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Financial Inclusion and Per Capita Real GSP Growth across Fifty US States and the District of Columbia: Evidences from Panel Cointegration and GMM Estimates

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ABSTRACT

This paper explores the effects of bank-based financial inclusion on per capita real GSP (Gross State Product) growth across all fifty U.S. states and the District of Columbia (DC) using panel data over 1997-2013. Estimates of Pedroni's panel vector error-correction model and GMM dynamic panel estimates reveal that increases in outstanding bank credit and bank-branch expansion exert significantly positive effects on per capita real GSP growth. Furthermore, rises in the number of other financial institutions and deposit mobilization by banks have subdued net positive effects on per capita real GSP growth.

Keywords: Financial Inclusion, Per Capita Real GSP growth, Panel Cointegration, Generalized Method of Moments

I. Introduction

Inclusive economic growth with financial inclusion of all segments of active-age population has been gaining global attention of policymakers and academicians since 2000. According to the World Bank estimate, 2.5 billion of the current world populations of 7.5 billion are excluded from access to formal financial network. This problem is common to both developing and developed countries in varying degrees. There are multiple definitions of financial inclusion, but none of them are universally accepted. In general, the basic tenet of financial inclusion is to bring low-income earners, less educated and small entrepreneurs under the formal banking and financial services. Benefits of financial inclusion are immense as

it spurs growth and household financial stability through diversifying funding base, lending base, savings channels and insurance. They are evidenced in various empirical studies by different authors all over the world.

In the 50 states of the USA and the DC (District of Columbia), there were nearly 9.6 million (7.7% of US households) unbanked in 2013. At the same time, 20% of US households remained under-banked in 2013. About 67.6 million adults and 25.3 million children fall in these two categories (1913, FDIC National Survey). A large percentage of poor households rely on predatory payday loans and pawnshops for weekly/biweekly loans against forthcoming paychecks. The effective annual interest rate on such loan ranges from 200% to 300% or even higher according to some estimates. Thus, making them bankable is imperative via financial inclusion to induce household financial stability. The 50 US states and the DC are different in terms of geography, demography, economic structures and levels of development. The US Fed recognizes the

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gravity of this situation and made financial inclusion a statutory requirement.

Franklin and Douglas (2001) conclude that financial systems are crucial to the allocation of resources in a modern economy. They channel household savings to the corporate sector and allocate investment funds among firms. They allow intertemporal smoothing of consumption by households and capital expenditures by firms. They also enable households and firms to share risks. These functions are common to the financial systems of most developed economies, although the forms of these financial systems differ widely.

This is evidenced by various studies in different countries by different authors all over the world. Burgess and Pande (2005), and Bruhn and Inessa (2009) and Ruiz (2010) study how expanding access to microfinance affects the economy. They found that increased access to both credit and savings services have positive impacts on the welfare of the society. Banerjee, et al (2010), Crépon et al (2011), and Karlan and Zinman (2010), however, suggest that expanding access to microfinance alone has no significant effect on the economy. So, it has been oversold in the news media of various forms all these years.

Microfinance institutions have played an important role in many countries in making short-term loans in small amounts at much lower rates relative to those charged by informal moneylenders. Again, studies of programs that increased access to both credit and savings services have found important welfare impacts [e.g., Burgess and Pande (2005), Aportela (1999), Bruhn and Inessa (2009) and, Ruiz (2010)]. The degree of banking exclusion varies across the world. The people, who are affected, include low-income earners and those who have tainted credit history for bad debt. These sections of people are excluded from the mainstream banking services because they do not have sufficient income to repay the loan or to keep asset as collateral for taking loan [e.g., Satya and Rupayan (2010), World Bank (2005)]

The primary objective of this paper is to empirically investigate the effects of financial inclusion principally through formal banking system on growth in overall real GDP per capita as one of the two goals (price stability and maximum economic growth) of US monetary policy. The vector of financial inclusion for this study includes number of commercial bank branches, number of other financial institutions, outstanding loans from commercial banks and outstanding deposits with commercial banks.

This is an issue of great importance motivating this study because lessons drawn from it have profound policy implications for many developed and developing countries. Moreover, such a rigorous study on this issue is in paucity, to our knowledge. To our belief, this study will definitely contribute to the existing body of related literature.

The remainder of the paper proceeds as follows. Section II briefly reviews related literature. Section III outlines the empirical methodology. Section IV reports results. Section V concludes with brief policy implications.

II. Brief Review of Related Literature

Policies for financial inclusion are vital components of majority of government strategies for inclusive sustainable growth all over the world without hurting the environment and compromising the future. Inclusiveness is an essential ingredient of any successful growth strategy. Three pillars of inclusive growth are: (1) maximize economic opportunities; (2) ensure economic well-being; and (3) ensure equal opportunities to economic opportunities. An inclusive growth strategy encompasses the key elements of an effective poverty reduction strategy and, more importantly, expands the development agendas. Developing inclusive financial systems which are financially and socially sustainable, as a poverty reduction strategy, should be given priority (Amit Bhandari, 2009). Beck, Demirgüç-Kunt and Levine (2007) have noticed a positive effect of financial inclusion on poverty reduction. Economies with higher levels of financial development experience faster reduction of poverty. This has been explained by an extensive body of empirical studies (e.g., White and Anderson (2001), Bourguignon (2003)). In an often cited cross-country study by Kraay (2004) proves that growth in average income explains 70 percent of the variation in poverty reduction (as measured by the headcount ratio) in the short-run, and as much as 97 percent in the long-run. Lopez and Servén (2004) suggest that the higher the inequality intensity, the poorer is the country. Higher per capita real GDP growth is vital to poverty reduction. This underscores the importance of promoting equitable growth through financial inclusion.

Importance of financial inclusion arises from the problem of financial exclusion of nearly 2.5 billion people

from the formal financial services across the world. According to World Bank (2005), financial exclusion includes four vital areas: savings, credit, transaction banking, and insurance. Broadly, financial exclusion can be defined as the inability to access basic financial services owing to complications accompanying with access, conditions, prices, marketing or self-exclusion in response to unfavorable experiences or perceptions of individuals/entities. The sections that are generally excluded are: marginal farmers, landless laborers, unorganized sector, urban slum dwellers, migrants, ethnic minorities and women. Some of the reasons for exclusion include lack of awareness, low income, social exclusion, illiteracy, sparse population in rural areas with poor infrastructure and lack of physical access, easy availability of informal credit, documenting procedures requiring proof of identity and address, high charges and penalties, etc. In short, there is no single over-riding factor that could explain financial exclusion. A variety of factors lead to financial exclusion, as stated above, and probably much more.

Another strand of literature explains financial exclusion also in the context of a larger issue of social exclusion of weaker sections of the population. While Leyshon and Thrift (1995) explain financial exclusion as such processes those aid to prevent some social groups and individuals from getting access to the formal financial system. Carbo et al. (2005) and Conroy (2005) opine that it is a state of inability of some poor and disadvantaged societal groups to access the financial system. Mohan (2006) reasons that financial exclusion implies the lack of access by some segments of the society to suitable, low-cost, fair and secure financial products and services from mainstream providers of financial services. Financial exclusion occurs mostly to people who are the disadvantaged sections of the society.

Another issue of interest is whether low level of financial inclusion is associated with high income inequality (Kempson et al., 2004). Beck et al. (2007) have examined financial sector outreach and its factors by employing cross-country data. Even, in the developed economies too, studies have revealed that the exclusion from the financial system occurs to low-income groups, the ethnic minorities, immigrants and others (e.g., Barr, 2004; Kempson and Whyley, 1998; Connoly and Hajaj, 2001). Studies by Leyshon and Thrift (1995) highlight that the geographical factor that people living in rural areas and in locations that are remote from financial centers are more likely to be financially excluded. As such, countries with low

levels of income inequality tend to have relatively high level of financial inclusion (Kempson and Whyley, 1998). In other words, the levels of financial inclusion inevitably rise in response to both prosperity, as measured by rise in per capita real GDP growth, and declining income inequalities. Another factor that can be related with financial exclusion is unemployment (Goodwin, et al. 2000). Evidence also suggests that the continued payment in cash is significantly related to financial exclusion (Kempson and Whyley, 1999). Informal sector accounts for a substantial share of unemployment in several less developed countries (ILO, 2002) which do not facilitate the process of financial inclusion. Formal employment entails inclusion and hence, the proportion of formal sector employment would be a key indicator of the degree of financial inclusion.

III. Empirical Methodology

Panel data as a combination of cross-sectional and time series observations are used in this study. This provides a convenient way to study phenomenon where a statistically adequate number of cross-sectional and time series observations are not obtainable. This augments quality and quantity of data. Otherwise, it would be impossible to use only one of these two dimensions for meaningful analyses (Gujarati, 2003). This study provides an example of such situation where incorporating observations on the variables over successive time periods allows to expand the informational content of the data. Furthermore, since the length of the time series is small compared to the number of cross-sections, the effects of autocorrelation are small if not negligible. Panel data estimation models include the constant coefficient (pooled), the fixed effects and the random effects regression models.

In order to test for the existence of a long-run equilibrium relationship among variables in heterogeneous panel consisting of 50 states of the USA and the DC over the period 1997-2013, the following model is specified:

$$y_{it} = \alpha_i + \beta_i x_{it} + \gamma_t D_{it} + e_{it} \quad (1)$$

Where, y = per capita real GSP (Gross State Product)

growth and x = vector of explanatory variables. $i=1, \dots, N$ and $t= 1, \dots, T$. The panel data set thus has altogether $N \times T$ observations.

In model (1), α_i shows the possibility of country fixed effects and β_i allows for heterogeneous cointegrating vectors. γ_t represents time-dependent common shocks, captured by common-time dummies (D_{it}), that might simultaneously affect all the 50 states and the DC included in this study. Model (1) is to be estimated by the proposed Pedroni (2000, 2001) panel Fully- Modified Ordinary Least Squares (FM-OLS) cointegration technique, which adjust for the presence of endogeneity and serial correlation in the data. This method is an appropriate technique, especially if there are endogenous macroeconomic factors that can cause co-movements in the above variables.

Before estimating model (1), it is required that the order of integration of the variables is determined by using panel unit root tests. If all variables are found to be I (1), then by using the Pedroni panel cointegration tests (1999, 2000, 2001), it is investigated whether they are co-integrated. These above mentioned tests and techniques are warranted to make sure that no spurious regression phenomenon exists in the estimation of β_i . In order to test for the presence of a unit root in the panel data series under study, panel unit root tests proposed by Im, Peseran and Shin (1997, 2003); Hadri (1999); Levin, Lin and Chu (2002) and Breitung (2000) are employed. For all these tests, the null hypothesis is non-stationarity. The rejection of the null hypothesis of no cointegration requires that the absolute values of the calculated test statistics exceed the respective critical values.

Subsequently, the following panel vector error-correction model in the spirit of (Engle and Granger, 1987) is estimated on the evidence of cointegrating relationship among variables of interest:

$$\Delta y_{it} = \alpha + \sum_{q=1}^k \beta \Delta y_{it-q} + \sum_{q=1}^l \phi \Delta x_{it-q} + \pi \hat{e}_{it-1} + \mu_{it} \quad (2)$$

Here, y = per capita real Gross State Product (GSP) growth and x = vector of explanatory variables.

For long-run convergence and causal relationship, the

estimated coefficient ($\hat{\pi}$) of the error-correction term (\hat{e}_{it-1}) is expected to be negative. The associated t-value indicates its statistical significance. The estimated β , and ϕ reveal short-run interactive feedback relationships. The appropriate lag-lengths are determined by the Akaike (1969) information criterion.

This study also invokes Generalized Method of Moments (GMM), as developed in Hansen (1982), for robust and efficient estimates. GMM is one of the most widely used econometric tools in finance. A set of moment conditions is used to estimate model parameters by GMM. In general, the number of moment conditions is larger than the number of model parameters. A model misspecification for over-identifying restrictions can be tested by GMM J-statistic. GMM does not require strong distributional assumptions for applications in finance. Since this paper employs panel data, GMM dynamic panel estimation is more appropriate than the original GMM estimation. On differencing of the regression equation, unobserved country/state- specific effects and the use of differenced lagged regressors eliminate parameter inconsistency arising from simultaneity bias (Arellano and Bond, 1991). Monte Carlo simulations of the model offer dramatic improvements in both efficiency and consistency (Blundell and Bond, 1997). The estimating regression equation in first-difference is specified as follows:

$$\Delta y_{it} = \alpha \Delta y_{it-1} + \beta' \Delta x_{it} + \Delta \varepsilon_{it} \quad (3)$$

Where, y = per capita real GSP growth and x = vector of weakly exogenous explanatory variables. Also, $\Delta \varepsilon_{it}$ is correlated with Δy_{it-1} . For estimation, the following moment conditions are used:

$$E[y_{it-q} (\Delta \varepsilon_{it})] = 0 \text{ for } q \geq 2; t = 3, \dots, T$$

$$E[x_{it-q} (\Delta \varepsilon_{it})] = 0 \text{ for } q \geq 2; t = 3, \dots, T$$

Annual data from 1997 through 2013 for US 50 states and the DC are available from the US Bureau of Labor Statistics, the US Department of Commerce, Financial Access Survey and the Federal Reserve of St. Louis.

IV. Results

LLC (Levine, Lin and Chu), Breitung, IPS (Im, Pesaran and Shin) and Hadri panel unit root tests are applied to determine non-stationarity in panel data series. The test results are reported as follows:

Table 1 reveals that the evidences, based on the above tests, are mixed for non-stationarity at different levels of significance in terms of the associated p-values. On first differencing of data, stationarity is restored in nonstationary variables, as shown in the lower segment of Table 1.

Subsequently, the Pedroni panel Cointegration testing is implemented consisting of a battery of seven tests. They are reported in Table 2 as follows:

Table 1. Panel Unit Root Tests

VARIABLES (LEVEL)	METHOD			
	LLC	Breitung	IPS	Hadri
GSP (Per Capita real GSP growth)	-13.4099 (0.0000)	-2.7045 (0.0034)	0.44662 (0.6724)	14.8754 (0.0000)
DEPT (Outstanding bank deposits)	9.2136 (1.0000)	20.7508 (1.0000)	15.1008 (1.000)	19.1128 (0.000)
LOAN (Outstanding bank loans)	-3.3376 (0.0000)	-1.4663 (0.0712)	-1.5172 (0.0646)	17.1575 (0.0000)
INST (Number of other financial institutions)	-10.5836 (0.0000)	-4.4221 (0.0000)	-11.6022 (0.0000)	5.2322 (0.0000)
BRAN (Number of bank branches)	8.2777 (1.0000)	7.7057 (1.0000)	5.3017 (1.0000)	5.7221 (1.0000)

VARIABLE (DIFFERENCE)	METHOD			
	LLC	Breitung	IPS	Hadri
GSP (Per Capita real GSP growth)	-7.9796 (0.0000)	-11.1810 (0.0000)	-18229 (0.0342)**	4.7664 (0.0000)
DEPT (Outstanding bank deposits)	-38.6553 (0.0000)	-35.0270 (0.0000)	-31.2214 (0.0000)	34.5030 (0.0000)
LOAN (Outstanding bank loans)	-5.7135 (0.0000)	-11.4106 (0.0000)	-5.8186 (0.0000)	-2.5828 (0.0000)
INST (Number of other financial institutions)	-8.3314 (0.0000)	-0.9863 (0.1664)	-5.0881 (0.0000)	13.2156 (0.0000)
BRAN (Number of bank branches)	-9.43413 (0.0000)	-7.9220 (0.0000)	1.07231 (0.8582)	11.0594 (0.0000)

Note: Associated P-values are reported in Parentheses.

Table 2. The Pedroni Panel Co-integration Test

Test	Constant trend	Constant + Trend
Panel v-Statistic	8.293318 (0.0000)*	10.08933 (0.0000)*
Panel rho-Statistic	2.322424 (0.0000)*	1.510235 -0.9345
Panel PP-Statistic	-9.10174 -0.9899	-21.6314 (0.0000)*
Panel ADF-Statistic	-8.78787 (0.0000)*	-18.5572 (0.0000)*
Group rho-Statistic	5.462206 -1	4.196788 -1
Group PP-Statistic	-9.00091 (0.0000)*	-22.3913 (0.0000)*
Group ADF-Statistic	-8.61581 (0.0000)*	-18.9308 (0.0000)*

Note: All reported values are asymptotically distributed as standard normal. Probability Statistics are within parentheses. * indicates rejection of the null hypothesis of no co-integration at 1% level of significance in terms of associated p-value.

As observed above, panel v-statistic, panel rho-statistic, panel ADF-statistic, group PP- statistic and group ADF-statistic with constant trend clearly reject the null hypothesis of no co-integration at 1% level of significance. Only panel PP-statistic and group rho-statistic suggest, otherwise. Again, with constant plus trend, the same tests confirm the above with regard to non-stationarity versus stationarity.

Finally, the panel vector error-correct model (2) is estimated. The estimates are presented as follows:

$$\begin{aligned} \Delta GSP_{it} = & 59.6116 + 0.1869\hat{e}_{it-1} + 0.5888\Delta GSP_{it-1} + 0.2176\Delta GSP_{it-2} \\ & (0.1011) \quad (-10.4781) \quad (17.8013) \quad (6.8586) \\ & + 2.06E-05\Delta DEPT_{it-1} + 5.78E-06\Delta DEPT_{it-2} + 5.08E-05\Delta LOAN_{it-1} \\ & (1.9547) \quad (0.5381) \quad (4.8124) \\ & + 2.55E-05\Delta LOAN_{it-2} + 7.0720\Delta INST_{it-1} - 6.5794\Delta INST_{it-2} \\ & (2.3689) \quad (1.4636) \quad (-1.9731) \\ & + 8.9180\Delta BRAN_{it-1} + 6.0742\Delta BRAN_{it-2} \\ & (11.600) \quad (12.2136) \end{aligned} \tag{2}$$

$$\bar{R}^2 = 0.4992, \quad F = 78.8473, \quad AIC = 22.3685$$

The coefficient of the error- correction term (\hat{e}_{it-1}) has expected negative sign that is overwhelmingly significant in terms of the associated t-value. This confirms convergence toward long-run equilibrium in the current change in per capita real GSP growth at slow speed since the numerical coefficient at (0.1869) is low. In terms of short-run dynamics, the net effect of deposit mobilization on the current change in real GSP growth per capita across

50 US states and the DC in terms of the sum of the lagged coefficients is positive but individually they are marginally significant in terms of the associated t-values. Likewise, the net effect of bank lending in the short-run is significantly positive in terms of the associated t-values. Furthermore, the net effect of the number of other financial institutions is marginally positive and statistically insignificant. The net effect of expansion in bank branches on per capita real GSP growth is highly positive and

statistically significant. \bar{R}^2 shows that nearly 50% of the change in per capita real GSP growth is explained by the changes in the lagged regressors. F-statistic is also highly significant. From the above findings, it is ostensibly clear that advancing bank-based financial inclusion through lending and branch expansion is highly important for all US states and the DC to boost per capita real GSP growth.

The estimates of model (3) are reported as follows:

Table 3. Method: Panel Generalized Method of Moments

Dependent Variable: ΔGSP_{it}

Variables	Coefficient	t-statistic	Prob.
ΔGSP_{it}	0.140327	2.80E+08	0
$\Delta LOAN_{it}$	0.385259	4.29E+09	0
$\Delta DEPT_{it}$	0.012108	NA	NA
$\Delta INST_{it}$	0.019571	NA	NA
$\Delta BRAN_{it}$	0.551858	4.90E+09	0
Effects Specification			
Cross-section fixed (orthogonal deviations)			
Mean dependent var	-0.03859	S.D. dependent var	0.03831
S.E. of regression	0.006093	Sum squared resid	0.028211
J-statistic	760	Instrument rank	15
Prob (J-statistic)	0		

As observed in Table 3, change in outstanding bank deposit and that in the number of other financial institutions across 50 US states and the DC have insignificantly positive effect on per capita real GSP growth. The effects of both outstanding bank loan and bank branches unleash highly significant positive influences on per capita real GSP growth. One percent increase in outstanding bank loan raises GSP growth by 0.38 percent. Likewise, one percent rise in the number of bank branches increases GSP growth by 0.55 percent. Thus, both dynamic panel cointegration estimates of model (2) and GMM dynamic panel estimates of model (3) lead to the same conclusion. However, differences are observed in magnitudes of coefficients and their associated respective t-values. GMM J-statistic at 760.00 also confirms no misspecification of the model.

V. Conclusions and Brief Policy Implications

Among the above four variables, greater access to bank credit and bank-branch expansion as means of financial inclusion greatly enhance economic well-being across fifty US states and the DC. So, the US monetary policy should attach greater importance to expanding further banking services to rural areas and economically disadvantaged segments of the population in addition to small businesses. They will add to consumer demand, GSP growth and employment. In closing, bank-based financial inclusion would promote equitable economic growth and household financial stability in the long-run resulting in narrowing income inequality and diminishing poverty. Expanding applications of modern information technologies and automation in this regard will be an added advantage. In closing, lessons drawn from this study may also apply to other countries.

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