# Fiscal Policy, Consumption Risk, and Stock Returns: Evidence from US States*

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**Abstract**

We find evidence that consumption volatility is lower in states that implement counter-cyclical fiscal policies. The lower consumption volatility in counter-cyclical states is confirmed using balanced budget amendments and the adoption of budget stabilization funds that determine state-level fiscal policy. In addition, firms headquartered in states with counter-cyclical fiscal policies have lower average stock returns provided their investors have a local investment bias. Thus, the consumption risk of geographically segmented investors appears to be sensitive to state-level fiscal policy. However, fiscal policy's impact on firm-level cash flow is limited to geographical concentrated firms that operate in a small number of states. Overall, by reducing consumption risk, we find evidence that counter-cyclical fiscal policies can influence asset prices.

Keywords: Fiscal Policy, Consumption, Local Bias, Stock Returns

# 1 Introduction

The impact of federal government stimulus on the national economy is highly controversial. These debates often center on the implications of different fiscal policies for unemployment (Battaglini and Coate, 2011) and investment (Alesina, Ardagna, Perotti, and Schiantarelli, 2002). In contrast, using state-level data, we investigate the impact of fiscal policy on consumption volatility as well as stock returns.[1] Korniotis (2008) highlights three advantages of using state-level data. First, state-level data has less measurement error than individual data. Second, income shocks are not fully diversified across states. Third, there is considerable cross-sectional variation in consumption across states. Finally, in contrast to country-level data, our analysis is able to utilize state-level balanced budget amendments as instruments since these amendments determine a state's fiscal policy but are exogenous with respect to the economic conditions that influence consumption risk and stock returns.

To the best of our knowledge, we are the first to investigate the impact of fiscal policy on consumption risk. Although Johnson, Parker, and Souleles (2006) along with Parker, Souleles, Johnson, and McClelland (2011) examine the consumption implications of tax rebates from the federal government, their analyses focus on national recessions in 2001 and 2008, respectively. In contrast, we examine the relationship between consumption volatility, defined over a long time series, and fiscal policy. Therefore, our empirical analysis exploits cross-sectional variation in state-level fiscal policy.

Our paper provides two main empirical contributions. First, we report that counter-cyclical fiscal policies lower consumption volatility. This finding has implications for the intertemporal sharing of consumption risk. Indeed, building on the first result, our second finding is that firms headquartered in states with counter-cyclical fiscal policies have lower average stock returns provided their investors have a local investment bias.[2] These return results verify the economic significance of cross-sectional variation in state-level consumption volatility. Intuitively, as local investors are more likely to be the marginal investor in a firm, the consumption risk of local investors is likely to influence its expected return. This intuition is consistent with the ability of correlated local trades to create local risk factors, as documented in Pirinsky and Wang (2006) and more recently by Korniotis and Kumar (2012) as well as

[1] The magnitude of state government expenditures parallels those of the federal government. In 2008, per capita federal expenditures were $9,831 versus average state expenditures of $9,335 according to the Statistical Abstract of the United States.

[2] Investors that overweight, relative to the market portfolio, firms headquartered in the same state are defined as having a local investment bias.

Kumar, Page, and Spalt (2012).

A state's fiscal policy is identified by the sensitivity of its budget deficit (surplus) to state-level economic conditions. The empirical methodology we implement to estimate each state's fiscal policy "beta" parallels the identification of state-level fiscal policy in Svec and Kondo (2012). Intuitively, counter-cyclical states have deficits in bad economic conditions and surpluses in good economic conditions, which results in a negative fiscal policy beta. Annual state-level economic conditions are defined by per capita growth rates in gross state product (GSP). Two proxies for consumption at the state-level are examined. The first proxy is retail sales (Ostergaard, Sorensen, and Yosha, 2002; Korniotis, 2008) and the second is electricity consumption (Da and Yun, 2010, Ferson and Lin, 2011). Both proxies indicate that consumption volatility is lower in counter-cyclical states.

To address the common dependence of government budget deficits and consumption on economic conditions, we find that a combination of balanced budget amendments and state-level stabilization funds explain differences in fiscal policy across states. The Advisory Commission on Intergovernmental Relations (1987) summarizes the stringency of each state's balanced budget amendment, many of which were adopted before the Civil War, by assigning states an ACIR score between zero and ten. A higher ACIR score corresponds to a more stringent balanced budget amendment. Furthermore, many states adopted budget stabilization ("rainy day") funds during our 1965 to 2008 sample period to institutionalize savings. States with lower ACIR scores and states that were early adopters of budget stabilization funds are more likely to implement counter-cyclical fiscal policies. Intuitively, counter-cyclical fiscal policies are facilitated by a less stringent balanced budget amendment and more accumulated savings in a budget stabilization fund. Our instrumental variables procedure constructs predicted fiscal policy betas for each state. These predicted fiscal policy betas confirm that consumption volatility is lower in counter-cyclical states.

We then examine two channels that enable the intertemporal sharing of consumption risk in counter-cyclical states to influence stock returns; a discount rate channel arising from the geographical segmentation of investors, and a cash flow channel arising from the geographical segmentation of firm operations.

Under the discount rate channel, the geographical segmentation of investors due to local investment biases provides a mechanism for the transmission of state-level economic shocks into the stock prices of firms headquartered in the same state.[3] After documenting significant

[3] The literature on local investment biases includes important contributions by Coval and Moskowitz (1999) as well as Grinblatt and Keloharju (2001).

return co-movement among firms headquartered in the same state, Pirinsky and Wang (2006) conclude that equity investment has a significant geographic component since trading activity is correlated among local investors. The return co-movement attributable to this correlated trading activity is not induced by commonality in fundamentals (earnings) or economic conditions. Korniotis and Kumar (2013) along with Kumar, Page, and Spalt (2012) confirm that the trades of local investors are correlated, and conclude that a firm's cost of capital depends on the location of its headquarters as a consequence. Hong, Kubik, and Stein (2008) also report that stock prices are higher in regions with a low population density due to inefficient risk sharing caused by local investment biases. Overall, the combination of correlated trading by local investors and their local investment biases creates a local consumption risk factor.

According to our first result, government budget deficits partially finance the consumption of investors in bad economic conditions. Therefore, the discount rate channel posits that counter-cyclical fiscal policies reduce the need for correlated equity sales by local investors in bad economic conditions, thereby lowering the equity return premium.[4] Flow betas that capture the sensitivity of state-level mutual fund flows to state-level economic conditions provide empirical support for the discount rate channel. Intuitively, a lower flow beta implies less redemptions from mutual funds in bad economic conditions. The flow betas indicate that local investors in counter-cyclical states sell less of their equity portfolio during bad economic conditions than their counterparts in pro-cyclical states.

To clarify, the discount rate channel does not assume the risk premium for firms headquartered in a state is entirely driven by this state's consumption risk. This scenario would coincide with perfect market segmentation for each state. Instead, invoking the prior literature that documents the geographic segmentation of investors, the discount rate channel posits that the lower consumption risk confronted by local investors in counter-cyclical states is sufficient to lower equity return premiums, holding other risk factors constant.

In contrast, local investment biases are irrelevant under the alternative cash flow channel. Instead, by smoothing the consumption of households (including non-investors), counter-cyclical fiscal policies lower firm-level cash flow risk according to the cash flow channel. This cash flow channel is predicted to be stronger for firms whose operations are concentrated in a small number of states with common fiscal policies. Overall, the discount rate and cash flow channels operate through investors and firms, respectively, since local investment biases

[4]Hong, Wang, and Yu (2008) demonstrate that firms capable of re-purchasing their own shares have lower return volatility after controlling for firm fundamentals. The price support provided by firm-level share repurchases parallels the price support from counter-cyclical fiscal policies at the state-level.

motivate the discount rate channel while local firm operations motivate the cash flow channel.

Lower fiscal policy betas, and lower predicted fiscal policy betas based on balanced budget amendments and budget stabilization funds, are associated with lower stock returns for firms that have a local investor base. Thus, counter-cyclical fiscal policies appear to lower equity returns through the discount rate channel. The average return for firms headquartered in counter-cyclical states with local investment biases is 2.03% lower than in pro-cyclical states. For firms with a local investor base, Fama-MacBeth regression coefficients indicate that return variation of 1.56% can be attributed to variation in state-level fiscal policy. Moreover, variation in consumption volatility is sufficient to explain variation in state-level equity premiums since implied risk aversion parameters, which condition on consumption volatility, are similar across counter-cyclical and pro-cyclical states.

We also extend our analysis to nearly 1,000 firm-level headquarter relocations. Firms relocating to a more counter-cyclical state subsequently have lower average stock returns. In addition, local investment biases migrate with firms since investors in the new headquarter state develop a local bias towards the firm while this bias declines in the old headquarter state. Pirinsky and Wang (2006)'s study of return co-movement also documents this migration.

Unlike consumption volatility, the cash flow volatility of firms headquartered in counter-cyclical states is not lower. This finding is consistent with firm cash flows being obtained from multiple states (and countries) with diverse fiscal policies. Therefore, using Garcia and Norli (2012)'s data on the state-level operations of individual firms, we compute firm-level cash flow betas by weighting the state-level fiscal policy betas according to each firm's operations. However, these firm-level cash flow betas are unable to explain stock returns. In conjunction with the importance of local investment biases, this evidence suggests that the cash flow channel is not the primary link between fiscal policy and stock returns. Korniotis and Kumar (2013) also conclude that a firm's cost of capital depends on the location of its headquarters due to the discount rate channel rather than the cash flow channel. Nonetheless, for geographically concentrated firms that operate in a small number of states, the cash flow betas derived from state-level fiscal policy are relevant to stock returns. This finding highlights the importance of the cash flow channel to firms whose operations are not diversified across the US.

The remainder of the paper begins with a brief literature review in Section 2. The data underlying our empirical study is then described in Section 3. Section 4 presents our state-level estimates for fiscal policy while Section 5 examines the empirical relationship between fiscal policy and consumption. The impact of fiscal policy on average state-level and firm-level returns are presented in Section 6 with our conclusions following in Section 7.

# 2 Literature Review

Julio and Yook (2012) conclude that investment declines around national elections while Durnev (2010) reports that investment is less sensitive to stock prices during election periods. In an international context, Brogaard and Detzel (2012) construct a country-specific proxy for economic policy uncertainty and report that greater uncertainty reduces investment. In contrast to this literature, our study examines the consumption and stock return implications of government fiscal policy.

Santa-Clara and Valkanov (2003) find that stock returns are higher and real interest rates are lower during Democratic presidencies. Belo, Gala, and Li (2013) report that the market is positively surprised by the spending policies of Democratic presidents and negatively surprised by those of Republican presidents. In contrast to the time-series methodology in these studies, we examine cross-sectional variation within the United States and find evidence that state-level balanced budget amendments exert a greater influence on fiscal policy than political affiliations. Moreover, consumption-based asset pricing rather than errors in investor expectations motivates our study.

Although our cross-sectional tests are limited to 50 states, international studies often involve fewer countries and their conclusions are complicated by differences in labor markets as well as legal, political, and monetary institutions (Acemoglu, Johnson, Robinson, and Thaicharoen, 2003). Moreover, balanced budget amendments at the country-level are not available to serve as instruments for fiscal policy.

For emphasis, balanced budget amendments apply to budget deficits. Therefore, state governments can adjust taxation and spending policies to ensure compliance with their balanced budget amendment. Consequently, when estimating state-level fiscal policy betas, our methodology focuses on budget deficits instead of government revenue or government expenditures separately. Moreover, the possibility that different government expenditures (highways versus health care for example) have distinct impacts on consumption is also ignored.

Fiscal policy also has tax implications. Specifically, counter-cyclical fiscal policies may increase tax uncertainty as higher taxes in good economic conditions are needed to fund deficits in poor economic conditions. In Croce, Kung, Nyuyen, and Schmid (2012)'s model, tax uncertainty can be as important as the level of taxation to the cost of equity. Although high-income households finance a large portion of government deficits under progressive tax rates, counter-cyclical fiscal policies that expand available tax credits during recessions benefit these households. The expectation of tax increases can also increase the savings and equity investments of high-income households.

# 3 Data

Annual state-level data on government revenue and expenditures as well as state-level data on GSP and income from 1965 to 2008 are obtained from the Statistical Abstract of the United States maintained by the United States Census. Individual firms are matched with specific states using the location of their headquarters in COMPUSTAT. As in Korniotis and Kumar (2013), our analysis of stock returns from CRSP focuses on 35 states that contain the headquarters of at least 20 firms.

## 3.1 State-Level Consumption Proxies

Two proxies for state-level consumption are examined. The first proxy is retail sales (Ostergaard, Sorensen, and Yosha, 2002; Korniotis, 2008) and the second is electricity consumption (Da and Yun, 2010; Ferson and Lin, 2011). As in Korniotis (2008), retail sales data on nondurables is scaled upward to account for services. This scale factor is the same across all states each year but varies over time. The annual scale factor equals aggregate per capita US consumption divided by average per capita US retail sales. Therefore, the scale factor assumes that services are the same proportion of consumption in every state. Our second proxy for state-level consumption is residential and commercial electricity usage. To the extent that many modern-day consumption activities use electricity, which cannot be stored, electricity usage is a suitable proxy for real-time consumption. Although electricity is measured in thousands of kilowatt hours, our analysis converts these units into percentage growth rates. Da and Yun (2010) demonstrate that electricity usage produces salient asset pricing implications that are consistent with consumption-based capital asset pricing models.

Panel A of Table 1 provides summary statistics for state-level returns, consumption, and gross state product (GSP). Consumption and GSP are expressed as annual per capita growth rates to mitigate the effects of interstate migration. As expected, consumption and GSP are more autocorrelated than stock returns.

According to Panel B of Table 1, GSP volatility has a positive cross-sectional correlation with the volatility of both consumption proxies. Although GSP volatility is partially influenced by fiscal policy since government spending is a large component of economic activity, a state's fiscal policy may also reflect its industrial composition. For example, states with pro-cyclical industries such as New York's finance industry may implement counter-cyclical fiscal policies in an attempt to lower GSP volatility. Consequently, industry-adjusted returns are examined when analyzing the implications of fiscal policy. Moreover, we include the volatility of each state's GSP as a control variable in later empirical tests along with state fixed effects.

The average time series correlation of 0.253 between the consumption growth proxies in Panel C of Table 1 indicates that within a particular state, electricity usage and retail sales fluctuate in tandem. Furthermore, the Consumer Expenditure survey (CE) contains detailed consumption data from households in four regions (Northeast, Midwest, South, and West). Therefore, we aggregate our consumption proxies according to each region and compute their time series correlations with the regional CE observations staring in 1984. In unreported results, the 0.206 correlation between the regional CE data and electricity consumption and the 0.245 correlation between the regional CE data and retail sales provide additional validation of our consumption proxies.

## 3.2 Fiscal Policy Instruments

State-level balanced budget amendments determine a state's ability to implement counter-cyclical fiscal policies. As balanced budget amendments were adopted before the start of our sample in 1965, they provide ideal instruments for fiscal policy. The stringency of each state's balanced budget amendment is summarized by an ACIR index produced by the Advisory Commission on Intergovernmental Relations (1987) ranging from 0 (least stringent) to 10 (most stringent). The weakest amendments require the governor to submit a balanced budget to the legislature or the legislature to propose a balanced budget. More stringent amendments require budget deficits to be corrected in the following fiscal year or within two fiscal years. Balanced budget amendments are either constitutional or statutory. Bohn and Inman (1996) emphasize the stringency provided by end-of-the-year (not beginning-of-the-year) constitutional budget restrictions (not statutory) enforced by independent (not politically appointed) state supreme courts. Consequently, the ACIR index accounts for the enforcement of balanced budget amendments.[5] Poterba (1994) also uses the ACIR variable to explain the response of fiscal policy to unexpected deficits.

[5]The National Conference of State Legislatures (2004) estimates that at least 75% of government expenditures are affected by balanced budget amendments, which apply to a state's general fund. Furthermore, Bohn and Inman (1996) conclude that balanced budget amendments are not responsible for transferring general fund deficits to other funds such as employee pensions.

Many states have also adopted budget stabilization funds. Wagner and Elder (2005) describe the deposit and withdrawal rules associated with these "rainy day" funds. These authors also document that government expenditures are less volatile in states with stabilization funds as a consequence of their ability to institutionalize government savings. Several states adopted budget stabilization funds during our sample period but these adoptions appear to be uncorrelated with prevailing economic conditions. In unreported results, GSP growth in the year of adoption and the preceding year are typically near the state's median growth rate. Thus, the adoption of a budget stabilization fund is not in response to particularly good or bad economic conditions.

Overall, balanced budget amendments and budget stabilization funds provide instruments to explore the relationships between fiscal policy, consumption volatility, and average stock returns. Unreported results confirm the effectiveness of balanced budget amendments and budget stabilization funds at curbing cumulative state-level debt. Indeed, all but three states have debt-to-GSP ratios that are either constant or declining during our sample period. Therefore, while high levels of government debt bolster Ricardian equivalence by increasing the possibility of tax increases, cumulative state-level debt is unlikely to affect our results.

### 3.3 Local Investment Biases

Local investment biases have been attributed to informational advantages (Ivković and Weisbenner, 2005), familiarity (Huberman, 2001), and social interactions (Ivković and Weisbenner, 2007). Korniotis and Kumar (2013b) find evidence that the motivation for local investment biases differ across investors with information-based and behavioral explanations applying to sophisticated and unsophisticated investors, respectively. Although the portfolios of retail investors have a closer conceptual link with the investment decisions of households, retail investor portfolios are only available from 1991 until 1996. Furthermore, Korniotis and Kumar (2013) document that retail investors display a similar local bias as institutional investors at the state-level. Indeed, the investment decisions of institutional investors likely reflect the investment preferences of the households that comprise their client base.[6]

Therefore, we compute local investment biases using data on institutional investor portfolios. The location of institutional investors is obtained from Nelson's Directory of Investment Managers. State-level local investment biases are constructed as deviations from the market portfolio using the methodology in Korniotis and Kumar (2013). In particular, a state's local

[6] Becker, Ivković, and Weisbenner (2011) report that local investment biases are capable of increasing the payout policies of firms located in areas that contain a high percentage of senior citizens.

investment bias is defined by a two-step procedure. First, we compute differences for each institutional investor between the percentage of local stocks (firms headquartered in the same state as the institutional investor) minus the percentage of local stocks in the market portfolio, with the latter being the same for all investors each quarter. Second, to create a state-level local bias measure, the investor-level local bias measures are then value-weighted according to the market value of the assets under management for each institutional investor.

As the local bias measures are computed after 1980 due to the availability of institutional ownership data, our return results are reported in both the post-1965 and post-1980 periods to address the possible influence of look-ahead bias.

# 4 Identification of Fiscal Policy

To identify states with counter-cyclical fiscal policies, we classify states according to their budget deficits in poor economic conditions and surpluses in good economic conditions. We first define annual budget deficits and surpluses as follows

$$\text{DEFICIT}_{i,t} = \frac{\text{Government Expenditures}_{i,t} - \text{Government Revenue}_{i,t}}{\text{GSP}_{i,t}} . \quad (1)$$

When positive, this state-year observation represents a budget deficit for state $i$ in year $t$. Conversely, when negative, this state-year observation represents a budget surplus. State-level revenue includes transfer payments from the federal government. In the above definition's denominator, GSP represents a dollar-denominated quantity.

The nature of each state's fiscal policy is estimated by the $\beta_{i,1}$ coefficient in the following specification

$$\text{DEFICIT}_{i,t} = \beta_{i,1}\,\text{GSP}^{Q}_{i,t} + \beta_{i,2}\,\text{INCOME}_{i,t} + \beta_{i,3}\,\text{GSP}^{Q}_{i,t-1} + \gamma\, X_i + \epsilon_{i,t} , \quad (2)$$

where $X$ denotes the vector of state fixed effects.

$\text{GSP}^{Q}$ represents a collection of indicator functions for GSP growth quintiles that identify boom and bust periods. For example, $\text{GSP}^{Q}_{i,t}$ equals 1 and 5 if per capita GSP growth in year $t$ is in the bottom quintile (bust period) and top quintile (boom period) of state $i$, respectively. An indicator function for the bottom quintile is used in our later analysis of borrowing constraints. The thresholds that determine the quintiles are computed at the individual state-level over the 1965 to 2008 sample period. As economic conditions are autocorrelated, equation (2) includes the lagged value $\text{GSP}^{Q}_{i,t-1}$. $\text{INCOME}_{i,t}$ refers to the annual per capita rate of income growth in state $i$ between year $t$ and year $t-1$. Svec and Kondo (2012) estimate a regression that parallels equation (2) to identify state-level fiscal policy. However, their study does not investigate the impact of fiscal policy on consumption or stock returns.

The fiscal policy beta $\beta_{i,1}$ measures the sensitivity of government budget deficits / surpluses to contemporaneous economic growth. A negative fiscal policy beta identifies a counter-cyclical fiscal policy since budget deficits occur in poor economic conditions and budget surpluses occur in good economic conditions. Panel B of Table 2 sorts the fiscal policy betas in ascending order to illustrate their cross-sectional variation across states. Although the average fiscal policy beta is near zero, 0.079, there is considerable variation in these betas across the 50 states. The five most counter-cyclical and five most pro-cyclical states have average fiscal policy betas of -3.43 and 4.49, respectively. Panel B also provides a relevant summary of state-level revenues and expenditures. Observe that lower fiscal policy betas are associated with more autocorrelated and less volatile government expenditures. These properties suggest that counter-cyclical states are less likely to reduce their expenditures in poor economic conditions than pro-cyclical states. Counter-cyclical states also adopt taxation strategies that ensure their revenue is less volatile and more persistent. Overall, the summary statistics in Panel B are consistent with the ability of the fiscal policy betas to capture cross-sectional differences in state-level fiscal policy.

For emphasis, state fixed effects capture variation in government spending across states. However, in unreported results, average government spending divided by GSP is similar across states, with the cross-sectional average of the state-level time series averages being 17.11%. The lack of variation around 17.11% indicates that our fiscal policy betas are not manifesting differences between "big government" versus "small government" states.

Higher economic growth decreases the dependent variable in equation (2) while increasing $GSP^Q$, thereby introducing a downward bias in the fiscal policy betas. Moreover, higher government spending results in a larger numerator and a larger denominator for DEFICIT. The net impact is that increased government spending increases budget deficits and $GSP^Q$, thereby introducing an upward bias in the fiscal policy betas. Despite these conflicting biases, applications of the fiscal policy betas involve their relative magnitude.

Furthermore, unreported results confirm that including additional lagged values of GSP growth in equation (2) yields $\beta_{i,1}$ estimates whose correlation with the fiscal policy betas in Table 2 exceeds 0.90. Thus, our fiscal policy betas are robust to the inclusion of additional lagged economic growth rates. Moreover, state-level fiscal policy betas estimated during the pre-1980 and post-1980 periods are similar. For example, states such as New York and Oregon consistently implement counter-cyclical fiscal policies.

# 5 Fiscal Policy and Consumption

According to Ricardian equivalence (Barro-Ricardo equivalence theorem), government budget deficits cannot stimulate consumption since households increase savings in anticipation of future tax increases. Furthermore, provided households can borrow against their future labor income, government spending is unnecessary to finance consumption. Conversely, higher government spending can increase consumption due to household borrowing constraints (Zeldes, 1989) and difficulties hedging unemployment (Hubbard, Skinner, and Zeldes, 1994).[7] Households may also prefer public indebtedness to private debt if government borrowing costs are lower.

Johnson, Parker, and Souleles (2006) along with Parker, Souleles, Johnson, and McClelland (2011) report that the fiscal policy of the US federal government impacts consumption during two national recessions. Using our state-level proxies for consumption, the next subsection examines whether household borrowing constraints reduce consumption in poor economic conditions. This reduction is a necessary condition for counter-cyclical fiscal policies to lower consumption volatility.

## 5.1 Borrowing Constraints and Consumption

State-level banking regulations provide natural proxies for household borrowing constraints. In particular, bank deregulation relaxed borrowing constraints. Becker (2007) documents that banking deregulation improved capital allocation with the US. Prior to the 1970s, interstate banking was not allowed and restrictions on the opening of intrastate branches were common. The economic motivation for these restrictions were the lucrative charter fees collected from banks by state governments. Kroszner and Strahan (1999) list three important deregulations of the banking industry that gradually occurred until the Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994. The first deregulation allowed mergers and acquisitions to facilitate intrastate branching ($j = 1$), the second allowed for intrastate branching ($j = 2$),

[7] Heaton and Lucas (1992) demonstrate that borrowing constraints limit the ability of investors to hedge reductions in labor income. In Lucas (1994), investors self-insure themselves against transitory shocks to labor income through precautionary savings. However, Brav, Constantinides, and Geczy (2002) find evidence that consumption is not completely insured while Storesletten, Telmer, and Yaron (2004) report that precautionary savings provide inadequate insurance against prolonged losses of income (unemployment).

and the third permitted interstate banking ($j = 3$). Table 1 of Kroszner and Strahan (1999) provides the year in which each of these banking deregulations was enacted for individual states.

The relationships between state-level consumption growth and borrowing constraints derived from bank deregulations are estimated by the following panel regression

$$\begin{aligned} \text{Consumption Growth}_{i,t} &= \beta_1 \, 1_{\text{Bad}_{i,t}} + \beta_2 \, 1_{\text{BC}^j_{i,t}} + \beta_3 \left( 1_{\text{Bad}_{i,t}} \times 1_{\text{BC}^j_{i,t}} \right) \\ &\quad + \gamma \, X_i + \epsilon_{i,t} \, . \end{aligned} \tag{3}$$

State fixed effects are included in the $X$ vector while the standard errors are clustered at the state-level. The dummy variable $1_{\text{Bad}}$ equals 1 in years when a state's GSP growth is in its bottom quintile. For state $i$, the budget constraint dummy variable $\text{BC}^j_{i,t}$ equals 1 in years before the relaxation of a particular banking restriction ($j$=1,2,3).

A negative $\beta_1$ coefficient indicates that consumption growth decreases during bad economic conditions, while the $\beta_2$ coefficient pertains to any trend in consumption growth during the sample period. Most importantly, a negative $\beta_3$ coefficient is consistent with consumption growth during bad economic conditions declining more in the presence of borrowing constraints.

According to Table 3, the $\beta_3$ coefficients are negative. Thus, consumption growth decreased more during bad economic conditions before each of the three bank deregulations. Thus, consumption became less sensitive to local economic conditions after the relaxation of borrowing constraints. The importance of borrowing constraints to consumption provides an avenue for government fiscal policy to influence consumption and motivates our next analysis that tests whether counter-cyclical fiscal policies lower consumption volatility. The impact of borrowing constraints on consumption also provides further indirect evidence that electricity usage and retail sales are suitable proxies for state-level consumption.

## 5.2 Fiscal Policy and Consumption Volatility

Counter-cyclical fiscal policies can mitigate household borrowing constraints and allow households to smooth consumption, especially when confronted by negative income shocks. However, the relationships between fiscal policy and consumption as well fiscal policy and stock returns are subject to endogeneity concerns since economic activity determines government tax revenue, household consumption, and stock returns. Therefore, we implement an instrumental variables procedure to address this endogeneity. In the first stage of our instrumental variables procedure, the fiscal policy betas are regressed on dummy variables defined by the ACIR index and the adoption of a budget stabilization fund in each state.

Balanced budget dummy variables are collectively denoted $\text{DACIR}_i$ and equal 0 when the ACIR index for state $i$ is three or below (weak balanced budget amendment), 1 when this index is between four and seven (moderate balanced budget amendment), and 2 when this index is eight or above (stringent balanced budget amendment). The DACIR variable summarizes the ACIR index using three values instead of ten to reduce the number of degrees of freedom in our later empirical specifications. Ricardian equivalence is predicted to be stronger in states with stringent balanced budget amendments since increases in government spending must be offset by higher taxes within a shorter time horizon.

The first stage of our instrumental variables procedure also involves the fraction of the 1965-2008 sample period that each state does not have a stabilization fund. This fraction denoted $\text{FNO}_i$ equals

$$\text{FNO}_i = \begin{cases} 0 & \text{if a stabilization fund was adopted by state } i \text{ before 1965} \\ 1 & \text{if state } i \text{ had no stabilization fund by 2008} \\ \frac{t-1965}{2008-1965} & \text{if state } i \text{ adopts a stabilization fund in year } t \end{cases}$$

The FNO variable for each state is recorded in Table 2, along with the year in which a state adopted its budget stabilization fund, if at all. A low FNO signifies a state was an early adopter of a budget stabilization fund. As reported in Knight and Levinson (1999), budget stabilization funds are intended to institutionalize savings based on prior tax revenue. Thus, all else being equal, a low FNO facilitates a counter-cyclical fiscal policy since the state has more accumulated savings.

In unreported results, the ACIR index and FNO are nearly independent. Indeed, their -0.049 correlation indicates that states with a stringent balanced budget amendment are more likely to adopt a budget stabilization fund. Thus, balanced budget amendments and budget stabilization funds reinforce common fiscal policy objectives. For example, the budget stabilization fund reduces the need for deficit financing that is disallowed by stringent balanced budget amendments.

The first stage of the instrumental variable estimation examines the extent to which the state-level fiscal policy betas denoted FP are explained by balanced budget amendments and budget stabilization funds according to the following cross-sectional regression

$$\text{FP}_i = \alpha_0 + \alpha_1 \, \text{DACIR}_i + \alpha_2 \, \text{FNO}_i + \alpha_3 \, ( \, \text{DACIR}_i \times \text{FNO}_i \, ) + \epsilon_i \, . \tag{4}$$

The positive coefficients for DACIR and FNO in Table 4 indicate that less stringent balanced budget amendments and the early adoption of a budget stabilization fund are associated with counter-cyclical fiscal policies. The early adoption of a budget stabilization fund coincides with greater accumulated savings. The negative coefficient for the interaction between DACIR and FNO indicates that budget stabilization funds inhibit counter-cyclical fiscal policies when they are combined with stringent balanced budget amendments. This finding is consistent with budget stabilization funds requiring deposits that limit expenditures.

In the second stage of the instrumental variables procedure, consumption volatility is regressed on the predicted values for FP from the first stage. The ability of these predicted betas to explain cross-sectional differences in consumption volatility is evaluated by the following cross-sectional regression

$$\text{Consumption Volatility}_i = \gamma_0 + \gamma_1 \hat{\text{FP}}_i + \epsilon_i \,, \quad (5)$$

where the predicted values $\hat{\text{FP}}$ are defined by the $\alpha$ estimates from equation (4)

$$\hat{\text{FP}}_i = \hat{\alpha}_0 + \hat{\alpha}_1 \, \text{DACIR}_i + \hat{\alpha}_2 \, \text{FNO}_i + \hat{\alpha}_3 \, ( \, \text{DACIR}_i \times \text{FNO}_i \, )$$

for state $i$. The instrumental variables estimation is conducted using a Generalized Method of Moments (GMM) procedure with the second stage accounting for the estimation error in the first stage. In unreported results, the findings are similar using Two-Stage Least Squares (2SLS) and Limited Information Maximum Likelihood (LIML).

The underlying $\alpha$ estimates are reported in Table 4 for the cross-section of 50 states and the 35 state subset comprising states with at least 20 firms. A positive $\gamma_1$ coefficient indicates that consumption volatility is lower in states that are predicted to implement counter-cyclical fiscal policies based on their balanced budget amendment and their decision to adopt a budget stabilization fund. According to Table 4, the $\gamma_1$ coefficients are positive for both consumption proxies.[8] Furthermore, these coefficients are positive across the entire cross-section of 50 states and the subset of 35 states that contain at least 20 firms. Indeed, for electricity, the $\gamma_1$ coefficient is 0.5227 ($t$-statistic of 4.51) across all 50 states and 0.6373 ($t$-statistic of 2.48) across the 35 state subset. Similar significance levels are reported when retail sales proxy for consumption. Therefore, the results in Table 4 provide evidence that households in states with weaker fiscal constraints have lower consumption volatility.

[8]The Craig-Watson F-statistic, which applies to the first stage of the instrumental variables procedure, is identical under the GMM, 2SLS and LIML methodologies. However, the critical values of this statistic vary across the methodologies. With our specification involving one endogenous regressor and three instruments, we obtain a significant F-statistic at the 10% level. The weak instrument dilemma is also mitigated by variation in the predicted fiscal policy betas and their high correlation with the original fiscal policy betas.

The proportion of Democratic versus Republican governors during the sample period was also included in the first stage of the instrumental variables procedure. In unreported results, this proportion has an insignificant coefficient and does not alter the predicted fiscal policy betas in the 50 state or 35 state cross-sections. Consequently, balanced budget amendments and budget stabilizations funds rather than political affiliations determine state-level fiscal policy. The lack of empirical support for political affiliations may arise from heterogeneity within the Republican and Democratic parties across the US.

# 6 Return Implications of Fiscal Policy

Salient differences in consumption volatility across states can have asset pricing implications if investors are segmented geographically and therefore confront different levels of consumption risk due to variation in fiscal policy, as well as variation in economic conditions. Indeed, market incompleteness due to insufficient risk sharing within the US has been reported by Korniotis (2008) along with Korniotis and Kumar (2013). We begin our analysis with state-level returns and local biases before examining firm-level returns and local biases.

## 6.1 State-Level Results

The first Fama-MacBeth regression (1973) conditions on our proxy for state-level local investment biases denoted LB. Recall that local investment biases are required to transmit state-level differences in consumption volatility into state-level return variation under the discount rate channel

$$\text{State Return}_i \;=\; \beta_0 + \beta_1 \left(\text{FP}_i \times \text{LB}_i\right) + \beta_2 \,\text{FP}_i + \beta_3 \,\text{GSP}_i + \beta_4 \,\text{GSP volatility}_i + \epsilon_i \,. \quad (6)$$

This regression controls for the average growth and volatility of each state's GSP, which partially reflects differences in each state's industrial composition. A positive $\beta_1$ coefficient is consistent with lower average returns for firms headquartered in counter-cyclical states that have local investment biases. The correlation between FP and LB is 0.125, indicating that local investment biases are only slightly more prevalent in procyclical states. In unreported results, the coefficient for LB is insignificant when this variable is included in equation (6) and its inclusion does not affect the other $\beta$ coefficients.

The state-level return in equation (6) is constructed by value-weighting the risk-adjusted returns of firms headquartered in a particular state. Value-weighting ensures that larger firms, which are more likely to have a geographically disperse investor base, exert more influence on state-level returns.

The first risk-adjustment procedure utilizes the Fama-French four-factor procedure from 1965 to 1975 and the DGTW procedure from 1976 to 2008 based on the characteristics-based methodology of Daniel, Grinblatt, Titman, and Wermers (1997). DGTW-returns are available starting in 1976 from the website of Russ Wermers, although our results are not sensitive to the use of the Fama-French four-factor procedure to risk-adjust returns throughout the entire sample period. The second risk-adjustment procedure is based on industry-adjusted returns using the Fama-French 48 industry classifications.

The positive $\beta_1$ coefficients in Panel A of Table 5 indicate that differences in fiscal policy can explain state-level return variation. However, the ability of fiscal policy to influence average stock returns is limited to states with local investment biases since the $\beta_2$ coefficients are insignificant. In particular, the $\beta_1$ coefficient for the interaction between FP and LB across the 1965 to 2008 sample period is 0.0880 ($t$-statistic of 2.75) with controls for state-level GSP characteristics. This positive coefficient indicates that firms headquartered in counter-cyclical states with local investment biases have lower average returns. The importance of local investment biases supports the discount rate channel but not the cash flow channel.

To confirm the results from equation (6), the median local bias splits the cross-section of 35 states into 17 above-median and 18 below-median subsets. We then estimate the following Fama-MacBeth regression

$$\text{State Return}_i \quad = \quad \beta_0 + \beta_1 \, \text{FP}_i + \beta_2 \, \text{GSP}_i + \beta_3 \, \text{GSP volatility}_i + \epsilon_i \,, \tag{7}$$

within the above-median local investment bias subset. This regression is also conducted with the predicted fiscal policy betas $\hat{\text{FP}}$ from the first stage of our instrumental variables procedure in equation (4) replacing FP. For the above regression specification, a positive $\beta_1$ coefficient is once again consistent with counter-cyclical fiscal policies lowering stock returns in the presence of local investment biases.

The results in Panel B of Table 5 support the findings in Panel A. The $\beta_1$ coefficients for FP and $\hat{\text{FP}}$ are positive in both the post-1965 and post-1980 sample periods. For example, after controlling for state-level differences in GSP characteristics during the entire sample period, the $\beta_1$ coefficient for the fiscal policy betas and their predicted values are 0.0098 ($t$-statistic of 3.69) and 0.0140 ($t$-statistic of 2.56), respectively.

The results in Panel C and Panel D based on industry-adjusted returns parallel those in Panel A and Panel B, respectively. Consequently, the ability of fiscal policy to explain state-level return variation is not driven by industry differences across states. Overall, the evidence in Table 5 is consistent with counter-cyclical fiscal policies lowering average stock returns in states with local investment biases.

## 6.2 Economic Significance

On average, the fiscal policy betas for states with above-median and below-median local investment biases are similar. However, there remains considerable cross-sectional variation in the fiscal policy betas among states with above-median local biases. Within this subset, the five states with the most counter-cyclical fiscal policies (SC, UT, MI, MN, IA) have an average fiscal policy beta of -1.77 while the five most pro-cyclical states (AR, KY, IL, LA, IN) have an average fiscal policy beta of 1.41. Therefore, we construct a "hypothetical" trading strategy that is short firms headquartered in the five counter-cyclical states and long those headquartered in the five pro-cyclical states.

As state-level returns are determined by value-weighting firm-level returns within each state, state-level returns for the long portfolio and short portfolio are equally-weighted across their five respective states. Monthly return differences between the long and short portfolios are then risk-adjusted according to the four factor model. However, as the long and short portfolios are based on fiscal policy betas estimated over the entire sample period, the trading strategy exploits cross-sectional variation that may not have been apparent to investors ex-ante. Therefore, the return spread attributable to this long-short strategy illustrates the economic importance of fiscal policy without offering any judgment on market efficiency.

The unadjusted return spread averages 17bp per month or 2.03% per year. In unreported results, after applying the four factor model, the alpha associated with the trading strategy increases to 21bp per month ($t$-statistic of 2.20) or 2.49% per year. Consequently, return variation attributable to state-level differences in fiscal policy cannot be explained by market risk, the value and size premiums, or momentum.

To understand whether higher consumption volatility in states with pro-cyclical fiscal policies can explain their higher equity premiums, we extract implied risk aversion parameters denoted $\phi$ from the following relationship

$$\text{State Return}_i - R_f \quad = \quad \beta_C \cdot \phi_i \cdot \sigma^2_{C,i} \,. \tag{8}$$

Under the assumption that the consumption beta denoted $\beta_C$ equals one, the implied risk aversion parameter equals

$$\phi_i \quad = \quad \frac{\text{State Return}_i - R_f}{\sigma^2_{C,i}} \,. \tag{9}$$

With a risk-free rate denoted $R_f$ of 4%, we extract $\phi_i$ for each state involved in our hypothetical trading strategy that produces the 2.03% return spread. For example, with electricity usage proxying for consumption, Michigan's consumption volatility is 0.039 while firms headquartered in Michigan have an average return of 0.088 according to Table 1. Therefore, the implied $\phi_i$ parameter for Michigan is $\frac{0.088-0.04}{(0.039)^2} = 32.06$, which falls to 21.76 when consumption is measured using retail sales.

The implied risk aversion parameters capture differences in state-level equity premiums that cannot be explained by differences in consumption volatility. Furthermore, since there is no difference in GSP volatility between the counter-cyclical and pro-cyclical states, differences in consumption volatility are not induced by differences in their economic variability.

In unreported calculations, there is little difference in the average implied risk aversion parameters across pro-cyclical and counter-cyclical states with above-median local investment biases. Indeed, for electricity usage, the average implied $\phi$ is 32.79 for the five most counter-cyclical states (SC, UT, MI, MN, IA) and 33.13 for the five most pro-cyclical states (AR, KY, IL, LA, IN). Therefore, despite slightly higher implied risk aversion in the pro-cyclical subset, the 2.03% difference in the equity premium can primarily be attributed to differences in consumption volatility.

Although the average cost of capital is lower in counter-cyclical states, firms may select the location of their headquarters based on geographic proximity to skilled labor, customers, and suppliers as well as potential acquisitions (Almazan, De Motta, Titman, and Uysal, 2010). Firms may also attempt to diversify their investor base beyond the state in which they are headquartered instead of relocating to a counter-cyclical state.

Finally, state governments may select their fiscal policy to balance consumption volatility and long-term consumption growth. This optimization is complicated by the uncertainty surrounding the impact of fiscal policy on long-term consumption growth. In unreported results, regressing long-term consumption growth in each state on its respective fiscal policy beta yields a positive coefficient for electricity consumption ($t$-statistic of 1.67) and retail sales ($t$-statistic of 0.57). Although this impact is insignificant, the possibility that counter-cyclical fiscal policies lower consumption growth provides an intuitive mean-variance tradeoff for states attempting to optimize their fiscal policy. Nonetheless, while different state governments may choose different fiscal policies from the "efficient consumption" frontier, the return implications of these choices remain interesting and are the focus of our paper.

## 6.3 Firm-Level Results

Our next set of Fama-MacBeth regressions allows for heterogeneity in the local bias of investors towards different firms headquartered in the same state. Thus, average stock returns and local biases are measured at the firm-level in the following specification

$$\text{Firm Return}_k \;=\; \beta_0 + \beta_1\,\text{FP}_i + \beta_2\,\text{GSP}_i + \beta_3\,\text{GSP volatility}_i + \epsilon_k\,. \tag{10}$$

In addition, the fiscal policy beta FP is replaced with its predicted counterpart $\hat{\text{FP}}$ from the first stage of our instrumental variables procedure in equation (4). A positive $\beta_1$ coefficient for either of these fiscal policy proxies is consistent with lower average stock returns in counter-cyclical states. The Fama-MacBeth regressions start in 1980 with the availability of institutional ownership data to compute firm-level local bias measures.

Firm-level returns are risk-adjusted using the DGTW procedure before the $\beta$ coefficients in equation (10) are estimated. The above regression is first estimated for all firms, which results in a sample of 2,510 firms on average per year in 49 states. Our second estimation is restricted to firms with a local investment base. A firm has a local investment base if institutional investors located in the same state as its headquarters hold 20% more than the firm's percentage weight in the market portfolio and these investors also own at least 5% of the firm's outstanding shares. The second criteria requires investors located in the same state as the firm to constitute a relatively large fraction of its shareholder base.[9] The local ownership threshold reduces our sample to an average of 666 stocks per year in 44 states.

Table 6 reports the results from the firm-level Fama-MacBeth regressions. The results in Panel A indicate that the $\beta_1$ coefficients are insignificant in the entire cross-section of firms. However, the coefficients for FP and $\hat{\text{FP}}$ are both positive in Panel B for firms with a local investor base. In particular, the $\beta_1$ coefficients are 0.0039 ($t$-statistic of 2.71) and 0.0051 ($t$-statistic of 2.92), respectively, after controlling for state-level economic growth and volatility. Therefore, consistent with the discount rate channel, local investment biases are required for variation in state-level fiscal policy to induce return variation.

To assess the economic significance of fiscal policy, consider a fluctuation from the average fiscal policy beta for the five most counter-cyclical states with above-mean local investment biases (SC, UT, MI, MN, IA) to this average for the five most pro-cyclical states with above-median local investment biases (AR, KY, IL, LA, IN). This fluctuation in the average fiscal policy beta, from -1.77 to 1.41, totals 3.18. Over an annual horizon, the $\beta_1$ coefficient of 0.0049

[9] This filter accounts for the possibility that local investors only comprise a small fraction of a firm's investor base, perhaps due to out-of-state fund managers from New York or Massachusetts.

in Panel B implies a return difference of $0.0049 \times 3.18 = 1.56\%$. This difference in the equity risk premium is similar in magnitude to the return spread from our hypothetical long-short trading strategy.

In unreported results, we increase the local bias threshold from 5% to 10% and then to 15%. Specifically, these thresholds require investors located in the same state as the firm to own at least 10% or 15% of its outstanding shares, and overweight the firm by 20% relative to its weight in the market portfolio. These more stringent local investor base thresholds reduce the sample to 318 firms in 39 states and 153 firms in 33 states, respectively. Furthermore, the $\beta_1$ coefficients are larger when local investment biases become more salient. For example, at the 15% threshold, the $\beta_1$ coefficient is 0.0096 ($t$-statistic of 2.35) for FP and 0.0083 ($t$-statistic of 2.96) for its predicted counterpart.

## 6.4 Headquarter Relocations and Local Bias Migrations

Relocations of firm headquarters provide a quasi-natural experiment to analyze the impact of fiscal policy on average stock returns. During the 1986 to 2006 period, 1,265 firms change the state in which they are headquartered according to Compact Disclosure data. We observe an even split between the number of firms relocating to states with more counter-cyclical and less counter-cyclical fiscal policies. Furthermore, these relocations occur across an array of industries.

We compute firm-level average returns before and after headquarter relocations. For firms that relocated their headquarters, these average return differences are then regressed on changes in the fiscal policy beta associated with a relocation of firm $k$'s headquarters from state $i$ to state $j$

$$\text{Return}_{k,j,t+} - \text{Return}_{k,i,t-} = \beta_0 + \beta_1 \left(\text{FP}_{k,j,t+} - \text{FP}_{k,i,t-}\right) + \gamma X + \epsilon_k \,, \tag{11}$$

where $t+$ refers to the firm's average annual return after the relocation and $t-$ refers to the firm's average annual return before the relocation in year $t$. The $X$ vector denotes state-level differences in GSP growth as well as its volatility before and after a headquarter relocation. Firms are required to have two years of returns before and after their headquarter relocation. This filter reduces our sample from 1,265 to 975. The first post-relocation return following a headquarter relocation is also removed from $\text{Return}_{k,j,t+}$ to further mitigate the possible influence of takeover premiums associated with mergers and acquisitions. Finally, we remove firms that relocated their headquarters multiple times during the sample. These filters create a final sample of 924 firm-level relocations. We also estimate equation (11) within a subset of firms that relocated their headquarters between states with above-median local investment biases. This local bias subset consists of 462 firm-level relocations. In a separate subset involving 462 headquarter relocations where at least one of the two states involved does not have an above-median local investment bias, the $\beta_1$ coefficients are smaller and less significant. This finding provides additional evidence regarding the important of local investment biases to the relationship between fiscal policy and stock returns.

According to Panel A of Table 7, the $\beta_1$ regression coefficient is positive without controls for state-level economic characteristics and with these control variables ($t$-statistics of 2.05 and 2.15, respectively). These positive coefficients indicate that firms relocating to states with more counter-cyclical fiscal policies subsequently have lower average stock returns.[10] For the subset of relocations between states with above-median local investment biases, the $\beta_1$ coefficients are larger in magnitude but less significant. The weaker significance may be attributable to the smaller sample size. Overall, the return implications of headquarter relocations confirm the importance of fiscal policy to average stock returns.

As emphasized earlier, local investment biases can arise from asymmetric information. In Van Nieuwerburgh and Veldkamp (2009), investors specialize in learning about local investments to differentiate their information from the information of other investors. Thus, when investors can choose what to learn, learning can exacerbate rather than mitigate informational asymmetries. Consistent with the predictions of Van Nieuwerburgh and Veldkamp (2009), Figure 1 indicates that local investment biases migrate with firm relocations. Specifically, when a firm relocates from state $i$ to state $j$, we observe that investors in state $i$ decrease their local bias towards the firm while investors in state $j$ develop a local bias. Similar evidence regarding return co-movement is reported in Pirinsky and Wang (2006).

Three portfolio weights are involved in the creation of Figure 1. The first represents a firm's weight in the market portfolio. The second and third weights are computed from the portfolios of institutional investors located in the same state as the firm's headquarters before and after the year $t$ relocation, respectively. For each firm, the means of all three firm-level weights are normalized to 1 over the five year $t-1$ to $t+3$ period. This normalization accounts for the tendency of active fund managers to overweight a small number of stocks. We then normalize the pre-relocation and post-relocation portfolio weights of institutional investors by the respective market portfolio weights of firms. Thus, values of the firm-level ratio that

[10]To clarify, equation (11) does not require headquarter relocations to occur between states whose fiscal policy betas differ in sign. For example, firms can relocate between counter-cyclical states, with the difference in the fiscal policy beta then reflecting whether the firm has relocated to a more or less counter-cyclical state.

exceed 1 indicate that investors have a local bias towards the firm. For investors located in both the pre-relocation state (state $i$) and post-relocation state (state $j$), the average ratios are plotted over the five-year horizon.

## 6.5 Fiscal Policy and Mutual Fund Flows

The discount rate channel predicts that the buy/sell decisions of investors are less sensitive to economic conditions in states with counter-cyclical fiscal policies. Specifically, during poor economic conditions, investors in counter-cyclical states are less prone to sell stocks than investors in pro-cyclical states. To test this prediction, we compute the percentage flow (FLOW) of mutual fund assets in each state every year starting from 1991. Mutual fund locations are identified by the state in which they are headquartered using the data in Hong, Kubik, and Stein (2005).[11] At the state-level, a time series regression of state-level mutual fund inflows and outflows on contemporaneous GSP growth is performed

$$\mathrm{FLOW}_{i,t} = \beta_{i,0} + \beta_{i,1}\,\mathrm{GSP}_{i,t} + \epsilon_{i,t} \tag{12}$$

for 9 of the 17 states with above-median local investment biases and sufficient mutual fund data. The $\beta_{i,1}$ coefficient is the designated flow beta. A small flow beta indicates that investors with local biases are less likely to sell stocks during poor economic conditions.

Flow betas are reported in Panel B of Table 7 along with the original fiscal policy betas from Table 2. The 0.395 correlation between these betas indicates that for mutual funds in counter-cyclical states (low FP beta), flows into mutual funds are less sensitive to state-level economic conditions (low flow beta). Thus, in the presence of local investment biases, counter-cyclical fiscal policies reduce the amount of local stocks sold during poor economic conditions.

Overall, the flow betas are consistent with the economic intuition underlying the discount rate channel. In particular, counter-cyclical fiscal policies enable government budget deficits to partially finance the consumption of investors in poor economic conditions. Thus, counter-cyclical fiscal policies reduce the need for investors to sell stocks during poor economic conditions. As a consequence, the systematic risk associated with stock ownership is lowered along with the equity premium.

[11] We are grateful to the authors for providing us with their data.

## 6.6 Cash Flow Channel

Counter-cyclical fiscal policies have the potential to lower firm-level cash flow volatility by reducing consumption volatility. To examine this cash flow channel, we apply our instrumental variables analysis to cash flow volatility.

State-level cash flow volatility is computed as the average cash flow volatility of all firms headquartered in each state. The underlying firm-level cash flow volatility equals the standard deviation of a firm's annual cash flow across the 1965 to 2008 sample period. As in Korniotis and Kumar (2013), cash flows are defined as earnings before extraordinary items minus total accruals, scaled by total assets. This instrumental variables procedure is applied to the 35 states with at least 20 firms. The first stage of the instrumental variables procedure proceeds identically to equation (4) while the second stage replaces consumption volatility with cash flow volatility

$$\text{Cash Flow Volatility}_i \quad = \quad \gamma_0 + \gamma_1 \, \hat{\text{FP}}_i + \epsilon_i \,, \tag{13}$$

where the predicted fiscal policy betas $\hat{\text{FP}}$ are defined by the $\alpha$ estimates from the first stage. The sign and interpretation of the $\alpha$ coefficients in Table 8 are identical to their counterparts in Table 4. Furthermore, a positive $\gamma_1$ coefficient in equation (13) indicates that cash flow volatility is lower in counter-cyclical states. However, Panel A of Table 8 does not provide convincing evidence that cash flow volatility is lower in counter-cyclical states since the $\gamma_1$ coefficient is insignificant ($t$-statistic of 1.27).

The dataset of Garcia and Norli (2012) enables us to examine the operations of individual firms across states. This data, which begins in 1994, records instances where state names occur in firm-level 10-K reports. Delaware and Washington are removed from the sample given the large number of firms incorporated in Delaware and the possibility that Washington refers to the US capital rather than the state. On average, firms operate in 7.9 states with the median being 5.5 states.

We compute firm-level cash flow fiscal policy betas denoted $\text{CF-FP}_k$ that weight the state-level fiscal policy betas in which a firm operates by the frequency of each state's name in the firm's 10-K report. The cash flow betas based on fiscal policy enable us to examine the cash flow channel using the following Fama-MacBeth regression starting in 1994

$$\text{Firm Return}_k \quad = \quad \beta_0 + \beta_1 \, \text{CF-FP}_k + \beta_2 \, \text{BM}_k + \beta_3 \, \text{SIZE}_k + \beta_4 \, \text{PRET}_k + \epsilon_k \,. \tag{14}$$

As this Fama-MacBeth regression uses firm-level CF-FP betas, the control variables are firm-level book-to-market (BM), size (SIZE), and past return (PRET) characteristics. PRET is defined as the firm's return over the prior twelve months after omitting the most recent month. Unlike previous regression specifications, the firm-level returns that comprise the dependent variable in equation (14) are not risk-adjusted due to the inclusion of these firm-level control variables.

An insignificant $\beta_1$ coefficient is evidence that the cash flow channel is not responsible for the return implications of fiscal policy. According to Panel B of Table 8, the $\beta_1$ coefficients are insignificant with and without firm-level controls. Indeed, their $t$-statistics range from 0.56 to 0.69 in various specifications. The R-squared measures in Panel B are also considerably lower than their counterparts in Table 5 that pertain to the discount rate channel.

Nonetheless, the lack of empirical support for the cash flow channel may be attributable to firms that operate across many states. Indeed, the cash flows of firms with diversified operations are unlikely to be affected by the fiscal policy of any individual state. Therefore, we restrict our next analysis to geographically concentrated firms whose operations are limited to a maximum of three states.

The results for geographically concentrated firms are reported in Table 9, with Panel A replicating the instrumental variables procedure in equation (refsecondcf). On average, 440 firms per month are geographically concentrated. Within this subset, the $t$-statistic for $\gamma_1$ increases to 1.76 from its earlier value of 1.27. Furthermore, the results in Panel B of Table 9 indicate that the firm-level cash flow betas are significant for geographically concentrated firms. In particular, the $t$-statistic for CF-FP equals 2.19 without controlling for book-to-market, size, and past return characteristics. This $t$-statistic declines slightly to 1.97 after the introduction of these firm-level control variables. Overall, for geographically concentrated firms, Panel B of Table 9 provides evidence consistent with the cash flow channel.

In unreported results, the CF-FP betas continue to be priced when the definition of geographically concentrated is relaxed to include at most five states. On average, 812 firms per month comprise this extended subset, with the CF-FP coefficients all having $t$-statistics above 2.50. However, once the definition of geographically concentrated expands to include firms that operate in as many as ten states, the $\beta_1$ coefficients are no longer significant in any specification.

Overall, the lower average returns associated with counter-cyclical fiscal policies can be partially attributed to the cash flow channel, but only for firms whose operations are geographically concentrated in a small number of states with similar fiscal policies.

### 6.7 State Demographics

According to the consumption-CAPM, an asset's expected return is a function of consumption volatility, an asset's consumption beta, and investor risk aversion. We find strong evidence that counter-cyclical fiscal policies lower consumption volatility, and weaker evidence that counter-cyclical fiscal policies lower firm-level cash flow betas. Furthermore, investor risk aversion does not differ significantly across states with counter-cyclical versus pro-cyclical fiscal policies.

Nonetheless, our fiscal policy betas may capture an omitted state characteristic. For example, counter-cyclical states may have more residents employed in knowledge-based industries whose education coincides with greater participation in the stock market (Campbell, 2006), and a lower equity premium as a consequence. Therefore, we regress our state-level fiscal policy betas on the amount of university R&D conducted in each state normalized by GSP as well as the proportion of the state's population with university degrees and the number of patents per capita. State-level university R&D and educational attainment data are obtained in 2007 while patent data is from 2010.

The following cross-sectional regression examines the impact of university R&D, per capita patents (PATENT), and per capita attainment of university degrees (UNI) on government fiscal policy

$$\mathrm{FP}_i \;=\; \beta_0 + \beta_1\,\mathrm{R\&D}_i + \beta_2\,\mathrm{PATENT}_i + \beta_3\,\mathrm{UNI}_i + \epsilon_i\,. \tag{15}$$

Negative coefficients for R&D, PATENT, and UNI indicate that counter-cyclical fiscal policies are more prevalent in states whose residents are more educated and more likely to be employed in a knowledge-based industry.

Table 10 reports that our fiscal policy betas are not capturing an omitted variable related to education or employment in knowledge-based industries. Indeed, all three independent variables (R&D, PATENT, and UNI) have insignificant $\beta$ coefficients.

# 7 Conclusions

Using state-level data within the United States, we find that consumption volatility is lower in states that implement counter-cyclical fiscal policies. Furthermore, firms headquartered in counter-cyclical states have lower average stock returns provided their investor base has a local investment bias. Variation in state-level fiscal policy induces return variation between

1.5% and 2.0% per year after controlling for common risk factors and industry differences across states. This return variation can be attributed to differences in consumption volatility.

The implications of fiscal policy for consumption risk and stock returns are confirmed using an instrumental variables analysis based on state-level balanced budget amendments and the adoption of budget stabilization funds. Moreover, average stock returns decrease following firm relocations to states with a more counter-cyclical fiscal policy. Equity trades by investors in counter-cyclical states are also less sensitive to economic conditions than the trades of investors in pro-cyclical states.

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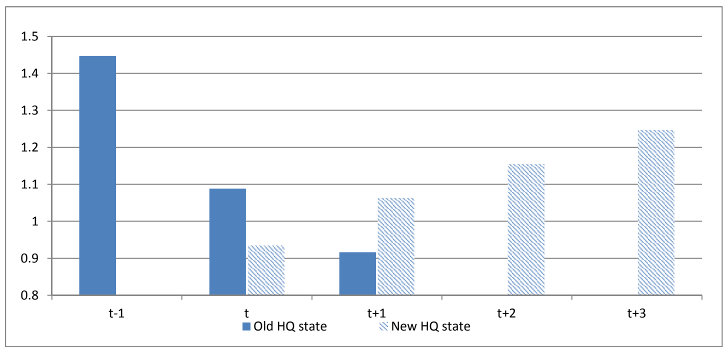


**Figure 1** This figure provides a visual illustration of the migrations in local investment biases that accompany firm headquarter relocations across states in year $t$. Three portfolio weights are involved in this analysis. First, each firm's weight in the market portfolio is computed. The second and third weights are computed from the portfolios of institutional investors located in the same state as the firm's headquarters before and after the relocation, respectively. For each firm, the means of all three firm-level weights are normalized to 1 over the five year $t-1$ to $t+3$ period. This normalization accounts for the tendency of active fund managers to overweight a small number of stocks. We then normalize the pre-relocation state and post-relocation state portfolio weights in each firm by the firm's weight in the market portfolio. Values of this firm-level ratio in excess of 1 indicate that investors have a local bias towards the firm.

# Table 1: Summary statistics

Panel A of this table summarizes the mean, median, standard deviation (Std. Dev.), minimum, maximum, $10^{th}$ and $90^{th}$ percentiles, and autocorrelation of the important variables in our study. Return represents that average value-weighted return of firms headquartered in the 35 states that have at least 20 firms. Summary statistics are also reported for both consumption proxies and their time series volatility across all 50 states. These consumption proxies are residential and commercial electricity usage as well as retail sales. Summary statistics for the per capita annual growth in gross state product (GSP) as well as its volatility are included in Panel A along with the volatility of annual per capita income growth. Panel B reports cross-sectional correlations between several variables in Panel A. These correlations are defined using the time series averages of each variable. In contrast, the correlations in Panel C represent averages of within state time series correlations.

Panel A: Summary statistics

| | N | Average | Std. Dev. | Minimum | 10th percentile | Median | 90th percentile | Maximum | Average autocorrelation |
|---|---|---|---|---|---|---|---|---|---|
| Average return | 35 | 0.112 | 0.021 | 0.078 | 0.088 | 0.107 | 0.129 | 0.192 | 0.036 |
| Electricity growth | 50 | 0.032 | 0.007 | 0.011 | 0.023 | 0.033 | 0.040 | 0.045 | 0.201 |
| Retail Sales growth | 50 | 0.065 | 0.010 | 0.051 | 0.053 | 0.064 | 0.078 | 0.095 | 0.514 |
| GSP growth | 50 | 0.075 | 0.010 | 0.057 | 0.064 | 0.075 | 0.088 | 0.101 | 0.469 |
| Electricity volatility | 50 | 0.049 | 0.013 | 0.029 | 0.036 | 0.047 | 0.067 | 0.085 | - |
| Retail Sales volatility | 50 | 0.048 | 0.009 | 0.036 | 0.039 | 0.047 | 0.055 | 0.096 | - |
| Income volatility | 50 | 0.034 | 0.010 | 0.027 | 0.027 | 0.031 | 0.039 | 0.090 | - |
| GSP volatility | 50 | 0.041 | 0.019 | 0.022 | 0.028 | 0.036 | 0.053 | 0.138 | - |

Panel B: Cross-sectional correlations based on time series averages for each state

| | Average return | Electricity volatility | Retail Sales volatility | GSP volatility |
|---|---|---|---|---|
| Average return | 1 | | | |
| Electricity volatility | 0.272 | 1 | | |
| Retail Sales volatility | 0.151 | 0.365 | 1 | |
| GSP volatility | -0.067 | 0.255 | 0.491 | 1 |

Panel C: Within state time-series correlations

| | Average return | Electricity growth | Retail growth | GSP growth |
|---|---|---|---|---|
| Average return | 1 | | | |
| Electricity growth | -0.027 | 1 | | |
| Retail growth | 0.081 | 0.253 | 1 | |
| GSP growth | 0.041 | 0.191 | 0.642 | 1 |

# Table 2: State-level variables

Panel A of this table reports on the state-level fiscal policy betas (FP) in equation (2) that measure the sensitivity of annual budget deficits to economic conditions. The ACIR index for each state represents the stringency of its balanced budget amendment on a scale of 0 (weak) to 10 (strong). The year in which each state's stabilization fund was adopted, if at all is also reported. The fraction of our 1965-2008 sample period for which a state did not have a stabilization fund is denoted FNO (fraction of no stabilization fund) and defined as $\frac{t-1965}{2008-1965}$ if state $i$ adopted a stabilization fund in year $t$, 0 if the adoption occurred before 1965, and 1 if the state had no stabilization fund during the sample period. The average state-level return variable represents the value-weighted average stock return of firms headquartered in each of the 35 states that have at least 20 firms. The state-level averages for both consumption proxies based on electricity consumption and retail sales are also reported along with their volatility as well as the growth and volatility of each state's gross state product (GSP). The local investment bias measure equals the percentage of local stocks in a particular state held by institutional investors minus the percentage of local stocks in the market portfolio. These deviations from the market portfolio are value-weighted according to the market value of each institutional investor's portfolio from 1980. The last column records whether the local investment bias for a state is above or below the cross-sectional median. Panel B orders states in ascending order according to their fiscal policy beta. In addition, the volatility and autocorrelation of each state's tax revenues as well as expenditures are reported.

Panel A: State-level fiscal policy betas and summary statistics

| State | FP | ACIR | Stab. Fund | FNO | Average return | Elect. growth | Retail growth | GSP growth | Elect. Vol | Retail Vol | GSP Vol | Local Bias | Relation to Median |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| AK | 8.509 | 6 | - | 1.000 | - | 0.038 | 0.079 | 1.097 | 0.080 | 0.096 | 0.138 | - | - |
| AL | 0.346 | 10 | 1927 | 0.000 | 0.116 | 0.036 | 0.067 | 1.072 | 0.049 | 0.043 | 0.033 | 0.131 | above |
| AR | 2.946 | 9 | - | 1.000 | 0.192 | 0.040 | 0.066 | 1.075 | 0.063 | 0.057 | 0.036 | 0.140 | above |
| AZ | -1.368 | 10 | 1990 | 0.581 | 0.101 | 0.027 | 0.086 | 1.094 | 0.043 | 0.054 | 0.046 | 0.023 | below |
| CA | 1.094 | 6 | 1976 | 0.256 | 0.124 | 0.019 | 0.064 | 1.077 | 0.039 | 0.045 | 0.034 | 0.025 | below |
| CO | 2.673 | 10 | 1982 | 0.395 | 0.078 | 0.028 | 0.072 | 1.087 | 0.045 | 0.054 | 0.040 | 0.010 | below |
| CT | -0.625 | 5 | 1979 | 0.326 | 0.113 | 0.031 | 0.058 | 1.071 | 0.041 | 0.047 | 0.034 | 0.014 | below |
| DE | -1.639 | 10 | 1979 | 0.326 | - | 0.037 | 0.066 | 1.078 | 0.057 | 0.045 | 0.033 | - | - |
| FL | 0.507 | 10 | 1959 | 0.000 | 0.088 | 0.031 | 0.081 | 1.091 | 0.047 | 0.053 | 0.039 | 0.012 | below |
| GA | 0.226 | 10 | 1976 | 0.256 | 0.126 | 0.036 | 0.075 | 1.084 | 0.049 | 0.049 | 0.035 | 0.081 | above |
| HI | -2.493 | 10 | - | 1.000 | - | 0.026 | 0.073 | 1.077 | 0.033 | 0.069 | 0.041 | - | - |
| IA | -1.049 | 10 | 1984 | 0.442 | 0.111 | 0.029 | 0.052 | 1.067 | 0.044 | 0.036 | 0.043 | 0.067 | above |
| ID | -0.406 | 10 | 1984 | 0.442 | - | 0.017 | 0.070 | 1.080 | 0.061 | 0.045 | 0.044 | - | - |
| IL | 1.001 | 4 | 2001 | 0.837 | 0.107 | 0.027 | 0.051 | 1.063 | 0.042 | 0.043 | 0.026 | 0.056 | above |
| IN | 0.452 | 10 | 1982 | 0.395 | 0.103 | 0.033 | 0.054 | 1.064 | 0.041 | 0.044 | 0.036 | 0.507 | above |
| KS | -0.273 | 10 | 1993 | 0.651 | 0.087 | 0.033 | 0.057 | 1.070 | 0.043 | 0.038 | 0.033 | 0.003 | below |
| KY | 2.080 | 10 | 1983 | 0.419 | 0.122 | 0.044 | 0.061 | 1.067 | 0.050 | 0.044 | 0.034 | 0.096 | above |
| LA | 0.570 | 4 | 1990 | 0.581 | 0.097 | 0.037 | 0.064 | 1.076 | 0.054 | 0.052 | 0.064 | 0.053 | above |
| MA | 0.796 | 3 | 1985 | 0.465 | 0.105 | 0.034 | 0.054 | 1.071 | 0.040 | 0.056 | 0.032 | 0.012 | below |
| MD | -1.444 | 6 | 1985 | 0.465 | 0.097 | 0.040 | 0.063 | 1.075 | 0.067 | 0.047 | 0.028 | 0.011 | below |
| ME | 3.688 | 9 | 1985 | 0.465 | - | 0.028 | 0.066 | 1.069 | 0.036 | 0.047 | 0.033 | - | - |
| MI | -1.518 | 6 | 1977 | 0.279 | 0.088 | 0.029 | 0.053 | 1.057 | 0.039 | 0.047 | 0.047 | 0.079 | above |
| MN | -1.468 | 8 | 1981 | 0.372 | 0.119 | 0.035 | 0.061 | 1.074 | 0.051 | 0.044 | 0.034 | 0.133 | above |
| MO | -0.077 | 10 | 1992 | 0.628 | 0.122 | 0.040 | 0.056 | 1.066 | 0.047 | 0.043 | 0.030 | 0.029 | below |
| MS | 1.376 | 9 | 1982 | 0.395 | - | 0.036 | 0.064 | 1.071 | 0.048 | 0.046 | 0.037 | - | - |
| MT | 0.758 | 10 | - | 1.000 | - | 0.031 | 0.058 | 1.068 | 0.085 | 0.046 | 0.044 | - | - |
| NC | 0.396 | 10 | 1991 | 0.605 | 0.107 | 0.037 | 0.073 | 1.079 | 0.048 | 0.049 | 0.030 | 0.078 | above |
| ND | 0.620 | 8 | 1987 | 0.512 | - | 0.045 | 0.056 | 1.073 | 0.059 | 0.039 | 0.088 | - | - |
| NE | 1.019 | 10 | 1983 | 0.419 | - | 0.030 | 0.054 | 1.070 | 0.050 | 0.036 | 0.039 | - | - |
| NH | 1.762 | 2 | 1987 | 0.512 | 0.120 | 0.032 | 0.077 | 1.084 | 0.046 | 0.054 | 0.042 | 0.000 | below |
| NJ | 0.482 | 10 | 1990 | 0.581 | 0.104 | 0.035 | 0.061 | 1.069 | 0.036 | 0.040 | 0.028 | 0.036 | below |
| NM | 3.805 | 10 | 1966 | 0.023 | - | 0.026 | 0.069 | 1.078 | 0.052 | 0.048 | 0.052 | - | - |
| NV | -1.606 | 4 | 1994 | 0.674 | 0.121 | 0.011 | 0.095 | 1.101 | 0.067 | 0.053 | 0.040 | -0.001 | below |
| NY | -4.183 | 3 | 1945 | 0.000 | 0.104 | 0.026 | 0.051 | 1.064 | 0.029 | 0.038 | 0.024 | 0.009 | below |
| OH | -0.672 | 10 | 1981 | 0.372 | 0.104 | 0.032 | 0.053 | 1.060 | 0.037 | 0.042 | 0.032 | 0.098 | above |
| OK | 3.359 | 10 | 1986 | 0.488 | 0.120 | 0.036 | 0.063 | 1.075 | 0.059 | 0.052 | 0.055 | 0.040 | below |
| OR | -4.044 | 8 | 1995 | 0.698 | 0.129 | 0.013 | 0.061 | 1.078 | 0.036 | 0.052 | 0.040 | 0.009 | below |
| PA | -0.765 | 6 | 1985 | 0.465 | 0.101 | 0.034 | 0.053 | 1.063 | 0.030 | 0.039 | 0.022 | 0.055 | above |
| RI | -1.804 | 10 | 1985 | 0.465 | - | 0.036 | 0.055 | 1.066 | 0.057 | 0.053 | 0.030 | - | - |
| SC | -2.422 | 10 | 1978 | 0.302 | 0.083 | 0.033 | 0.074 | 1.079 | 0.046 | 0.047 | 0.036 | 0.059 | above |
| SD | 3.087 | 10 | 1991 | 0.605 | - | 0.033 | 0.058 | 1.075 | 0.042 | 0.041 | 0.052 | - | - |
| TN | -0.342 | 10 | 1972 | 0.163 | 0.107 | 0.027 | 0.072 | 1.077 | 0.080 | 0.047 | 0.033 | 0.065 | above |
| TX | -0.523 | 8 | 1987 | 0.512 | 0.110 | 0.030 | 0.074 | 1.088 | 0.047 | 0.049 | 0.048 | 0.037 | below |
| UT | -2.372 | 10 | 1986 | 0.488 | 0.099 | 0.027 | 0.078 | 1.084 | 0.037 | 0.052 | 0.038 | 0.071 | above |
| VA | -0.892 | 8 | 1992 | 0.628 | 0.122 | 0.037 | 0.070 | 1.081 | 0.043 | 0.045 | 0.029 | 0.113 | above |
| VT | -0.028 | 0 | 1988 | 0.535 | - | 0.030 | 0.068 | 1.075 | 0.073 | 0.045 | 0.040 | - | - |
| WA | -1.222 | 8 | 1981 | 0.372 | 0.158 | 0.021 | 0.071 | 1.080 | 0.053 | 0.045 | 0.036 | 0.047 | below |
| WI | -0.073 | 6 | 1985 | 0.465 | 0.133 | 0.029 | 0.056 | 1.067 | 0.040 | 0.041 | 0.029 | 0.029 | below |
| WV | -0.272 | 10 | 1994 | 0.674 | - | 0.038 | 0.055 | 1.059 | 0.034 | 0.045 | 0.034 | - | - |
| WY | -4.004 | 8 | 1982 | 0.395 | - | 0.043 | 0.067 | 1.081 | 0.067 | 0.057 | 0.085 | - | - |
| Average | 0.079 | 8 | 1982 | 0.479 | 0.112 | 0.032 | 0.065 | 1.075 | 0.049 | 0.048 | 0.041 | 0.064 | NA |

Panel B: State-level fiscal policy betas in ascending order and summary statistics for deficit components

| | | volatilities | | autocorrelations | |
|---|---|---|---|---|---|
| State | FP | revenue / GSP | expenditures / GSP | revenue | expenditures |
| NY | -4.183 | 0.0583 | 0.0420 | 0.9413 | 0.9629 |
| OR | -4.044 | 0.0856 | 0.0396 | 0.8023 | 0.9051 |
| WY | -4.004 | 0.1053 | 0.0671 | 0.8934 | 0.8989 |
| HI | -2.493 | 0.0594 | 0.0635 | 0.8924 | 0.8886 |
| SC | -2.422 | 0.0524 | 0.0522 | 0.9708 | 0.9692 |
| UT | -2.372 | 0.0525 | 0.0505 | 0.9303 | 0.9296 |
| RI | -1.804 | 0.0689 | 0.0575 | 0.9141 | 0.9224 |
| DE | -1.639 | 0.0574 | 0.0551 | 0.9033 | 0.7332 |
| NV | -1.606 | 0.0544 | 0.0487 | 0.6982 | 0.5809 |
| MI | -1.518 | 0.0740 | 0.0481 | 0.8925 | 0.9238 |
| MN | -1.468 | 0.0548 | 0.0402 | 0.9143 | 0.9022 |
| MD | -1.444 | 0.0663 | 0.0469 | 0.8875 | 0.9248 |
| AZ | -1.368 | 0.0520 | 0.0453 | 0.8448 | 0.8443 |
| WA | -1.222 | 0.0507 | 0.0675 | 0.8753 | 0.7460 |
| IA | -1.049 | 0.0449 | 0.0420 | 0.9159 | 0.8776 |
| VA | -0.892 | 0.0558 | 0.0390 | 0.8740 | 0.8926 |
| PA | -0.765 | 0.0552 | 0.0454 | 0.9522 | 0.9612 |
| OH | -0.672 | 0.0537 | 0.0386 | 0.9650 | 0.9815 |
| CT | -0.625 | 0.0608 | 0.0461 | 0.8984 | 0.9063 |
| TX | -0.523 | 0.0435 | 0.0387 | 0.9600 | 0.9406 |
| ID | -0.406 | 0.0512 | 0.0374 | 0.9237 | 0.9210 |
| TN | -0.342 | 0.0465 | 0.0473 | 0.9437 | 0.8974 |
| KS | -0.273 | 0.0456 | 0.0389 | 0.9331 | 0.9422 |
| WV | -0.272 | 0.0506 | 0.0475 | 0.9671 | 0.9660 |
| MO | -0.077 | 0.0520 | 0.0407 | 0.9435 | 0.9571 |
| WI | -0.073 | 0.0874 | 0.0370 | 0.8357 | 0.9482 |
| VT | -0.028 | 0.0452 | 0.0511 | 0.9183 | 0.8423 |
| GA | 0.226 | 0.0443 | 0.0458 | 0.9294 | 0.8924 |
| AL | 0.346 | 0.0640 | 0.0383 | 0.8847 | 0.9542 |
| NC | 0.396 | 0.0436 | 0.0491 | 0.9648 | 0.9468 |
| IN | 0.452 | 0.0468 | 0.0413 | 0.9625 | 0.9647 |
| NJ | 0.482 | 0.0508 | 0.0527 | 0.9565 | 0.9506 |
| FL | 0.507 | 0.0524 | 0.0456 | 0.9518 | 0.9458 |
| LA | 0.570 | 0.0546 | 0.0525 | 0.9257 | 0.9206 |
| ND | 0.620 | 0.0650 | 0.0464 | 0.8093 | 0.8362 |
| MT | 0.758 | 0.0524 | 0.0428 | 0.9371 | 0.9428 |
| MA | 0.796 | 0.0446 | 0.0471 | 0.9316 | 0.9248 |
| IL | 1.001 | 0.0563 | 0.0469 | 0.9433 | 0.9571 |
| NE | 1.019 | 0.0437 | 0.0466 | 0.9434 | 0.8873 |
| CA | 1.094 | 0.0694 | 0.0382 | 0.8103 | 0.9248 |
| MS | 1.376 | 0.0641 | 0.0401 | 0.8905 | 0.9682 |
| NH | 1.762 | 0.0438 | 0.0638 | 0.8617 | 0.7925 |
| KY | 2.080 | 0.0516 | 0.0457 | 0.9554 | 0.9632 |
| CO | 2.673 | 0.0446 | 0.0450 | 0.8662 | 0.8202 |
| AR | 2.946 | 0.0582 | 0.0538 | 0.9463 | 0.9332 |
| SD | 3.087 | 0.0688 | 0.0452 | 0.2506 | 0.4971 |
| OK | 3.359 | 0.0542 | 0.0410 | 0.9384 | 0.9416 |
| ME | 3.688 | 0.0612 | 0.0426 | 0.9475 | 0.9569 |
| NM | 3.805 | 0.0703 | 0.0438 | 0.8352 | 0.9438 |
| AK | 8.509 | 0.4767 | 0.0916 | 0.4985 | 0.8393 |
| Average | 0.079 | 0.0653 | 0.0476 | 0.8866 | 0.8974 |
| Correlation with FP | | 0.4910 | 0.2180 | -0.3679 | -0.0791 |

## Table 3: Borrowing constraints and consumption

The table reports on the implications of borrowing constraints arising from bank deregulations on consumption growth during bad economic conditions. A panel regression estimates the following specification in equation (3), Consumption Growth$_{i,t} = \beta_0 + \beta_1 1_{\text{Bad}_{i,t}} + \beta_2 1_{\text{BC}^j_{i,t}} + \beta_3 \left(1_{\text{Bad}_{i,t}} \times 1_{\text{BC}^j_{i,t}}\right) + \gamma X_i + \epsilon_{i,t}$. State fixed effects are included in the $X$ vector while the standard errors are clustered at the state-level. The dummy variable $1_{\text{Bad}}$ equals 1 in years where state $i$'s GSP growth is in its bottom quintile (GSP$^Q_i$ equals 1 for state $i$). The BC$^j_{i,t}$ dummy variable equals 1 in any year $t$ before the relaxation of each of the three banking restrictions ($j$=1,2,3) in state $i$. The first deregulation ($j = 1$) allowed mergers and acquisitions to facilitate intrastate branching, the second ($j = 2$) permitted intrastate branching, and the third ($j = 3$) permitted interstate banking. Table 1 of Kroszner and Strahan (1999) provides the year in which each of these banking deregulations were enacted for each state. $t$-statistics are reported in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively. These statistics have standard errors clustered at the state-level.

| | Electricity | | | Retail Sales | | |
|---|---|---|---|---|---|---|
| | M&A Allowed | Intrastate Allowed | Interstate Allowed | M&A Allowed | Intrastate Allowed | Interstate Allowed |
| Interaction | -0.0124** | -0.0145** | -0.0156*** | -0.0153** | -0.0116* | -0.0237*** |
| | *-2.04* | *-2.50* | *-3.11* | *-2.23* | *-1.86* | *-5.07* |
| BC Dummy | 0.0365*** | 0.0335*** | 0.0311*** | 0.0387*** | 0.0408*** | 0.0473*** |
| | *15.15* | *13.70* | *21.74* | *15.77* | *15.73* | *23.03* |
| Bad Economy Dummy | -0.0076*** | -0.0063** | -0.0051* | -0.0364*** | -0.0350*** | -0.0279*** |
| | *-2.94* | *-2.28* | *-1.82* | *-9.27* | *-7.85* | *-8.17* |
| Observations | 2,244 | 2,244 | 2,244 | 2,244 | 2,244 | 2,244 |
| R-squared | 0.131 | 0.110 | 0.116 | 0.307 | 0.311 | 0.383 |

## Table 4: Fiscal policy and consumption volatility

The table reports the results from an instrumental variables procedure based on the ACIR index as well as the presence of state-level stabilization funds. The ACIR index ranges from 0 to 10, with higher values corresponding to more stringent balanced budget restrictions on a state's fiscal policy. To capture the impact of balanced budget amendments, $\text{DACIR}_i$ equals 0 when the ACIR index for state $i$ is three or below, 1 when this index is between four and seven, and 2 when this index is eight or above. The fraction of our 1965-2008 sample period for which a state did not have a stabilization fund is denoted FNO (fraction of no stabilization fund) and defined as $\frac{t-1965}{2008-1965}$ if state $i$ adopted a stabilization fund in year $t$, 0 if the adoption occurred before 1965, and 1 if the state had no stabilization fund during the sample period. The state-level fiscal policy betas (FP) are measured according to equation (2) as the sensitivity of annual budget deficits to economic conditions. In the first stage, predicted values for the fiscal policy betas denoted $\hat{\text{FP}}$ are obtained from the cross-sectional regression in equation (4), $\text{FP}_i = \alpha_0 + \alpha_1 \, \text{DACIR}_i + \alpha_2 \, \text{FNO}_i + \alpha_3 \, ( \, \text{DACIR}_i \; \times \; \text{FNO}_i \, ) + \epsilon_i$. In the second stage, the volatility of state-level consumption is regressed on these predicted fiscal policy betas from the first stage, as in the cross-sectional regression specified in equation (5), $\text{Consumption Volatility}_i = \gamma_0 + \gamma_1 \, \hat{\text{FP}}_i + \epsilon_i$. The predicted values are defined by the $\alpha$ estimates from equation (4) as $\hat{\text{FP}}_i = \hat{\alpha}_0 + \hat{\alpha}_1 \, \text{DACIR}_i + \hat{\alpha}_2 \, \text{FNO}_i + \hat{\alpha}_3 \, ( \, \text{DACIR}_i \; \times \; \text{FNO}_i \, )$ for state $i$. The $\gamma_1$ coefficients have been multiplied below by 100 for ease of interpretation. $t$-statistics are reported in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively.

Panel A: Instrumental variables regression for 50 states

| | | Electricity volatility | Retail Sales volatility |
|---|---|---|---|
| | first stage | second stage | second stage |
| Interaction (DACIR x FNO) | -6.4037*** | | |
| | *-3.28* | | |
| DACIR | 2.7338*** | | |
| | *2.84* | | |
| FNO | 12.4005*** | | |
| | *3.64* | | |
| Predicted FP | | 0.5227*** | 0.3755*** |
| | | *4.51* | *3.08* |
| Constant | -5.2815*** | 4.8853*** | 4.7780*** |
| | *-3.14* | *23.55* | *41.31* |
| Number of Observations | 50 | 50 | 50 |
| R-squared | 0.225 | 0.151 | 0.188 |

Panel B: Instrumental variables regression for 35 states with return data

| | | Electricity volatility | Retail Sales volatility |
|---|---|---|---|
| | first stage | second stage | second stage |
| Interaction (DACIR x FNO) | -3.7056* | | |
| | *-1.97* | | |
| DACIR | 1.6249* | | |
| | *1.90* | | |
| FNO | 7.3853** | | |
| | *2.29* | | |
| Predicted FP | | 0.6373** | 0.1934** |
| | | *2.48* | *2.39* |
| Constant | -265.2709*** | 4.7986*** | 4.7254*** |
| | *-5.45* | *22.36* | *48.41* |
| Number of Observations | 35 | 35 | 35 |
| R-squared | 0.152 | 0.097 | 0.059 |

## Table 5: State-level regressions

The table reports the coefficients from Fama-MacBeth regressions that examine the stock return implications of fiscal policy after classifying firms into states depending on the location of their headquarters. The results are based on the following specification for the value-weighted average returns of firms in state $i$, $\text{State Return}_i = \beta_0 + \beta_1\,(\text{FP}_i \times \text{LB}_i) + \beta_2\,\text{FP}_i + \beta_3\,\text{GSP}_i + \beta_4\,\text{GSP volatility}_i + \epsilon_i$. Panel A and Panel C examine 35 states with at least 20 firms using the interaction between state-level local biases denoted LB and the state-level fiscal policy betas denoted FP. Panel B and Panel D focus on a subset of 17 states where local investment biases are above the median using the following regression specification, $\text{State Return}_i = \beta_0 + \beta_1\,\text{FP}_i + \beta_2\,\text{GSP}_i + \beta_3\,\text{GSP volatility}_i + \epsilon_i$. GSP growth represents the average per capita annual growth rate in a state's gross state product while GSP volatility represents the standard deviation of this growth rate over the sample period. The regression specifications in the first two panels are based on returns that are risk-adjusted using the Fama-French four-factor procedure from 1965 to 1975 and the DGTW procedure from 1976 to 2008. The later two panels examine industry-adjusted returns. For all firms headquartered in a particular state, state-level returns are then formed by value-weighting these firm-level returns. The local investment bias measure equals the percentage of local stocks in a particular state held by institutional investors minus the percentage of local stocks in the market portfolio. These deviations from the market portfolio are value-weighted according to the market value of each institutional investor's portfolio from 1980. Therefore, we estimate the Fama-MacBeth regressions starting in 1965 as well as 1980 until 2008. The state-level fiscal policy betas are measured according to equation (2) as the sensitivity of annual budget deficits to economic conditions. Predicted fiscal policy betas are defined by the first stage of our instrumental variables procedure in equation (4) that predicts state-level fiscal policy betas using balanced budget amendments and budget stabilization funds defined by the DACIR and FNO variables in Table 2. $t$-statistics are in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively.

Panel A: Fama-MacBeth regressions using DGTW-adjusted state-level returns for 35 states

| | 1965-2008 | | | | 1980-2008 | | | |
|---|---|---|---|---|---|---|---|---|
| Interaction of FP and LB | | 0.0604** | 0.0925*** | 0.0880*** | | 0.0694** | 0.0962*** | 0.0891** |
| | | *2.30* | *3.13* | *2.75* | | *2.26* | *2.93* | *2.48* |
| FP | 0.0010 | | -0.0031 | -0.0020 | 0.0017 | | -0.0026 | -0.0010 |
| | *0.49* | | *-1.31* | *-0.84* | *0.71* | | *-0.95* | *-0.36* |
| GSP growth | | | | 0.2052 | | | | 0.3409 |
| | | | | *1.20* | | | | *1.49* |
| GSP volatility | | | | -0.5721 | | | | -0.9828** |
| | | | | *-1.42* | | | | *-2.10* |
| Constant | -0.0014 | -0.0018 | -0.0026 | -0.1989 | -0.0032 | -0.0038 | -0.0045 | -0.3251 |
| | *-0.52* | *-0.64* | *-0.99* | *-1.09* | *-1.02* | *-1.15* | *-1.46* | *-1.32* |
| Observations | 1,540 | 1,540 | 1,540 | 1,540 | 1,015 | 1,015 | 1,015 | 1,015 |
| R-squared | 0.039 | 0.037 | 0.071 | 0.173 | 0.035 | 0.034 | 0.065 | 0.177 |

Panel B: Fama-MacBeth regressions using DGTW-adjusted state-level returns across 17 states with above-median local biases

| | 1965-2008 | | | | 1980-2008 | | | |
|---|---|---|---|---|---|---|---|---|
| FP | 0.0083*** | 0.0098*** | | | 0.0086*** | 0.0101*** | | |
| | *3.17* | *3.69* | | | *2.94* | *3.34* | | |
| Predicted FP | | | 0.0131** | 0.0140** | | | 0.0184*** | 0.0170** |
| | | | *2.59* | *2.56* | | | *3.17* | *2.59* |
| GSP growth | | 0.1877 | | 0.0433 | | 0.4774 | | 0.2304 |
| | | *0.75* | | *0.17* | | *1.55* | | *0.72* |
| GSP volatility | | -0.3557 | | -0.1026 | | -0.6376 | | -0.3603 |
| | | *-0.90* | | *-0.25* | | *-1.32* | | *-0.70* |
| Constant | -0.0013 | -0.1792 | -0.0003 | -0.0359 | 0.0005 | -0.4714 | 0.0025 | -0.2218 |
| | *-0.28* | *-0.67* | *-0.06* | *-0.14* | *0.09* | *-1.45* | *0.44* | *-0.66* |
| Observations | 748 | 748 | 748 | 748 | 493 | 493 | 493 | 493 |
| R-squared | 0.077 | 0.250 | 0.070 | 0.234 | 0.071 | 0.255 | 0.073 | 0.246 |

Panel C. Fama-MacBeth regressions using industry-adjusted state-level returns for 35 states

| | 1965-2008 | | | | 1980-2008 | | | |
|---|---|---|---|---|---|---|---|---|
| Interaction of FP and LB | | 0.0734*** | 0.0968*** | 0.0880*** | | 0.0747** | 0.1120*** | 0.1008** |
| | | *2.83* | *3.09* | *2.78* | | *2.52* | *2.91* | *2.68* |
| FP | 0.0020 | | -0.0023 | -0.0012 | 0.0013 | | -0.0036 | -0.0020 |
| | *1.18* | | *-1.10* | *-0.61* | *0.70* | | *-1.45* | *-0.86* |
| GSP growth | | | | 0.1530 | | | | 0.3121* |
| | | | | *1.18* | | | | *2.04* |
| GSP volatility | | | | 0.0117 | | | | -0.2616 |
| | | | | *0.04* | | | | *-0.68* |
| Constant | 0.0051* | 0.0044 | 0.0038 | -0.1563 | 0.0004 | -0.0002 | -0.0011 | -0.3161* |
| | *1.85* | *1.61* | *1.36* | *-1.13* | *0.13* | *-0.05* | *-0.35* | *-1.92* |
| Observations | 1,540 | 1,540 | 1,540 | 1,540 | 1,015 | 1,015 | 1,015 | 1,015 |
| R-squared | 0.038 | 0.047 | 0.081 | 0.178 | 0.030 | 0.041 | 0.078 | 0.177 |

Panel D. Fama-MacBeth regressions using industry-adjusted state-level returns across 17 states with above-median local biases

| | 1965-2008 | | | | 1980-2008 | | | |
|---|---|---|---|---|---|---|---|---|
| FP | 0.0092*** | 0.0099*** | | | 0.0100*** | 0.0105*** | | |
| | *3.32* | *3.54* | | | *2.92* | *3.02* | | |
| Predicted FP | | | 0.0110** | 0.0124** | | | 0.0141** | 0.0144* |
| | | | *2.09* | *2.21* | | | *2.09* | *1.99* |
| GSP growth | | -0.0066 | | -0.1814 | | 0.3354 | | 0.0946 |
| | | *-0.03* | | *-0.84* | | *1.28* | | *0.35* |
| GSP volatility | | 0.2013 | | 0.4048 | | -0.0023 | | 0.2180 |
| | | *0.49* | | *0.94* | | *0.00* | | *0.44* |
| Constant | 0.0068* | 0.0188 | 0.0072* | 0.1979 | 0.0041 | -0.3389 | 0.0050 | -0.0921 |
| | *1.79* | *0.08* | *1.95* | *0.86* | *0.83* | *-1.22* | *1.07* | *-0.32* |
| Observations | 748 | 748 | 748 | 748 | 493 | 493 | 493 | 493 |
| R-squared | 0.115 | 0.293 | 0.081 | 0.251 | 0.115 | 0.299 | 0.093 | 0.266 |

# Table 6: Firm-level regressions

The table reports the results from Fama-MacBeth regressions that examine the firm-level stock return implications of fiscal policy. Firm-level stock returns are risk-adjusted using the Fama-French four-factor procedure from 1965 to 1975 and the DGTW procedure from 1976 to 2008. Panel A contains the results for all firms starting in 1980 based on the following Fama-MacBeth specification, $\text{Firm Return}_k = \beta_0 + \beta_1\,\text{FP}_i + \beta_2\,\text{GSP}_i + \beta_3\,\text{GSP volatility}_i + \epsilon_k$. An average of 2,510 firms per year across 49 states are examined. GSP growth represents the average per capita annual growth rate in a state's gross state product while GSP volatility represents the standard deviation of this growth rate over the sample period. In Panel B, the sample is restricted to firms that have local investment biases. These firms have institutional investors located in the same state as the firm's headquarters that hold 20% more than the firm's percentage weight in the market portfolio, with these investors also owning at least 5% of the firm's outstanding shares. This restriction requires investors located in the same state as a firm's headquarters to have a local investment bias and constitute a relatively large fraction of the firm's shareholder base. The results in Panel B are obtained from an average of 666 stocks per year across 44 states. The state-level fiscal policy betas are measured according to equation (2) as the sensitivity of annual budget deficits to economic conditions. Predicted fiscal policy betas are defined by the first stage of our instrumental variables procedure in equation (4) that predicts state-level fiscal policy betas using balanced budget amendments and budget stabilization funds defined by the DACIR and FNO variables in Table 2. *t*-statistics are in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively.

Panel A: Fama-MacBeth regressions for all firms

| | | | | |
|---|---|---|---|---|
| FP | 0.0017 | 0.0017 | | |
| | *0.91* | *1.31* | | |
| Predicted FP | | | 0.0010 | 0.0026* |
| | | | *0.66* | *1.85* |
| GSP growth | | 0.7033** | | 0.8564** |
| | | *2.05* | | *2.28* |
| GSP volatility | | -0.6407 | | -0.7873 |
| | | *-1.10* | | *-1.37* |
| Constant | 0.0053 | -0.7160** | 0.0054 | -0.8715** |
| | *0.93* | *-2.00* | *0.99* | *-2.22* |
| R-squared | 0.001 | 0.006 | 0.001 | 0.005 |

Panel B: Fama-MacBeth regressions for firms with local investment biases

| | | | | |
|---|---|---|---|---|
| FP | 0.0049** | 0.0039*** | | |
| | *2.05* | *2.71* | | |
| Predicted FP | | | 0.0032** | 0.0051*** |
| | | | *2.57* | *2.92* |
| GSP growth | | 0.2213 | | 0.3201 |
| | | *0.40* | | *0.52* |
| GSP volatility | | 0.3410 | | 0.2363 |
| | | *0.46* | | *0.29* |
| Constant | 0.0101 | -0.2428 | 0.0115* | -0.3433 |
| | *1.22* | *-0.42* | *1.87* | *-0.53* |
| R-squared | 0.003 | 0.008 | 0.002 | 0.009 |

## Table 7: Robustness Tests

Panel A of this table examines firm-level average returns before and after headquarter relocations during the 1986 to 2008 sample period. Firm-level return differences are regressed on changes in the fiscal policy betas (FP) associated with the relocations of firm $k$ from state $i$ to state $j$ as in equation (11), $\text{Return}_{k,j,t+} - \text{Return}_{k,i,t-} = \beta_0 + \beta_1 \left(\text{FP}_{k,j,t+} - \text{FP}_{k,i,t-}\right) + \gamma X + \epsilon_k$ where $t+$ refers to the firm's average annual return after the relocation and $t-$ refers to the firm's average annual return before the relocation in year $t$. The control variables in the $X$ vector are differences in state-level GSP growth and GSP volatility. These differences pertain to the pre-relocation and post-relocation headquarter states. Firms are required to have at least two years of stock returns before and after their headquarter relocation. Furthermore, returns in the first year after a relocation are eliminated from $\text{Return}_{k,j,t+}$ to ensure that takeover premiums arising from mergers and acquisitions are not influencing our results. The cross-sectional regression is conducted across the full sample of 924 firm-level relocations as well as a subset of 462 firm-level relocations between two states with above-median local investment biases (as defined in Table 2) and a separate subset involving 462 headquarter relocations where at least one of the two states involved does not have an above-median local investment bias. Panel B reports state-level flow betas denoted $\beta_{i,1}$ in equation (12), $\text{FLOW}_{i,t} = \beta_{i,0} + \beta_{i,1}\,\text{GSP}_{i,t} + \epsilon_{i,t}$ that regress annual mutual fund flows on contemporaneous state-level economic growth. The flow betas from these time series regressions, which are conducted at the state-level, are listed along with the original FP betas for 9 states with above-median local investment biases and sufficient mutual fund flow data. $t$-statistics are in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively.

Panel A: Headquarter relocations

| | All states | | Local bias states | | All other states | |
|---|---|---|---|---|---|---|
| Δ FP | 0.0175** | 0.0194** | 0.0245* | 0.0273* | 0.0150 | 0.1064 |
| | *2.05* | *2.15* | *1.69* | *1.86* | *1.38* | *1.35* |
| Δ GSP growth | | -1.6416* | | -2.4208* | | -1.1368 |
| | | *-1.71* | | *-1.70* | | *-0.86* |
| Δ GSP volatility | | -0.8292 | | -0.5954 | | -0.7662 |
| | | *-0.35* | | *-0.20* | | *-0.19* |
| Constant | 0.0090 | 0.0030 | 0.0259 | 0.0135 | -0.0073 | -0.0097 |
| | *0.39* | *0.12* | *0.84* | *0.42* | *-0.21* | *-0.27* |
| Observations | 924 | 924 | 462 | 462 | 462 | 462 |

Panel B: Fiscal policy and fund flows

| State | Fiscal Policy Beta | Flow Beta |
|---|---|---|
| MI | -1.518 | 1.620 |
| MN | -1.468 | -2.284 |
| VA | -0.892 | -15.381 |
| PA | -0.765 | -0.014 |
| OH | -0.672 | 1.772 |
| GA | 0.226 | 5.236 |
| NC | 0.396 | 0.450 |
| IL | 1.001 | 2.125 |
| KY | 2.080 | 3.108 |

## Table 8: Fiscal policy and cash flow risk

This table examines an alternative cash flow channel through which counter-cyclical fiscal policies can lower average stock returns. Panel A replicates the instrumental variables estimation in Table 4 by replacing state-level time series consumption volatility with state-level cash flow volatility. State-level cash flow volatility is computed as the average cash flow volatility of all firms headquartered in each state, with cash flow volatility at the firm-level computed as the standard deviation of annual cash flows across the 1965 to 2008 sample period. Firm-level cash flows are defined as earnings before extraordinary items minus total accruals, scaled by total assets. Panel B examines the firm-level stock return implications of firm-level CF-FP$_k$ estimates that weight state-level fiscal policy betas for all states in which firm $k$ operates. Data from 1994 to 2008 on the state-level operations of firms is obtained from Garcia and Norli (2010). The Fama-MacBeth regressions examine the contribution of fiscal policy to stock returns through the cash flow channel. The control variables in this regression are firm-level book-to-market (BM), size (SIZE), and past return (PRET) characteristics where PRET is defined as the firm's return over the prior twelve months after omitting the most recent month. The firm-level returns that form the dependent variable of the regression in Panel B are not risk-adjusted due to the inclusion of firm characteristics as control variables. *t*-statistics are in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively.

Panel A: Instrumental variables procedure for cash flow volatility

| | | Cash flow volatility |
|---|---|---|
| | first stage | second stage |
| Interaction (DACIR x FNO) | -3.7056** | |
| | *-1.97* | |
| DACIR | 1.6249* | |
| | *1.90* | |
| FNO | 7.3853** | |
| | *2.29* | |
| Predicted CF-FP | | 0.3226 |
| | | *1.27* |
| Constant | -3.4447** | 3.7369*** |
| | *-2.37* | *12.01* |
| Number of Observations | 35 | 35 |
| R-squared | 0.152 | 0.004 |

Panel B: Fama-MacBeth regressions using cash flow beta

| | | | |
|---|---|---|---|
| CF-FP | 0.0001 | 0.0002 | 0.0001 |
| | *0.61* | *0.69* | *0.56* |
| BM | | 0.0014 | 0.0021*** |
| | | *1.17* | *2.05* |
| SIZE | | 0.0000 | 0.0000 |
| | | *-0.01* | *0.02* |
| PRET | | | 0.0037* |
| | | | *1.71* |
| Constant | 0.0072*** | 0.0062 | 0.0044 |
| | *3.10* | *0.95* | *0.65* |
| Number of Firms | 1,931 | 1,931 | 1,931 |
| R-squared | 0.000 | 0.020 | 0.032 |

## Table 9: Geographically concentrated firms and cash flow risk

This table examines an alternative cash flow channel through which counter-cyclical fiscal policies lower average stock returns. The sample of firms is restricted to those with operations in three states or less. Data from 1994 to 2008 on the state-level operations of firms is obtained from Garcia and Norli (2010). Panel A replicates the instrumental variables estimation in Table 8 for firms operating in three states or less. Once again, state-level cash flow volatility is computed as the average cash flow volatility of all firms headquartered in each state, with cash flow volatility at the firm-level computed as the standard deviation of annual cash flows across the 1965 to 2008 sample period. Firm-level cash flows are defined as earnings before extraordinary items minus total accruals, scaled by total assets. Panel B examines the firm-level stock return implications of firm-level CF-FP$_k$ estimates that weight the state-level fiscal policy betas for all states in which firm $k$ operates. The Fama-MacBeth regressions examine the contribution of fiscal policy to average stock returns for geographically concentrated firms operating in three states or less. The control variables in this regression are firm-level book-to-market (BM), size (SIZE), and past return (PRET) characteristics where PRET is defined as the firm's return over the prior twelve months after omitting the most recent month. The firm-level returns that form the dependent variable of the regression in Panel B are not risk-adjusted due to the inclusion of firm characteristics as control variables. $t$-statistics are in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively.

Panel A: Instrumental variables procedure for cash flow volatility

| | | Cash flow volatility |
|---|---|---|
| | first stage | second stage |
| Interaction (DACIR x FNO) | -3.7056** | |
| | *-1.97* | |
| DACIR | 1.6249* | |
| | *1.90* | |
| FNO | 7.3853** | |
| | *2.29* | |
| Predicted CF-FP | | 2.3146* |
| | | *1.76* |
| Constant | -3.4447** | 10.1774*** |
| | *-2.38* | *7.16* |
| Number of Observations | 35 | 35 |
| R-squared | 0.152 | 0.026 |

Panel B: Fama-MacBeth regressions using cash flow beta

| | | | |
|---|---|---|---|
| CF-FP | 0.0006*** | 0.0006*** | 0.0005*** |
| | *2.09* | *2.19* | *1.97* |
| BM | | 0.0022* | 0.0031*** |
| | | *1.69* | *2.40* |
| SIZE | | 0.0000 | 0.0002 |
| | | *0.06* | *0.36* |
| PRET | | | 0.0037*** |
| | | | *2.18* |
| Constant | 0.0084*** | 0.0069 | 0.0029 |
| | *3.36* | *1.03* | *0.45* |
| Number of Firms | 440 | 440 | 440 |
| R-squared | 0.000 | 0.024 | 0.039 |

## Table 10: Education and fiscal policy

This table reports on the relationship between fiscal policy and three proxies for educational attainment and the likelihood of being employed in a knowledge-based industry. University-level expenditures on research and development within each state normalized by GSP (R&D), the number of patents per capita (PATENT), andthe proportion of the population with university degrees (UNI) are examined. The following cross-sectional regression examines the impact of our proxies on the state-level fiscal policy betas (FP), as in equation (15), $\text{FP}_i = \beta_0 + \beta_1 \text{R\&D}_i + \beta_2 \text{PATENT}_i + \beta_3 \text{UNI}_i + \epsilon_i$. $t$-statistics are reported in parentheses below each of the estimates with ***, **, and * denoting statistical significant at the 1%, 5%, and 10% levels, respectively.

| | FP |
|---|---|
| R&D | 10.6057 |
| | *0.43* |
| PATENT | -0.0018 |
| | *-1.32* |
| UNI | -199.8298 |
| | *-1.19* |
| Constant | 1.7477 |
| | *1.41* |
| Observations | 50 |
| R-squared | 0.092 |