Banking Conditions and the Effects of Monetary Policy: Evidence from U.S. States

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Abstract

Using data from U.S. states, this paper examines empirically how the effect of monetary policy on output depends on banking conditions. It is found that when a state's banking sector starts out with a low capital-asset ratio, its subsequent output growth is more sensitive to changes in the Federal funds rate or other indicators of monetary policy. This result is consistent with the existence of a 'bank capital channel' as well as a conventional bank lending channel. I attempt to distinguish between these two explanations by including a bank liquidity variable.

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1. Introduction

Theories of how monetary policy affects the real economy differ in the role they accord to financial intermediaries. Most models abstract from financial intermediaries altogether, focusing directly on the saving/investment decisions of households and firms. This abstraction is justified if financial intermediaries are Modigliani-Miller agents that frictionlessly allocate funds between the other agents in the economy, so that they are no more than a 'veil' on the 'nonfinancial' economy. Other theories suggest a nontrivial role for banks or other financial intermediaries by incorporating financial imperfections that prevent banks from providing frictionless intermediation. A consequence of this nontrivial role of banks is that the effects of monetary policy on the real economy may depend on the financial structure of banks, which itself can depend on the stance of monetary policy. For example, according to the 'bank lending channel' thesis, monetary policy is more potent if banks have low levels of liquid assets. According to a different theory, the 'bank capital channel', bigger effects of monetary policy on bank lending are to be expected if banks have low equity relative to existing bank capital requirements.[2]

Motivated by these theories, the purpose of this paper is to document if and how monetary policy effects on output depend on the financial condition of the banking sector. Data from U.S. states is employed in order to address some of the main identification issues that arise in interpreting the predictive power of the banking variables in terms of different theories of the monetary transmission mechanism.[3] The main finding is that when a state's banking sector starts out with a low capital-asset ratio, its subsequent output growth is more sensitive to changes in the Federal funds rate or other indicators of monetary policy. I will argue that this finding is precisely what is to be expected based on both the 'bank lending channel' and the 'bank capital channel', although other evidence presented favors the latter.

[2] I will return to each of these examples below.

[3] John Driscoll (2004) uses state-level data to identify loan supply movements attributable to a 'bank lending channel' by looking at variation in money demand across U.S. states. Carlino and DeFina (1999) examine the variation in the response to monetary policy shocks across U.S. states by estimating identified VARs.

The rest of the paper is organized as follows. The next section reviews the literature and explains what predictions the 'bank lending channel' and the 'bank capital channel' make about the dependence of monetary policy effects on banks' financial structure. This also motivates the selection of bank variables to be examined. Section three presents the empirical model. Section four presents and discusses the results, followed by several checks for alternative explanations. The final section concludes.

2. Why would banks' financial structure affect the monetary transmission mechanism? A selective review of the literature

According to the lending view monetary policy affects output at least in part through its impact on the supply of bank loans to firms.[4] Two conditions are necessary for a distinct bank lending channel to be operative. First, some firms do not have perfect substitutes for bank loans available as a means of financing their activities. In other words, bank loans are special to firms. Second, by changing the quantity of reserves available to the banking system, monetary policy can affect the supply of bank loans. This requires that banks do not fully insulate their loan supply from shocks to their reserves. Whether these conditions hold in reality, is a controversial question. For example, Romer and Romer (1990) argue that bank loan supply is effectively insulated from reserve shocks because banks can frictionlessly switch to alternative forms of finance by issuing CDs or other securities. Kashyap and Stein (1995) and Stein (1998), however, have countered that the type of Modigliani Miller logic Romer and Romer appeal to will fail, if there is asymmetric information about the value of the bank's assets. In that case, as Stein's model shows, adverse selection leads to a 'lemon's premium' in the market for risky bank liabilities. Since most nonreservable bank liabilities are not insured, they are therefore at least somewhat risky so that the market for them is likely to be imperfect.

[4] I use the term 'lending view' to refer specifically to the role of bank intermediated loans, in contrast to the 'broad credit channel' in which bank loans do not play a special role. See Bernanke and Blinder (1988) for an exposition of the lending channel. Recognition that assets other than money and bonds play a role in the monetary transmission mechanism dates back further (e.g. Brainard (1964)). Kashyap and Stein (1994) provide an overview of some of the work relating to the lending view of monetary policy.

If this is the case, then, as Stein's model makes clear, banks with a low buffer stock of liquid assets should cut back their lending more in response to a monetary tightening. The reason is that banks that have large amounts of very liquid securities have the option of selling those when faced with an outflow of reserves and deposits. In contrast, banks with few liquid assets face a choice between cutting back on lending or issuing uninsured managed liabilities. Because the latter is costly, some of the adjustment will take place through a reduction in loan supply.

Using bank-level data in a test of the second prerequisite for the lending channel (the Fed affects the supply of bank loans), Kashyap and Stein (2000) find evidence in favor of exactly this prediction: less liquid banks reduce lending more when monetary policy tightens.[5] Thus, *to the extent that the lending channel is quantitatively important, we should expect monetary policy effects on output to be stronger when banks are less liquid.*

According to an alternative theory, the bank capital channel, monetary policy affects bank lending in part through its impact on bank equity capital. In a separate paper (Van den Heuvel, 2009), a dynamic bank asset and liability management model is presented that formalizes this 'channel' and analyzes its consequences for monetary policy.[6] The model incorporates the risk based capital requirements of the Basle Accord and an imperfect market for bank equity. These two conditions imply a failure of the Modigliani-Miller logic for the bank: its lending will depend on the bank's financial structure, as well as on lending opportunities and market interest rates. When equity is sufficiently low, because of loan losses or some other adverse shock, the bank will cut

[5] Evidence from aggregate U.S. time series on the lending channel is not conclusive. Loans move roughly contemporaneously with output. Bernanke and Blinder (1992) find that loans, like output, fall after some lag following monetary tightening, as measured by a positive innovation to the Federal funds rate. While this is consistent with the lending view, it is also consistent with a pure money channel, if the fall in loans is due to declining demand for credit, rather than a shrinking supply. The decline in loan demand could be due to the fall in output caused by the monetary contraction in a pure money channel economy. Distinguishing between movements in loan demand and movements in loan supply constitutes a difficult identification problem, especially since we do not observe one interest rate summarizing the effective cost of bank loan finance. This depends not only on the contractual interest rate, but also on collateral requirements, the extent of rationing, etc.
[6] See also Van den Heuvel (2002), which provides a summary of the bank capital channel and a comparison of its implications to the bank lending channel.

lending due to the capital requirement and the cost of issuing new equity.[7]

Another crucial feature of the model, besides capital adequacy regulations and an imperfect market for bank equity, is the maturity transformation performed by banks which exposes them to interest rate risk. A consequence of this is that a monetary tightening, by raising the short interest rate, lowers bank profits. Unless the bank can cut dividends substantially, this will result over time in lower bank capital and, given the failure of the Modigliani-Miller logic, less lending. Thus, monetary policy affects the supply of bank loans through its effect on bank equity. This dynamic effect, the bank capital channel, amplifies the standard interest rate channel of monetary policy. As with the lending channel, if some firms are bank-dependent (and other firms do not pick up the slack), the amplification of the lending response translates into a larger output response.

The size and the dynamics of the effect are highly dependent on the initial level and distribution of capital among banks. The reason is that banks the capital requirement affects bank behavior more when bank equity is low relative to the regulatory minimum. In that situation, there is less room for capital to absorb adverse shocks without cutting back on lending. Thus, *the amplification is much stronger for banks that start out with already low capital than for very well-capitalized banks.*[8]

To sum up the discussion so far, *based on the lending channel*, we expect output effects of monetary policy actions to be larger, *if bank liquidity is low*. If the *bank capital channel* is quantitatively important, we expect output effects of monetary policy actions to be larger, *if bank capital is low*. Of course, the two channels are by no means mutually exclusive.

Before testing these predictions, it is worth making one more point. While the bank lending channel is in essence a 'liquidity phenomenon' (if all banks always have sufficient cash or liquid securities, or can access a frictionless market for some managed

[7] Even when the capital requirement is not currently binding, the model shows that a low capital bank may optimally forego profitable lending opportunities now, in order to lower the risk of future capital inadequacy. This is interesting since in reality, and in the model, as calibrated with U.S. data, most banks are not at the capital constraint at any given time.

[8] The lending response of a bank with capital so low that the capital requirement is actually binding at the time of the shock may exhibit an initial delay, as lending is already depressed due to that binding constraint (see Van den Heuvel (2009) for details). Hence, if bank equity is low, the monetary policy effects on

liability, there is no lending channel), there may very well be a connection between the strength of the lending channel and the level of bank capital. To see this, consider two banks with the same quality assets, but different liability structures – bank one, say, has less equity, and more debt, than bank two. Suppose further that, following a contractionary monetary policy shock, both banks face an equal outflow of reservable deposits. Thus, both banks need to issue managed liabilities, such as large denomination CDs, to keep lending at normal level. Even though both banks have equally risky assets, bank one's CDs are more risky, because bank one has less equity to absorb future losses. Consequently, they are more exposed to any asymmetric information problems concerning the value of the bank's assets and thus command a larger 'lemon's premium'. Hence, following the contractionary monetary shock, bank one will optimally choose to issue fewer CDs and cut back lending by more than the better capitalized bank two. The conclusion is that, other things being equal, the lending channel is likely to be stronger for banks with lower levels of capital. Thus, a finding that monetary policy effects are larger when bank capital is low would be consistent with either banking channel.

3. The empirical model and data

The goal is to examine whether, at any given point in time, output in states with poor inherited banking conditions is more responsive to monetary policy actions than output in states where banks start out in better financial shape. As mentioned, data from U.S. states are employed to take advantage of both cross-sectional and time-variation in banking conditions. This is valid only if bank loan markets are not perfectly integrated across U.S. states. A certain degree of imperfect integration is certainly to be expected before the Riegle-Neal Interstate Banking and Branching Act of 1994. The Riegle-Neal Act lifted most restrictions still in effect on interstate branching and mergers of banks. For this reason, the sample period is set to 1969 – 1995, as the branching restrictions were formally lifted by the Act on September 29, 1995. (Merger restrictions were not lifted

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lending via the bank capital channel may be weak initially, but will be much larger after one or several until June 1997.) To the extent that the gradual lifting at the state level of cross-state branching restrictions since the 1980s has created a national market for bank loans, this will make it harder for us to find the effects of state-level banking variables.[9] Finally, it is worth noting that geographic specialization may also result in imperfect integration.

Several annual panel data models are estimated; a limitation of the data, to be described in more detail below, is that they are available only at annual frequency. The following is a baseline specification:

$$\begin{aligned}\Delta y_{it} = \alpha_i &+ (\beta_{US} + \delta_{US} c_{it-1})\Delta y_{USt} + (\beta_M + \delta_M c_{it-1})\Delta M_t + \beta_{c1} c_{it-1} \\ &+ (\beta_{US1} + \delta_{US1} c_{it-2})\Delta y_{USt-1} + (\beta_{M1} + \delta_{M1} c_{it-2})\Delta M_{t-1} \\ &+ (\beta_{y1} + \delta_{y1} c_{it-2})\Delta y_{it-1} + \beta_{c2} c_{it-2} + \varepsilon_{it}\end{aligned} \tag{1}$$

where Δy_{it} is the output growth of state *i* in year *t*, Δy_{USt} is the same quantity for the U.S., ΔM_t is an indicator of the change in the stance of monetary policy, with positive values indicating a loosening of monetary conditions, and c_{it-1} is (some transformation of) the aggregate capital asset ratio of all commercial banks in state *i* at the *end* of year $t - 1$ (explained in more detail below). α_i is a fixed effect for state *i*, which is included in some specifications.

The δ coefficients measure how the sensitivity of state output growth to U.S. output growth, monetary policy shocks and lagged state output growth depends on lagged banking conditions. One lag of the regressors is included to remove any autocorrelation in the error term. Independent variables dated $t - 1$ are interacted with c_{it-2} so that δ_{M1} does not pick up any reverse causality running from, say, ΔM_{t-1} to c_{it-1}.[10]

Based on the bank capital channel, one would expect to find $\delta_M < 0$ and $\delta_{M1} < 0$. That is, when a monetary expansion occurs ($\Delta M > 0$), states with a well capitalized

quarters.

[9] Jayaratne and Strahan (1996) and Morgan, Rime and Strahan (2004) examine the effects of the lifting of branching restrictions at the state level on state output growth and fluctuations.

[10] For example, suppose that some states for some reason happen to be more affected by a rise in the funds rate. These states are likely to see their banking conditions, as well as output, worsen more than other states, which would cause a variable such as $c_{it-1}\Delta M_{t-1}$ to be significant even if banking conditions are irrelevant.

banking sector ($c_{it\text{-}1} > 0$) enjoy a smaller output expansion. Put differently, when the Fed raises rates ($\Delta M < 0$), states with a poorly capitalized banking sector ($c_{it\text{-}1} < 0$) suffer a *larger drop* in income, over the next one or two years.

To examine the importance of bank liquidity, the following relation is also estimated:

$$\begin{aligned}\Delta y_{it} = \alpha_i &+ (\beta_{US} + \delta_{US} s_{it-1})\Delta y_{USt} + (\beta_M + \delta_M s_{it-1})\Delta M_t + \beta_{s1} s_{it-1} \\ &+ (\beta_{US1} + \delta_{US1} s_{it-2})\Delta y_{USt-1} + (\beta_{M1} + \delta_{M1} s_{it-2})\Delta M_{t-1} \\ &+ (\beta_{y1} + \delta_{y1} s_{it-2})\Delta y_{it-1} + \beta_{s2} s_{it-2} + \varepsilon_{it}\end{aligned} \tag{2}$$

where s_{it} is (some transformation of) the ratio of investment securities to total assets of all commercial banks in state i, year t. Based on the bank lending channel, we would again expect $\delta_M < 0$ and $\delta_{M1} < 0$.

As a measure of economic activity annual total personal income by state is used, which is available from the Bureau of Economic Analysis of the U.S. Department of Commerce. U.S. personal income is from the same source. To compute real personal income, these series are deflated by the U.S. GDP-deflator, since there is no complete set of state-level price indices. The use of personal income data, rather than some other measure of economic activity, is motivated by the limited availability of state-level data over a reasonable time span.

The state banking data are from the FDIC’s Historical Statistics on Banking. The state capital asset ratios is defined as total capital of all FDIC-insured commercial banks in the state divided by total assets of all FDIC-insured commercial banks in the state. Figure 1 shows the cross-sectional mean of state capital asset ratios for each year, as well as the minimum and maximum values, and U.S. personal income growth. A liquidity ratio is also used. This is defined as investment securities divided by total assets of all FDIC-insured commercial banks in the state. At the state level, these data are available at the annual frequency. As mentioned, the sample period is 1969-1995.

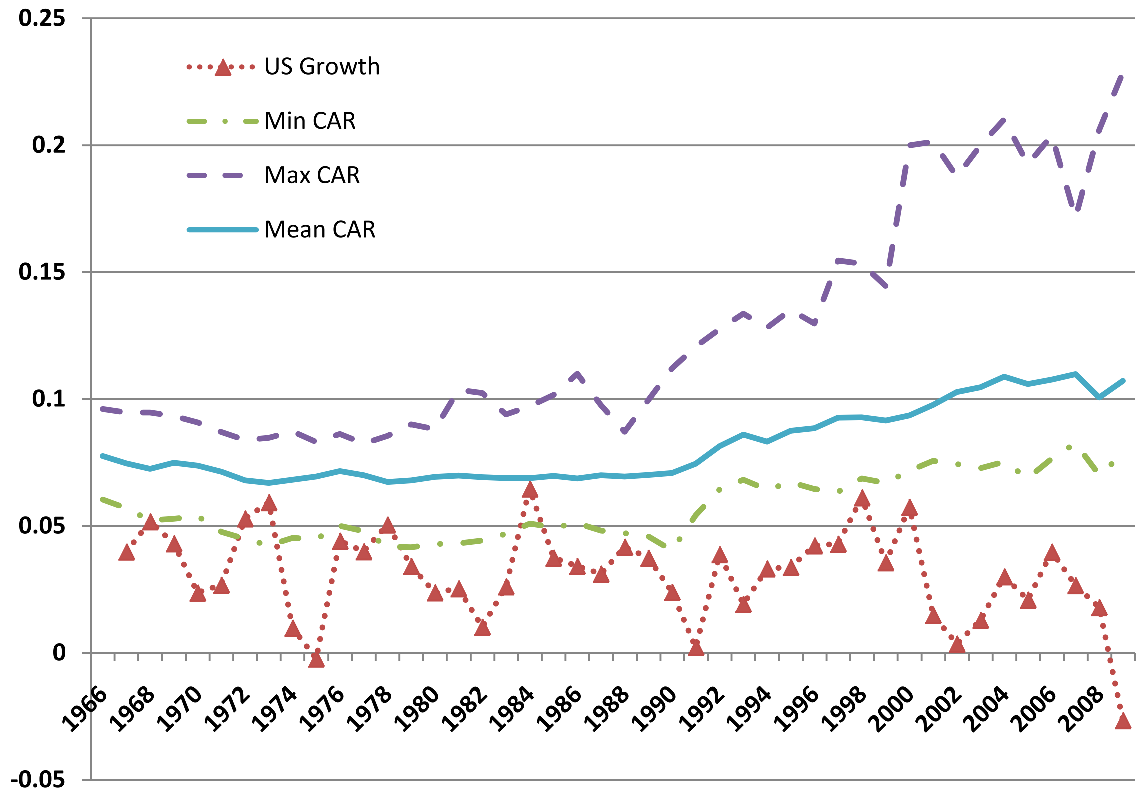

Figure 1. Bank capital asset ratios and U.S. income growth.

To measure the stance of monetary policy, I use, alternatively, the Federal funds rate and the 'overall indicator' of monetary policy constructed by Ben Bernanke and Ilian Mihov (1998), based on their identified VAR. Both are converted to annual averages and then differenced. The differenced Federal Funds rate is multiplied by (–1) to maintain the convention that positive values of ΔM correspond to looser monetary policy.

A few remarks about using the Federal Funds rate as a monetary policy indicator are in order. It is well understood that movements in the Federal Funds rate reflect both 'normal' reactions of the monetary authority to economic conditions, such as inflation and business cycle conditions, as well as what can be usefully thought of as 'random' monetary policy shocks. If all that matters to banks is the sum of the two components,[11]

[11] This is the case in the model of Van den Heuvel (2001) because loan demand is held constant in the face of interest rate shocks.

then there is no need to distinguish between the two. However, if the variables that the Fed ordinarily reacts to affect state output growth independently, a familiar identification problem may arise. In the empirical exercise of this paper the identification problem will manifest itself only if two further conditions are both satisfied. First, the variables in the Fed's 'reaction function' that independently affect state output do so *differentially, depending on states' capital asset or liquidity ratios*. And second, they are *imperfectly correlated with U.S. output growth*, since that variable is already included in (1) alongside the Funds rate. If both these conditions are met, then the interaction coefficients will pick up not only the differential effects of monetary policy but also of the variables that the Fed is reacting to.

While the second condition is most likely satisfied (e.g. inflation), it is less obvious that the first condition is. At a minimum, it seems unlikely that some variable in the Fed's reaction function would have a differential impact on output via the financial structure of the banking sector, while the same would not be true for monetary policy. Nonetheless, measuring the magnitude of the differential impact of monetary policy actions by the interaction coefficients δ is, strictly speaking, valid only under the assumption that the variables in the Fed's reaction function, other than U.S. output growth, have no direct effect on state output growth that is dependent on state-level banking conditions.[12]

With the annual state-level data, it is hard to see how the identification problem could be solved conclusively without some such assumption. Any high frequency series of identified exogenous monetary policy shocks will within the year affect other economic variables whose impact on state output growth could conceivably depend on state-level banking conditions.

In addition, I use the Bernanke-Mihov overall indicator as an alternative measure of the stance of monetary policy. This indicator is constructed based on an identified VAR, which controls for the usual reaction of the Fed to prevailing economic conditions.

[12] A similar issue arises in Kashyap and Stein (1995), Kashyap and Stein (2000), and Kishan and Opiela (2001), and other studies.

Thus, employing this indicator can be expected to mitigate the potential problem.[13] In addition, it will provide a useful check on the robustness of the results.

4. Results

As a reference point, it is useful to consider what happens if one estimates model (1) and (2) under the restriction that $\delta_M = \delta_{M1} = 0$. Using the Federal Funds rate,[14] the result of that regression is (standard errors in parenthesis):

$$\underset{}{\Delta y_{it}} = \alpha_i + \underset{(0.07)}{0.94}\Delta y_{USt} + \underset{(0.05)}{0.06}\Delta M_t + \underset{(0.03)}{0.27}\Delta y_{it-1} - \underset{(0.06)}{0.27}\Delta y_{USt-1} - \underset{(0.06)}{0.04}\Delta M_{t-1}$$

The coefficient on dy_{USt} is approximately equal to one and the coefficient on dM_t is approximately equal to zero, reflecting the almost tautological fact that states respond *on average* neither more nor less than the U.S. to changes in the funds rate. In fact, the result is statistically and economically indistinguishable from

$$\Delta y_{it} - \Delta y_{USt} = \alpha_i + 0.27(\Delta y_{it-1} - \Delta y_{USt-1}) .$$

[13] The indicator is constructed from a semi-structural VAR that identifies monetary policy shocks. It contains an endogenous component of monetary policy, as well as the VAR-identified monetary policy shocks. Bernanke and Mihov's methodology takes into account potentially time-varying operating procedures of the central bank.

[14] Using the Bernanke Mihov indicator produces very similar results.

Table 1. Capital Asset Ratio and Federal Funds Rate

Variable:	(a) Capital Asset Ratio: $c_{it} = C_{it}$	(b) Deviation from state mean: $c_{it} = C_{it} - \overline{C}_i$	(c) Dev. from state and time mean: $c_{it} = C_{it} - \overline{C}_i - \overline{C}_t + \overline{\overline{C}}$
$c_{it-1}\Delta y_{USt}$	- 19.50** (6.18)	- 30.15** (8.73)	-28.04** (9.89)
$c_{it-2}\Delta y_{USt-1}$	- 6.74 (5.66)	- 12.34 (7.53)	-22.56** (8.67)
$c_{it-1}\Delta M_t$	- 12.10** (4.37)	- 26.30** (6.74)	-30.81** (7.96)
$c_{it-2}\Delta M_{t-1}$	2.88 (4.64)	14.48* (6.77)	7.41 (7.96)
$c_{it-2}\Delta y_{it-1}$	-2.86 (2.81)	-0.69 (2.90)	-2.72 (3.23)
c_{it-1}	0.37 (0.22)	0.46 (0.25)	0.45 (0.30)
c_{it-2}	0.29 (0.22)	0.35 (0.26)	0.60 (0.31)
$\Sigma\Delta M$ -test	- 9.21* (4.65)	-11.82 (7.39)	-23.40** (9.17)

Note: ΔM equals the *negative* of the change in the Federal Funds rate.
Standard errors are in parenthesis.
* indicates significance at the 0.05 level; ** at the 0.01 level.

4.1 Capital Asset Ratios

Table 1, column (a) presents the results of specification (1) with c_{it} equal to the (untransformed) state-level capital asset ratio (C_{it}). Only the interaction coefficients (δ) are reported. The prediction of the bank capital channel is born out: δ_M is significantly less than zero (at the 0.01 level), which means that states that start out year *t* with low bank capital, subsequently have lower (higher) output growth if the Federal Funds rate is increased (decreased) than other states or times. (Recall that an increase in the Funds rate corresponds to $dM < 0$.) In other words, their output growth is more sensitive to changes in the Federal Funds rate.

A second interesting result is that output growth of low capital state-years is also more sensitive to U.S. output growth, as indicated by the significantly negative coefficient on $c_{it-1}dy_{USt}$. One interpretation is that poor bank capitalization amplifies not only monetary shocks, as emphasized above, but also real shocks. Incidentally, an implication of this is that a monetary contraction affects states with poor banking conditions disproportionately, not only because of a rising Funds rate but also because it is likely to depress nationwide economic activity. In this sense, δ_M *underestimates* the true differential impact of the Federal Funds rate on low versus high capital states. A likelihood ratio test overwhelmingly rejects the hypothesis $\delta_M = \delta_{M1} = 0$.[15]

Are these results driven by specific states? If so, we might be especially reluctant to interpret as causal the correlation bank capital on the one hand and the effect of U.S. monetary and real shocks on state output on the other hand. For example, due to differences in the sectoral composition of output, some states tend to be more cyclical than others. If these states also happen to have low bank capital, then this could explain the result even if bank capital is itself irrelevant.[16] While one might expect banks in these states to in fact choose to maintain a larger buffer stock of equity on average, it seems prudent to check that the result is not driven by some state-specific bias.

This can be achieved by deducting state means from the capital asset ratio. That is, we use the following transformation of the capital asset ratio in specification (1):

$$c_{it} = C_{it} - \overline{C}_i \equiv C_{it} - \frac{1}{T}\sum_{s=1}^{T} C_{is}$$

where $\overline{C}_i$ is state *i*'s mean capital asset ratio over the sample period. Defined this way, c_{it} has zero mean for each state and any significance of δ cannot be attributed to factors that vary only across states. This transformation removes close to half of the variance of C_{it}: the standard deviation of $(C_{it} - \overline{C}_i)$ is 0.0082 versus a standard deviation of 0.010 for C_{it}. Column (b) of table 1 reports the results using this transformation.

[15] The same is true for all alternative specifications using the capital asset ratio below.

[16] Note that state fixed effects do not correct for this potential problem because the interactive specification measures a second derivative.

As can be seen, the results are similar to column (a). If anything the results are even stronger. δ_M , the coefficient on $c_{it-1}\Delta M_t$, is even larger in absolute value, although the *additional* sensitivity when capital is low is reversed after one year, as indicated by the coefficient on $c_{it-2}\Delta M_{t-1}$. The bank-dependent amplification of U.S. output shocks is also larger. No reversal of the additional effect occurs. Hence, removing state-specific bias strengthens the conclusion that low bank capital translates greater sensitivity to U.S. monetary and real shocks. This is exactly what one would expect if banks in more cyclical states optimally choose to hold more capital on average.

If the results in column (a) are not driven by state-specific bias, could they be the result of time-specific bias? As can be seen in figure 1, there is a general upward movement in state capital asset ratios starting around 1990, when the Basle Accord, which established risk-based capital requirements, was being implemented. Suppose monetary policy became more potent around that time for some other reason, then this could conceivably drive our results. To check for this, year means are also deducted from the capital asset ratios. Column (c) of table 1 reports the results of estimating (1) using the following transformation:

$$c_{it} = C_{it} - \overline{C}_i - \overline{C}_t + \overline{\overline{C}}$$

where $\overline{C}_t \equiv \frac{1}{N}\sum_{j=1}^{N} C_{jt}$ and $\overline{\overline{C}} \equiv \frac{1}{T}\sum_{t=1}^{T} \overline{C}_t$.[17] This lowers the standard deviation of c_{it} to 0.0066 compared to 0.0082 in column (b) and 0.010 in column (a).

Again, the results are, if anything, more pronounced using this ‘bias-correction’: low capital state-years show a greater response to changes in the Federal Funds rate and U.S. output. This is what one would expect if the rise in capital asset ratios toward the end of the sample is a response to effectively more stringent capital regulation.

To obtain an idea of the economic size of the effect, we examine the differential impact of a one standard deviation change in the Federal Funds rate (dM_t = 0.024), where

[17] Deducting only year means, and not state means, yields similar results (not reported).

the difference is between the states with the highest and the lowest (transformed) capital asset ratio in a typical year. To be precise, the time-average of the yearly differences between the highest and lowest values of c_{it} is computed: $(1/T)\sum_{t=1}^{T}(\max_i c_{it} - \min_i c_{it})$. For the raw capital asset ratios this number is 0.050; after deducting state (and time) means this number is 0.034. Taking the latter, more conservative, number, we can compute the differential response after one year as $0.034 \times 0.024 \times \delta_M$. After two years, the cumulative effect equals $0.034 \times 0.024 \times (\delta_M + \delta_{M1})$.[18] Using the point estimates, the resulting numbers are:

Column:	(a)	(b)	(c)
1 year:	- 0.010	- 0.021	- 0.025
2 years:	- 0.008	- 0.010	- 0.019

Thus, on a one year horizon the differential response of state income growth to a one standard deviation shock to the Funds rate ranges from - 1.0% for the raw capital asset ratio to - 2.5% after taking out state and year effects. After two years, the effects are somewhat smaller in absolute value. It is important to note that these numbers do not take into account any effect that the changes in the Federal Funds rate have on U.S. output. To the extent that a monetary policy induced rise in the Funds rate lowers U.S. growth, the true magnitudes are larger in absolute value than the above numbers.

[18] This ignores the interaction with the lagged dependent variable, $c_{it-2}dy_{it-1}$, which is always insignificant.

Table 2. Capital Asset Ratio and Bernanke-Mihov Indicator

Variable:	(a) Capital Asset Ratio: $c_{it} = C_{it}$	(b) Deviation from state mean: $c_{it} = C_{it} - \overline{C}_i$	(c) Dev. from state and time mean: $c_{it} = C_{it} - \overline{C}_i - \overline{C}_t + \overline{\overline{C}}$
$c_{it-1}\Delta y_{USt}$	-16.25** (5.81)	-16.12* (8.19)	-20.16* (9.74)
$c_{it-2}\Delta y_{USt-1}$	-4.09 (5.98)	-13.73 (7.80)	-20.82* (9.09)
$c_{it-1}\Delta M_t$	-3.96 (2.24)	-6.53* (2.97)	-9.39* (3.82)
$c_{it-2}\Delta M_{t-1}$	1.75 (2.16)	1.12 (2.75)	-0.27 (3.41)
$c_{it-2}\Delta y_{it-1}$	-2.78 (2.82)	-0.60 (2.91)	-2.75 (3.25)
c_{it-1}	0.20 (0.21)	0.03 (0.24)	0.18 (0.29)
c_{it-2}	0.26 (0.23)	0.43 (0.27)	0.55 (0.31)
$\Sigma\Delta M$ - test	-2.21 (2.64)	-5.41 (3.48)	-9.66* (4.53)

Note: ΔM equals the change in the overall monetary policy indicator by Bernanke and Mihov (1998). Standard errors are in parenthesis.
* indicates significance at the 0.05 level; ** at the 0.01 level.

Table 2 reports the results of estimating (1) using the Bernanke-Mihov indicator of monetary policy, instead of the Federal Funds rate. Again, state means and state and year means are deducted from the capital asset ratios in columns (b) and (c), respectively. The results are broadly similar. δ_M is negative in all three cases and significant (at the 0.05) when the state-specific 'bias correction' is used (columns (b) and (c)). The sensitivity of state output growth to U.S. growth is also larger for low capital state-years. Using the same methodology as for the Funds rate, on a one year horizon, the differential response of state income growth to a one standard deviation change in the Bernanke-Mihov indicator (ΔM = 0.047) ranges from - 0.6% for the raw capital asset ratio to - 1.5% after taking out state and year effects. For two years, it ranges from - 0.4% to - 1.6%.

Table 3. Liquidity Ratio and Bernanke-Mihov Indicator

Variable:	(a) Liquidity Ratio: $s_{it} = S_{it}$	(b) Deviation from state mean: $s_{it} = S_{it} - \bar{S}_i$	(c) Dev. from state and time mean: $s_{it} = S_{it} - \bar{S}_i - \bar{S}_t + \bar{\bar{S}}$
$s_{it-1}\Delta y_{USt}$	0.43 (0.94)	2.81 (1.50)	2.30 (1.79)
$s_{it-2}\Delta y_{USt-1}$	3.21** (1.02)	3.68* (1.58)	2.12 (1.85)
$s_{it-1}\Delta M_t$	-0.33 (0.38)	0.03 (0.59)	0.28 (0.71)
$s_{it-2}\Delta M_{t-1}$	1.11** (0.36)	1.34* (0.55)	0.79 (0.65)
$s_{it-2}\Delta y_{it-1}$	-1.23** (0.40)	-0.79 (0.53)	-0.63 (0.65)
s_{it-1}	-0.04 (0.05)	-0.12* (0.06)	-0.11 (0.07)
s_{it-2}	0.00 (0.05)	-0.02 (0.06)	0.00 (0.06)
$\Sigma\Delta M$ - test	0.77 (0.46)	1.37 (0.73)	1.07 (0.85)

Note: ΔM equals the change in the overall monetary policy indicator by Bernanke and Mihov (1998). Standard errors are in parenthesis.
* indicates significance at the 0.05 level; ** at the 0.01 level.

4.2 Liquidity Ratios

As mentioned, the bank lending channel thesis of the monetary transmission mechanism predicts that, other things being equal, monetary policy effects are larger when banks are less liquid. To see if we can find evidence for this prediction, model (2) is estimated using various transformations of the liquidity ratio, defined as the ratio of investment securities to total assets, for s_{it}. Table 3 reports the results using the Bernanke-Mihov indicator for ΔM_{it}. Using the Federal Funds rate produces similar results.

As can be seen, there is no support for the prediction that $\delta_M < 0$ or $\delta_{M1} < 0$. The coefficients are either not significantly different from zero or have the 'wrong' sign. This is true whether untransformed liquidity ratios are used, as in column (a), deviations from state means (b), or deviations from both state and year means (c). Thus, we find no

Table 4. Capital Asset Ratio, Liquidity Ratio and Bernanke-Mihov Indicator

Variable:	(a) Ratio:	(b) Deviations from state mean:	(c) Dev. from state and time mean:
$c_{it-1}\Delta y_{USt}$	-20.19** (6.40)	-16.06 (8.37)	-19.82* (9.80)
$c_{it-2}\Delta y_{USt-1}$	-11.32 (6.51)	-16.17* (7.93)	-20.60* (9.14)
$c_{it-1}\Delta M_{t}$	-3.45 (2.44)	-7.25* (3.08)	-9.51* (3.85)
$c_{it-2}\Delta M_{t-1}$	0.02 (2.35)	0.71 (2.81)	-0.04 (3.44)
$s_{it-1}\Delta y_{USt}$	1.62 (1.03)	3.14* (1.56)	1.97 (1.80)
$s_{it-2}\Delta y_{USt-1}$	4.09** (1.09)	3.99* (1.64)	1.78 (0.34)
$s_{it-1}\Delta M_{t}$	-0.19 (0.41)	0.23 (0.61)	0.29 (0.71)
$s_{it-2}\Delta M_{t-1}$	1.00* (0.39)	1.29* (0.56)	0.78 (0.65)

Note: ΔM equals the change in the overall monetary policy indicator by Bernanke and Mihov (1998). Standard errors are in parenthesis.
* indicates significance at the 0.05 level; ** at the 0.01 level.

evidence in favor of the prediction that state output growth is more sensitive to monetary policy actions when bank liquidity is low.

Table 4 shows that including both capital asset and liquidity ratios in a bivariate interactive specification confirms these results. It contains the results from estimating the following relation: [19]

$$\begin{aligned}\Delta y_{it} = {} & \alpha_i + (\beta_{US} + \delta_{US} c_{it-1} + \lambda_{US} s_{it-1})\Delta y_{USt} + (\beta_M + \delta_M c_{it-1} + \lambda_M s_{it-1})\Delta M_t \\ & + (\beta_{US1} + \delta_{US1} c_{it-2} + \lambda_{US1} s_{it-2})\Delta y_{USt-1} + (\beta_{M1} + \delta_{M1} c_{it-2} + \lambda_{M1} s_{it-2})\Delta M_{t-1} \\ & + (\beta_{y1} + \delta_{y1} c_{it-2} + \lambda_{y1} s_{it-2})\Delta y_{it-1} + \beta_{c1} c_{it-1} + \beta_{s1} s_{it-1} + \beta_{c2} c_{it-2} + \beta_{s2} s_{it-2} + \varepsilon_{it}\end{aligned} \tag{3}$$

Including both banking conditions does not seem to reduce the predictive power of the capital asset ratio in interaction with the monetary policy indicator or U.S. output. At the

[19] For brevity interactive coefficients on the constant and lagged dependent variable are not reported.

same time, nor does it do much to improve the predictive power of the liquidity ratio. The results in table 4 are for the Bernanke-Mihov indicator, but using the Federal Funds rate again produces very similar results.

4.3 Capital Asset Ratios and Local Income Growth

The finding that income growth in states that start out the year with a poorly capitalized banking sector is more sensitive to subsequent changes in the Federal Funds rate or the Bernanke-Mihov indicator, is consistent with the bank capital channel thesis. As explained, it is also consistent with the lending channel, although the results using the liquidity ratio provide little support for the latter interpretation.

Other interpretations of the finding are, of course, possible. The bank capital asset ratio is not a quantity that is exogenous to business cycle conditions. Suppose that monetary policy effects on output do not causally depend on bank health, but on some other variable that co-varies with banks' capital asset ratios over the business cycle. In that case, the interactive coefficients δ would show up as significant, even though there is no causal effect of bank capital. One candidate example for such an alternative variable might be the financial health of nonfinancial firms, if a broad credit channel is operative.

We can go some way toward exploring if such an alternative interpretation is behind the results by including lagged state income growth alongside the capital asset ratio in a bivariate interactive specification. If local business cycle conditions other than bank capital are driving the results, we would expect the inclusion of lagged state income growth to greatly reduce the significance of bank capital. Table 5 reports the results of estimating the following relation:

$$\begin{aligned}\Delta y_{it} = \alpha_i &+ (\beta_{US} + \delta_{US} c_{it-1} + \lambda_{US} \Delta\tilde{y}_{it-1})\Delta y_{USt} + (\beta_M + \delta_M c_{it-1} + \lambda_M \Delta\tilde{y}_{it-1})\Delta M_t \\ &+ (\beta_{US1} + \delta_{US1} c_{it-2} + \lambda_{US1} \Delta\tilde{y}_{it-2})\Delta y_{USt-1} + (\beta_{M1} + \delta_{M1} c_{it-2} + \lambda_{M1} \Delta\tilde{y}_{it-2})\Delta M_{t-1} \\ &+ (\beta_{y1} + \delta_{y1} c_{it-2} + \lambda_{y1} \Delta\tilde{y}_{it-2})\Delta y_{it-1} + \beta_{c1} c_{it-1} + \beta_{y1} \Delta\tilde{y}_{it-1} + \beta_{c2} c_{it-2} + \beta_{y2} \Delta\tilde{y}_{it-2} + \varepsilon_{it}\end{aligned} \tag{4}$$

where $d\tilde{y}_{it}$ is state income growth, its deviation from the state's mean income growth, or its deviation from state and time means, respectively in columns (a), (b) and (c). That is,

Table 5. Capital Asset Ratio, Income Growth and Bernanke-Mihov Indicator

Variable:	(a) Ratio/Growth rate:	(b) Deviations from state mean:	(c) Dev. from state and time mean:
$c_{it-1}\Delta y_{USt}$	-15.70** (5.80)	-12.34 (8.33)	-18.83* (9.72)
$c_{it-2}\Delta y_{USt-1}$	-3.63 (5.98)	-13.38 (7.74)	-23.17* (9.10)
$c_{it-1}\Delta M_t$	-3.78 (2.26)	-6.45* (3.06)	-9.94** (3.82)
$c_{it-2}\Delta M_{t-1}$	3.47 (2.16)	4.19 (2.76)	1.54 (3.42)
$\Delta\tilde{y}_{it-1}\Delta y_{USt}$	5.87** (2.26)	5.67* (2.51)	3.69 (2.37)
$\Delta\tilde{y}_{it-2}\Delta y_{USt-1}$	0.52 (0.95)	-0.44 (0.97)	-0.06 (1.17)
$\Delta\tilde{y}_{it-1}\Delta M_t$	0.52 (0.88)	0.17 (0.96)	0.59 (1.00)
$\Delta\tilde{y}_{it-2}\Delta M_{t-1}$	1.76** (0.62)	2.05** (0.66)	3.15** (0.73)

Note: ΔM equals the change in the overall monetary policy indicator by Bernanke and Mihov (1998)
Standard errors are in parenthesis.
* indicates significance at the 0.05 level; ** at the 0.01 level.

in each case, the same transformation is applied to $d\tilde{y}_{it}$ as to c_{it}.

The inclusion of lagged state income growth as an interactive variable does not, apparently, alter the estimates of δ much at all. Low bank capital still translates into greater sensitivity to U.S. monetary and output shocks. Interestingly, *high* state growth seems to lead to somewhat greater sensitivity to changes in U.S. income and the monetary policy indicator, although the effects are small in size.

5. Concluding remarks

When a U.S. state's banking sector starts out with a low capital-asset ratio, its subsequent output growth is more sensitive to changes in the Federal funds rate or the Bernanke-Mihov indicator of monetary policy. This finding survives removing state- or time effects, or both, from the capital asset ratios and is not purely driven by correlation of the ratio with local output growth. One interpretation of this result has been suggested: that a 'bank capital channel' (Van den Heuvel 2009) is operative, whereby monetary policy affects bank lending through its effects on bank capital.

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