar{\lambda}\bar{\beta} \quad (\text{add and subtract } \bar{\lambda}\beta_i) \\ &= \left[\gamma_i - \bar{\gamma} + \bar{\lambda}\tilde{\beta}_i \right] + \tilde{\lambda}_t\beta_i = \underbrace{\left[\gamma_i - \bar{\gamma} + \bar{\lambda}\tilde{\beta}_i \right]}_{\text{State component}} + \underbrace{\tilde{\lambda}_t\bar{\beta}}_{\text{Time component}} + \underbrace{\tilde{\lambda}_t\tilde{\beta}_i}_{\text{State-time interaction}}. \end{aligned}$$

In more compact notation, we have

$$\mu_{it} - \bar{\mu} = P_i + A_t + R_{it},$$

where

$$P_i = \frac{1}{T} \sum_t (\mu_{it} - \bar{\mu}) = \tilde{\gamma}_i + \bar{\lambda} \tilde{\beta}_i, \quad (\text{A.1})$$

$$A_t = \frac{1}{S} \sum_i (\mu_{it} - \bar{\mu}) = \tilde{\lambda}_t \bar{\beta}, \quad (\text{A.2})$$

$$R_{it} = \mu_{it} - \bar{\mu} - P_i - A_t = \tilde{\lambda}_t \tilde{\beta}_i. \quad (\text{A.3})$$

State component P_i – persistent risk premium – captures higher risk premia due to higher state fixed effects γ_i and higher betas on factors with positive average $\bar{\lambda}^n$. Time component A_t captures common changes in predicted returns over time – time-varying aggregate risk premia. State-time interaction component R_{it} captures differential responses of more/less exposed states to movements in risk premia. In other words, it is a measure of time-varying relative risk premia.

Total predicted-return variance decomposition. Define the total variance of predicted returns as

$$V_\mu = \frac{1}{ST} \sum_{i=1}^S \sum_{t=1}^T (\mu_{it} - \bar{\mu})^2.$$

Because P_i , A_t , and R_{it} are mutually orthogonal in an $S \times T$ panel,

$$V_\mu = V_P + V_A + V_R,$$

where

$$V_P = \frac{1}{S} \sum_{i=1}^S P_i^2, \quad V_A = \frac{1}{T} \sum_{t=1}^T A_t^2, \quad V_R = \frac{1}{ST} \sum_{i=1}^S \sum_{t=1}^T R_{it}^2.$$

Thus we have a decomposition that can apportion the return variance between persistent, aggregate time-varying and relative time-varying risk premia:

$$Var_{i,t}(\mu_{it}) = Var_i(\gamma_i + \bar{\lambda} \beta_i) + Var_t(\lambda_t \bar{\beta}) + Var_{i,t}[(\lambda_t - \bar{\lambda})(\beta_i - \bar{\beta})].$$

Factor contributions to variance. To determine which factors or groups of factors are associated with each of the V_P , V_A , V_R , we perform a covariance-based decomposition that can rank factors in relative importance as contributors to total variance of a particular component. Covariance-based decomposition uses the relation

$$Var\left(\sum_n x_n\right) = \sum_n Cov\left(x_n, \sum_n x_n\right).$$

Let ϕ_n^C denote the contribution of factor $n = 1, \dots, N$ to risk premium component $C = \{P, A, R\}$, with

$$V_C = \sum_n \phi_n^C.$$

If one factor amplifies the variation of a component, i.e. $\phi_n^C > V_C$, another factor must offset it and have $\phi_l^C < 0$.

Using (A.1), write the fixed effect contribution to the persistent risk premium component V_P as

$$\phi_0^P = Cov_i(\tilde{\gamma}_i, P_i),$$

and, for factor $n \geq 1$,

$$\phi_n^P = Cov_i(\bar{\lambda}^n \tilde{\beta}_i^n, P_i), \quad \text{with} \quad \sum_{n=0}^N \phi_n^P = V_P.$$

Similarly, from (A.2) we have for $n \geq 1$

$$\phi_n^A = Cov_t(\tilde{\lambda}_t^n \bar{\beta}^n, A_t), \quad \text{with} \quad \sum_{n=1}^N \phi_n^A = V_A,$$

and from (A.3) we have for $n \geq 1$

$$\phi_n^R = Cov_{i,t}(\tilde{\lambda}_t^n \tilde{\beta}_i^n, R_{it}) = \sum_{l=1}^N Cov_t(\lambda_t^n, \lambda_t^l) Cov_i(\beta_i^n, \beta_i^l).$$

Grouped factor decomposition. Let g index the subvectors of λ_t and β_i corre-

sponding to just the factors included in group $g \subseteq \{1, \dots, N\}$. Because we have covariance-based contributions of each individual factor, the group contribution is simply the sum of contributions of the factors included in the group:

$$\phi_g^C = \sum_{n \in g} \phi_n^C, \quad C = \{P, A, R\}.$$

Contributions to total and cross-sectional variance. The variance decompositions report the shares (percentages) s_g^C and ω_g^C that group g contributes to the total variance $V_\mu = V_P + V_A + V_R$ and to the cross-sectional variance $V_P + V_R$, respectively. The shares are defined as follows:

$$s_g^C = \frac{\phi_g^C}{V_\mu}, \quad C = \{P, A, R\}, \quad (\text{A.4})$$

$$\omega_g^C = \frac{\phi_g^C}{V_P + V_R}, \quad C = \{P, R\}. \quad (\text{A.5})$$

Tables [A1–A5](#) report the variance decompositions in the factor model for the pre-COVID sample and the three COVID subperiods.

Table A1: Variance Decomposition: Pre-COVID Sample, Baseline Model

Factor group, g	Contrib. to predicted returns, $\lambda_g \beta_g$, annual bp	Share in total variance, percent			Share in cross-sectional variance, percent	
		Persistent, s_q^P	Aggregate, s_q^A	Relative, s_q^R	Persistent, ω_q^P	Relative, ω_q^R
State FE (γ_i)	1	0.00	0.0	0.00	0.17	0.0
Market	541	0.00	96.7	0.84	0.18	33.2
Style (mom/val/carry/lr)	2	0.02	0.7	1.36	0.89	53.8
Treasury (slope/convex)	-4	0.00	0.1	0.30	0.02	11.8
TOTAL (factors+FE)	540	0.03	97.5	2.49	1.26	98.7

Note: “Persistent”, “Aggregate” and “Relative” refer to the persistent (P), aggregate time-varying (A) and relative time-varying (R) risk premium components defined in Equations A.1–A.3. Column 1 reports the contribution of each factor group to the average predicted return, in annual basis points. Columns 2–4 report the shares (in percent) of the total variance of predicted returns, $V_\mu = V_P + V_A + V_R$, attributable to each component for each factor group, as defined in Equation A.4. Columns 5–6 report the corresponding shares of cross-sectional variance, $V_P + V_R$, as defined in Equation A.5. The TOTAL row sums over all factor groups and the state fixed effects. Data are daily state municipal bond indices from S&P; COVID factors are from IHME. This note applies to Tables A1–A5.

Table A2: Variance Decomposition: COVID Period 3/1–5/15, Baseline Model, Dropping Collinear Factors

Factor group, g	Contrib. to predicted returns, $\lambda_g \beta_g$, annual bp	Share in total variance, percent			Share in cross-sectional variance, percent	
		Persistent, s_q^P	Aggregate, s_q^A	Relative, s_q^R	Persistent, ω_q^P	Relative, ω_q^R
State FE (γ_i)	3	0.02	0.0	0.00	1.45	0.0
Market	-1964	0.01	98.4	0.71	0.70	62.7
Style (mom/lr)	144	0.02	0.5	0.12	1.37	10.5
Treasury (slope/convex)	71	0.03	0.0	0.23	2.67	20.6
TOTAL (factors+FE)	-1746	0.07	98.9	1.06	6.20	93.8

Note: See notes for Table A1.

Table A3: Variance Decomposition: COVID Period 3/1–5/15, Baseline + COVID Factors, Dropping Collinear Factors

Factor group, g	Contrib. to predicted returns, $\lambda_g \beta_g$, annual bp	Share in total variance, percent			Share in cross-sectional variance, percent	
		Persistent, s_q^P	Aggregate, s_q^A	Relative, s_q^R	Persistent, ω_q^P	Relative, ω_q^R
State FE (γ_i)	8	0.02	0.0	0.00	1.00	0.0
Market	-1963	0.00	99.8	1.11	0.19	69.1
Style (mom/lr)	46	0.04	-0.7	0.10	2.20	5.9
Treasury (slope/convex)	-10	0.01	0.1	0.15	0.43	9.3
COVID factors	171	0.01	-0.9	0.18	0.53	11.4
TOTAL (factors+FE)	-1746	0.07	98.4	1.54	4.36	95.6

Note: See notes for Table A1.

Table A4: Variance Decomposition: COVID Period 6/1–8/15, Baseline + COVID Factors, Dropping Collinear Factors

Factor group, g	Contrib. to predicted returns, $\lambda_g \beta_g$, annual bp	Share in total variance, percent			Share in cross-sectional variance, percent	
		Persistent, s_q^P	Aggregate, s_q^A	Relative, s_q^R	Persistent, ω_q^P	Relative, ω_q^R
State FE (γ_i)	4	0.70	0.0	0.0	3.9	0.0
Market	1531	0.98	82.1	1.2	5.4	6.5
Style (mom/lr)	-34	2.56	0.2	4.2	14.2	23.4
Treasury (slope/convex)	-7	0.16	1.5	2.3	0.9	12.7
COVID factors	-1	0.48	-1.8	5.5	2.7	30.3
TOTAL (factors+FE)	1492	4.88	81.9	13.2	27.0	73.0

Note: See notes for Table A1.

Table A5: Variance Decomposition: COVID Period 9/1–12/15, Baseline + COVID Factors, Dropping Collinear Factors

Factor group, g	Contrib. to predicted returns, $\lambda_g \bar{\beta}_g$, annual bp	Share in total variance, percent			Share in cross-sectional variance, percent	
		Persistent, s_g^P	Aggregate, s_g^A	Relative, s_g^R	Persistent, ω_g^P	Relative, ω_g^R
State FE (γ_i)	4	0.18	0.0	0.0	3.1	0.0
Market	736	0.09	92.5	1.2	1.5	20.8
Style (mom/lr)	4	-0.01	1.0	1.2	-0.2	19.7
Treasury (slope/convex)	4	0.01	1.0	1.5	0.1	25.5
COVID factors	4	0.10	-0.4	1.6	1.8	27.6
TOTAL (factors+FE)	751	0.37	94.2	5.5	6.3	93.7

Note: See notes for Table A1.