



# The Long-Run Relationship Between Consumption and Housing Wealth in the Eighth District States

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The authors examine the long-run relationship between consumption and housing wealth for the seven individual states in the Federal Reserve System's Eighth District. Given that state-level consumption data are not available, the authors develop a novel proxy for state-level consumption based on state-level data for personal income and savings income. They use this consumption proxy to estimate a cointegrating relationship between consumption spending and housing wealth, stock market wealth, and income in each of the Eighth District states. Their results indicate that increases in housing wealth produce sizable increases in consumption for most of the states in the Eighth District. Interestingly, the authors also find that consumption typically responds much more strongly to changes in housing wealth than to changes in stock market wealth. Their results imply that the strong increases in housing prices and home construction over the past decade have helped to buoy consumption and decrease saving in most Eighth District states. (JEL C32, C33, E21)

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**H**igh levels of consumption spending—which have driven the personal saving rate below zero during the past year—together with continued increases in housing prices are two U.S. economic facts that currently receive considerable attention in both the popular and financial press. It is natural to speculate that these two facts are linked, and analysts have posited that the strong increases in housing wealth experienced over the past decade in the United States have played an important role in stimulating household spending. There is also concern that a slowing of the housing market in the near future will depress household spending and help precipitate a general economic slowdown. For example, Ben Bernanke, current Chairman of the Board of Governors of the Federal Reserve System, remarked in early 2006 that “given the

substantial gain in house prices and the high levels of home construction activity over the past several years, prices and construction could decelerate more rapidly than currently seems likely. Slower growth in home equity, in turn, might lead households to boost their saving and trim their spending relative to current income by more than is now anticipated” (Bernanke, 2006).

The general interest in the link between consumption spending and housing wealth, along with its potential interest to policymakers, motivates the present paper, where we undertake a formal econometric analysis of the long-run relationship between consumption spending and housing wealth. We concentrate on the relationship between consumption spending and housing wealth in the seven states of the Federal Reserve System's Eighth District (Arkansas, Illinois, Indiana, Kentucky, Missouri, Mississippi, and Tennessee).

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There is already a large body of research, falling under the rubric of the “wealth effect,” that examines the relationship between consumption spending and household wealth.<sup>1</sup> This literature either focuses on the response of consumption to changes in financial wealth alone—especially stock market wealth—or assumes that all forms of wealth are viewed equivalently by households. As stressed by Case, Quigley, and Shiller (2005), this is a potentially important drawback to this literature: Households may view different forms of wealth differently, so that consumption can respond differently to changes in financial compared with housing wealth. For example, financial market frictions, due to certain types of liquidity constraints created by information asymmetries, may make it easier for households to increase their consumption by borrowing against increases in housing values, as evidenced by the sharp rise in home equity loans that have accompanied the strong increases in housing values over the past decade.<sup>2</sup> In addition, households may separate their wealth into different “mental accounts,” so that changes in different categories have different effects on household consumption (the psychology of “framing”).

In contrast to the substantial literature on the wealth effect, there is a relatively small literature that specifically examines the response of consumption spending to changes in housing wealth.<sup>3</sup> Nevertheless, some recent studies suggest that there are important differences in how consumption responds to changes in financial and housing wealth. Using aggregate U.S. data, Benjamin, Chinloy, and Jud (2004) estimate that the marginal propensity to consume from real estate wealth is approximately four times larger than the marginal propensity to consume from financial wealth. Using a panel of U.S. state-level data, Case, Quigley, and Shiller (2005) find that household wealth has a significant and sizable effect on household consumption, an effect that is significantly larger than that of stock

market wealth. The present paper contributes to this recent literature by analyzing the long-run relationship between consumption spending and housing wealth in the Eighth District states.

Our econometric methodology involves estimating a cointegrating relationship between real consumption spending, housing wealth, stock market wealth, and income (a “long-run consumption function”) in each of the Eighth District states. A challenge in estimating a long-run consumption function for individual states is that state-level consumption data are not readily available. We develop a novel proxy for consumption spending at the state level that allows us to estimate a cointegrating relationship that is informative about the long-run relationship between actual (but unobserved) consumption and housing wealth in each Eighth District state. In analyzing cointegrating relationships, we pay careful attention to the integration properties of all the variables appearing in our model. Unit root tests, including unit root tests for heterogeneous panels, indicate that all of the variables in our model (more precisely, their log-levels) are integrated processes, so that it is appropriate to consider potential cointegrating relationships.

We estimate cointegrating relationships using a number of well-known procedures, and we find that consumption is significantly and positively related to housing wealth in most of the Eighth District states. We also find that housing wealth typically has a much stronger effect on consumption than stock market wealth. Panel cointegration tests support the existence of cointegrating relationships in a significant portion of the Eighth District states. Our finding of a significant and sizable housing wealth effect on consumption is in line with the recent studies of Benjamin, Chinloy, and Jud (2004) and Case, Quigley, and Shiller (2005), and our results support the conjecture that increases in housing wealth over the past decade have contributed significantly to strong consumption growth.

There is an important way in which our results differ from Benjamin, Chinloy, and Jud (2004) and Case, Quigley, and Shiller (2005): The homogeneity assumptions implicit in both of these studies are likely to be inappropriate and can mask important differences in the responses of consumption to

<sup>1</sup> See, for example, the surveys in Ludvigson and Steindel (1999), Poterba (2000), and Davis and Palumbo (2001).

<sup>2</sup> See Greenspan and Kennedy (2005).

<sup>3</sup> Case, Quigley, and Shiller (2005) provide a survey of this literature, and they note that most of the studies in this area are micro studies of consumer behavior. They conclude that there is “much ambiguity in the interpretation of statistical results.”

changes in housing wealth across regions of the United States. For example, we find that the consumption response to a change in housing wealth is much stronger in Illinois than it is in Arkansas. In summary, the housing wealth effect is not uniform across the Eighth District states.

The next section describes our estimation strategy, and the following section reports our estimation results.

## ESTIMATION STRATEGY

An important problem in analyzing the relationship between consumption and housing wealth in individual states is that consumption data are not readily available at the state level. In this section, we outline our strategy of using a proxy for state-level consumption that enables us to analyze the long-run relationship between consumption and housing wealth in the individual states of the Eighth District.

A state's household consumption is clearly equal to the difference between a state's personal disposable income and personal saving. While state-level income data are available from the Bureau of Economic Analysis (BEA), state-level personal saving data are not available. However, the BEA does report personal savings income, which consists of dividend, interest, and rental income from prior accumulated savings. It is likely that permanent changes in household saving will lead to permanent changes in the flow of income derived from accumulated savings; we exploit this likely link between saving and savings income to construct a proxy for consumption at the state level that enables us to analyze the long-run relationship between consumption and housing wealth at the state level. More specifically, we use available data to construct a proxy (personal disposable income minus personal savings income) for actual—but unavailable—consumption (personal disposable income minus personal saving). If there is a stable long-run relationship between actual consumption and our proxy, then we can use our proxy to analyze the long-run relationship between actual consumption and housing wealth. We emphasize that we view our consumption proxy only as a useful long-run proxy,

such that it will not necessarily be informative with respect to short-run dynamics.<sup>4</sup>

Let  $c_{i,t}^S$  equal the log-level of the difference between personal disposable income and personal saving in state  $i$ , and let  $c_{i,t}^{DIR}$  equal the log-level of the difference between personal disposable income and savings income in state  $i$  (after both differences have been converted to real terms). Assuming  $c_{i,t}^S, c_{i,t}^{DIR} \sim I(1)$  (as is likely to be the case), a stable long-run relationship exists between  $c_{i,t}^S$  and  $c_{i,t}^{DIR}$  when these two variables are cointegrated [ $c_{i,t}^S, c_{i,t}^{DIR} \sim CI(1,1)$ ], and the cointegrating relationship can be expressed as

$$(1) \quad c_{i,t}^S = \alpha_i + \beta_i c_{i,t}^{DIR} + e_{i,t},$$

where  $e_{i,t}$  is an  $I(0)$  disturbance term with an unconditional mean of zero. If a cointegrating relationship of the form in equation (1) exists for each individual state in the Eighth District, we can exploit this to analyze the long-run relationship between consumption spending and housing wealth for each state.

Consider the possibility of the existence of a stable long-run relationship between consumption and housing wealth, as well as stock market wealth and income, in state  $i$ :

$$(2) \quad c_{i,t}^S = \gamma_i + \delta_{i,hw} hw_{i,t} + \delta_{i,sw} sw_{i,t} + \delta_{i,y} y_{i,t} + u_{i,t},$$

where  $hw_{i,t}$  is the log-level of real housing wealth in state  $i$ ,  $sw_{i,t}$  is the log-level of real stock market wealth in state  $i$ ,  $y_{i,t}$  is the log-level of real personal disposable income in state  $i$ , and  $u_{i,t}$  is a stationary, zero-mean disturbance term.<sup>5</sup> Equation (2) is a standard type of specification for a long-run con-

<sup>4</sup> Given the household budget constraint that labor income plus savings income equals consumption plus saving, using personal income minus personal savings income as a proxy for consumption essentially assumes that labor income serves as a proxy for consumption. We also note that changes in rates of return potentially affect savings income in ways that have an impact on saving behavior, but it is likely that these effects are small relative to the long-run effect that we isolate. Overall, whether personal disposable income minus personal savings income is a reasonable proxy for consumption over the long run is an empirical matter, and we present evidence below that there is a stable long-run relationship between our proxy for consumption and actual consumption for aggregate U.S. data.

<sup>5</sup> We could also allow a linear time trend in equation (2), but this does not affect our results in important ways, as the estimates for equation (3) reported in Table 2 change little if we include a linear time trend in equations (2) or (3). Complete results with a linear time trend included are available from the authors upon request.

sumption function; see, for example, Davis and Palumbo (2001). While we ideally would analyze equation (2) directly, as discussed above, we cannot estimate equation (2) directly because state-level data for  $c_{i,t}^S$  are not available. However, we can use equation (1) to substitute for  $c_{i,t}^S$  in equation (2):

$$(3) \quad c_{i,t}^{DIR} = \zeta_i + \theta_{i,hw} hw_{i,t} + \theta_{i,sw} sw_{i,t} + \theta_{i,y} y_{i,t} + \varepsilon_{i,t},$$

where  $\zeta_i = (\gamma_i - \alpha_i) / \beta_i$ ;

$\theta_{i,j} = \delta_{i,j} / \beta_i$  for  $j = hw, sw, y$ ; and  $\varepsilon_{i,t} = (u_{i,t} - e_{i,t}) / \beta_i$

Note that  $\varepsilon_{i,t}$  is a stationary, zero-mean process, as both  $u_{i,t}$  and  $e_{i,t}$  are stationary, zero-mean processes.

Equation (3) provides considerable information about the parameters of interest in equation (2). Note the following:

- (i)  $\delta_{i,k} = 0$  implies  $\theta_{i,k} = 0$  (assuming  $\beta_i < \infty$ );
- (ii)  $\beta_i > 0$  implies  $sign(\theta_{i,k}) = sign(\delta_{i,k})$ ;
- (iii)  $\delta_{i,j} > \delta_{i,k}$  implies  $\theta_{i,j} > \theta_{i,k}$ ;
- (iv)  $\beta_i \approx 1$  implies  $\delta_{i,k} = \theta_{i,k}$ .

According to (i), we can use equation (3) to analyze the statistical significance of the slope parameters in equation (2). According to (ii), if  $\beta_i > 0$ , as it almost surely is, the signs of the slope coefficients in equation (3) are the same as those in equation (2). According to (iii), we can also compare the relative sizes of the slope parameters in equation (2) using equation (3). Finally, according to (iv), insofar as  $\beta_i$  approaches unity,  $\delta_{i,k}$  approaches  $\theta_{i,k}$ .

In the next section, we estimate the cointegrating relationship in equation (3) using standard procedures: ordinary least squares (OLS), fully modified OLS (FMOLS; Phillips and Hansen, 1990), and dynamic OLS (DOLS; Saikkonen, 1991, and Stock and Watson, 1993). While OLS is super-consistent, it is subject to an endogeneity bias that renders conventional inferential procedures invalid.<sup>6</sup> The FMOLS and DOLS procedures address the endogeneity bias and permit valid inference. Of course, to treat equation (3) as a cointegrating relationship,

<sup>6</sup> Loosely speaking, the endogeneity bias exists when there is feedback from the left-hand-side variable to the right-hand-side variables in equation (3), as there almost surely is in our applications.

$c_{i,t}^{DIR}$ ,  $hw_{i,t}$ ,  $sw_{i,t}$ , and  $y_{i,t}$  all need to be integrated processes. We test for a unit root in these variables using the familiar augmented Dickey and Fuller ([ADF] 1979) test, as well as a more-powerful panel version of the test from Im, Pesaran, and Shin (2003). For equation (3) to be a valid cointegrating relationship, it obviously must be the case that  $c_{i,t}^{DIR}$ ,  $hw_{i,t}$ ,  $sw_{i,t}$ , and  $y_{i,t}$  are cointegrated. We test for cointegration using the well-known augmented Engle and Granger ([AEG] 1987) two-step test and a more-powerful panel version of the test from Pedroni (1999, 2004).

The key to our estimation strategy is the existence of a stable long-run relationship between  $c_{i,t}^S$  and  $c_{i,t}^{DIR}$ . Although we obviously cannot test for the existence of such a relationship for each state, we can test whether the variables are cointegrated in aggregate U.S. data. Evidence of cointegration between these two variables at the national level is highly suggestive that similar cointegrating relationships exist at the state level. Using BEA data for 1975:Q1–2004:Q4, we construct observations for  $c_{US,t}^S$  and  $c_{US,t}^{DIR}$ , the aggregate counterparts to  $c_{i,t}^S$  and  $c_{i,t}^{DIR}$ .<sup>7</sup> The ADF statistics for  $c_{US,t}^S$  and  $c_{US,t}^{DIR}$  are  $-2.66$  and  $-1.71$ , respectively, and in neither case can the null hypothesis of a unit root be rejected at conventional significance levels (the 5 percent critical value equals  $-3.45$ ), indicating that  $c_{US,t}^S, c_{US,t}^{DIR} \sim I(1)$ .<sup>8</sup> The AEG statistic (which includes a constant in the potential cointegrating relationship) equals  $-3.64$ , so that the null hypothesis of no cointegration between  $c_{US,t}^S$  and  $c_{US,t}^{DIR}$  can be rejected at the 5 percent significance level (the 5 percent critical value equals  $-3.34$ ), indicating that  $c_{US,t}^S, c_{US,t}^{DIR} \sim CI(1,1)$ . We expect  $\beta_{US}$  in equation (1) to be positive and relatively close to unity, and the OLS, FMOLS, and DOLS estimates of  $\beta_{US}$  in equation (1) all equal 1.13. The finding of a cointegrating relationship between  $c_{US,t}^S$  and  $c_{US,t}^{DIR}$  at the national level increases our confidence that similar cointegrating relationships exist at the state level, and the estimates of

<sup>7</sup> The observations are converted to real terms using the personal consumption expenditure deflator.

<sup>8</sup> Given the obvious upward drift in  $c_{US,t}^S$  and  $c_{US,t}^{DIR}$ , a constant and linear trend are included the ADF tests. The number of lags included in the ADF and AEG tests is determined using a top-down procedure based on a maximum lag of four quarters. We obtain similar results using the “state of the art” unit root tests in Ng and Perron (2001).

**Table 1****ADF Test Results, Eighth District States, 1975:Q1–2004:Q4**

State	$c_{i,t}^{DIR}$	$hw_{i,t}$	$sw_{i,t}$	$y_{i,t}$
AR	-3.08	-0.46	-2.80	-3.59*
IL	-2.00	-1.95	-2.88	-2.14
IN	-1.54	-0.91	-2.77	-1.82
KY	-2.51	-0.99	-2.83	-3.07
MO	-1.71	-1.96	-3.22	-1.98
MS	-1.87	-0.38	-2.94	-2.26
TN	-3.19	-1.24	-3.13	-3.60*
Panel test				
$W_{tbar}$	-0.80	0.68	0.87	-0.89

NOTE: The table reports the ADF statistic, which corresponds to the null hypothesis that the variable has a unit root against the one-sided (lower-tail) alternative hypothesis that the variable is stationary; the 5 percent critical value equals -3.45. The  $W_{tbar}$  statistic corresponds to the null hypothesis that each of the variables in the panel has a unit root against the one-sided (lower-tail) alternative hypothesis that at least a portion of the variables in the panel is stationary; the 5 percent critical value equals -1.645. \*Significant at the 5 percent level.

$\beta_{US}$  at the national level further indicate that estimation of equation (3) at the state level will be informative about the parameters in equation (2) at the state level.<sup>9</sup>

## ESTIMATION RESULTS

Quarterly data for 1975:Q1–2004:Q4 for personal income, savings income (dividends, interest, and rental income), and the personal consumption expenditure (PCE) deflator from the BEA are used to construct state-level observations for  $c_{i,t}^{DIR}$ . The Office of Federal Housing Enterprise Oversight (OFHEO) provides a housing price index for each state. The Census Bureau provides annual housing quantity data for each state, and we interpolate the annual data to obtain quarterly observations for the quantity of housing in each state. The housing quantity data end in 2004:Q4, which is the endpoint for all of our samples. We obtain real housing wealth by multiplying housing quantity by housing price and dividing by the PCE deflator. Real stock

wealth is obtained from quarterly aggregate S&P 500 stock market capitalization data available from Global Financial Data. We compute real stock market wealth for state  $i$  by first multiplying the proportion of aggregate U.S. dividends paid to state  $i$  by aggregate S&P 500 stock market capitalization and then dividing by the PCE deflator.

Table 1 reports unit root test results for  $c_{i,t}^{DIR}$ ,  $hw_{i,t}$ ,  $sw_{i,t}$ , and  $y_{i,t}$  for Arkansas, Illinois, Indiana, Kentucky, Missouri, Mississippi, and Tennessee. The ADF statistics indicate that we almost always fail to reject the null hypothesis that the variables are unit root processes.<sup>10</sup> A potential drawback to using the ADF statistic is that it may have limited power against persistent, but stationary, alternatives. In light of this, we also employ the more powerful Im, Pesaran, and Shin (2003) panel unit root test based on the  $W_{tbar}$  statistic, which is essentially an average of the individual ADF statistics. From Table 1, we can see that the null hypothesis that each variable contains a unit root cannot be rejected at the 5 percent significance level using the panel test, so we have substantial evidence that  $c_{i,t}^{DIR}$ ,  $hw_{i,t}$ ,  $sw_{i,t}$ ,  $y_{i,t} \sim I(1)$  for each of the Eighth District

<sup>9</sup> Case, Quigley, and Shiller (2005) derive retail sales observations from county-level sales tax data to construct a proxy for consumption. It is unclear how reliable this consumption proxy is, and Case, Quigley, and Shiller (2005) do not examine the relationship between retail sales and consumption on the national level to get a feel for reliability.

<sup>10</sup> We obtain similar results using the Ng and Perron (2001) unit root tests.

**Table 2****Coefficient Estimates for Equation (3), Eighth District States, 1975:Q1–2004:Q4**

State	$\hat{\theta}_{i,hw}^{OLS}$	$\hat{\theta}_{i,sw}^{OLS}$	$\hat{\theta}_{i,y}^{OLS}$	$\hat{\theta}_{i,hw}^{FMOLS}$	$\hat{\theta}_{i,sw}^{FMOLS}$	$\hat{\theta}_{i,y}^{FMOLS}$	$\hat{\theta}_{i,hw}^{DOLS}$	$\hat{\theta}_{i,sw}^{DOLS}$	$\hat{\theta}_{i,y}^{DOLS}$
AR	0.030* (0.007)	0.004* (0.002)	0.967* (0.008)	0.024* (0.012)	0.004 (0.004)	0.969* (0.016)	0.016 (0.021)	-0.003 (0.008)	0.992* (0.030)
IL	0.075* (0.007)	0.009* (0.002)	0.864* (0.013)	0.083* (0.010)	0.011* (0.003)	0.844* (0.020)	0.090* (0.012)	0.006 (0.004)	0.852* (0.023)
IN	0.067* (0.005)	-0.004* (0.002)	0.956* (0.007)	0.070* (0.009)	-0.006* (0.003)	0.959* (0.012)	0.072* (0.014)	-0.009 (0.005)	0.966* (0.019)
KY	0.067* (0.005)	-0.001 (0.001)	0.928* (0.005)	0.069* (0.009)	-0.004 (0.003)	0.933* (0.010)	0.069* (0.012)	-0.006 (0.004)	0.939* (0.015)
MO	0.060* (0.005)	0.008* (0.002)	0.920* (0.006)	0.063* (0.009)	0.007 (0.004)	0.919* (0.011)	0.065* (0.027)	0.007 (0.010)	0.918* (0.034)
MS	0.040* (0.005)	0.001 (0.002)	0.971* (0.006)	0.043* (0.009)	-0.002 (0.004)	0.977* (0.012)	0.043 (0.025)	-0.001 (0.010)	0.976* (0.035)
TN	0.059* (0.006)	-0.003 (0.002)	0.958* (0.005)	0.063* (0.012)	-0.005 (0.005)	0.960* (0.011)	0.062* (0.029)	-0.009 (0.012)	0.968* (0.028)

NOTE: Standard errors are given in parentheses; 0.00 indicates less than 0.005; \* denotes significance at the 5 percent level.

states.<sup>11</sup> Given these results, we proceed to estimate equation (3) and to test for cointegration among the variables in equation (3) for the individual states in the Eighth District.

Estimation results for equation (3) are reported in Table 2. The table reports OLS, FMOLS, and DOLS point estimates and corresponding standard errors for  $\theta_{i,hw}$ ,  $\theta_{i,sw}$ , and  $\theta_{i,y}$ . As noted above, the OLS standard errors are generally not valid for inference, and we include them only for the sake of completeness. The first thing to notice in Table 2 is that all of the estimates of  $\theta_{i,hw}$  are positive, indicating a positive long-run relationship between consumption spending and housing wealth in the states of the Eighth District. When using FMOLS (DOLS), seven (five) of the estimates are significant, the exceptions being the DOLS estimates for Arkansas and Mississippi. Overall, there is strong evidence in Table 2 for a positive and significant relationship between consumption and housing wealth for most of the states in the Eighth District.

<sup>11</sup> In order to account for a degree of cross-sectional dependence, a common time component is subtracted from each variable before computing the  $W_{ibar}$  statistics. For more discussion on issues relating to panel unit root tests, see the recent survey in Breitung and Pesaran (2005).

The estimates of  $\theta_{i,sw}$  in Table 2 are all smaller than the corresponding estimates of  $\theta_{i,hw}$ , and fewer of the estimates are significant.<sup>12</sup> This indicates that the housing wealth effect on consumption is generally much stronger than the stock market wealth effect in the Eighth District. The  $\theta_{i,y}$  estimates are all positive and reasonably close to unity, so that the values seem economically plausible.

The finding of a stronger housing wealth effect in comparison with a stock market wealth effect is in line with the results in Benjamin, Chinloy, and Jud (2004) and Case, Quigley, and Shiller (2005). However, the results in Table 2 also point to a potential problem with the approaches of both these studies. Both are based on the implicit homogeneity assumption that the cointegrating coefficients are the same across all states ( $\theta_{i,k} = \theta_k$  for all  $i$ ), whereas Table 2 shows that the cointegrating coefficients can differ substantially across states. For example, the  $\theta_{i,hw}$  estimates for Illinois, Indiana, Kentucky, Missouri, and Tennessee are typically around two to three times larger than the  $\theta_{i,hw}$  estimates for Arkansas and Mississippi. Imposing homogeneity

<sup>12</sup> In fact, a number of the  $\theta_{i,sw}$  estimates are negative, the opposite sign predicted by theory.

across states can thus mask important differences in the long-run relationship between consumption and housing wealth across regions, differences that can arise from differences in demographics, institutions, and other factors across regions.<sup>13</sup>

Finally, it is important to test for the existence of cointegrating relationships in the Eighth District states. The coefficient estimates reported in Table 2 assume the existence of a cointegrating relationship, and we have a spurious regression if the variables are not cointegrated. Applying the AEG test to the residuals in equation (3) for each state, we cannot reject the null hypothesis of no cointegration for any of the seven states, as the AEG statistics range from  $-1.51$  to  $-3.06$ , while the 5 percent critical value equals  $-4.10$ . However, we can employ the more powerful group  $t$  panel cointegration test of Pedroni (1999, 2004), which is essentially an average of the individual AEG statistics. The null hypothesis for this test is no cointegration for each of the panel members, and the one-sided (lower-tail) alternative is that a cointegrating relationship holds for a significant portion of the panel members. The (normalized) group  $t$ -statistic equals  $-2.66$ ; given a 5 percent critical value of  $-1.645$ , we can reject the null hypothesis of no cointegration. We thus have evidence that a cointegrating relationship holds for at least a significant number of the states in the Eighth District.<sup>14</sup>

## CONCLUSION

This paper examines the long-run relationship between consumption spending and housing wealth in the states of the Federal Reserve's Eighth District. The consumption-housing wealth relationship has received limited attention at the state level, in part

<sup>13</sup> Allowing for heterogeneity also gives a more accurate picture of the true underlying relationships because it avoids the bias that can result from pooling (Pesaran and Smith, 1995).

<sup>14</sup> To account for a degree of cross-sectional dependence, a common time component is subtracted from each variable before computing the group  $t$ -statistic. There are a number of econometric issues involved with panel cointegration tests, including how cross-sectional dependence is handled and the potential existence of cross-sectional cointegration; see the survey in Breitung and Pesaran (2005). We deal with these issues more extensively in ongoing research in which we consider the long-run relationship between consumption spending and housing wealth in each of the 50 individual U.S. states.

because of the lack of consumption data at this level. We develop a novel proxy for consumption at the state level that can be constructed on a quarterly basis since 1975, and this proxy can be used in a cointegration framework to analyze the long-run relationship between consumption spending and housing wealth. Our estimation results show that housing wealth exerts a significant and sizable influence on consumption spending for most of the states in the Eighth District, and this influence is typically stronger than that of stock market wealth. Our results imply that the strong increases in housing prices and home construction over the past decade have helped to buoy consumption in most of the states of the Eighth District; they also imply that sharp decreases in housing prices and home construction in the future will have a depressing effect on consumer spending.

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