Mohan Bijapur

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Do financial crises erode potential output? Evidence from OECD inflation responses

Mohan Bijapur*

UEL Royal Docks Business School, Docklands Campus, University Way, London E16 2RD, UK

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ABSTRACT

This paper presents evidence from a panel investigation of OECD countries that inflationary pressures tend to be stronger during recovery from financial crises compared to recovery from non-crisis economic downturns, indicating impairment in productive potential.

1. Introduction

As the global economy begins to recover from the most severe financial crisis since the Great Depression, concerns have been raised that the recent credit crunch has impaired the productive potential of the worst affected economies such that inflationary pressures may emerge as the economic recovery gathers pace. Although the recent empirical literature has for the most part focused on the impacts on aggregate demand, it has long been recognized that financial crises (and the associated shortages in the availability of bank credit) can also affect aggregate supply. Widespread credit rationing can constrain current production by restricting the availability of working capital for firms, and also reduce future production by constraining investment spending and the future capital stock (Blinder, 1987). These "supply" effects would manifest through decreasing the amount of spare capacity which normally opens up following economic downturns, such that inflationary pressures would be stronger than otherwise. In this paper, we test for evidence of these effects in the aftermath of 9 financial crises since 1990. We include the currently ongoing recoveries from the recent global financial crisis in our estimation sample in order for our results to be of greater relevance to current events, but it needs mentioning that these episodes represent incomplete cycles to date.

We estimate stylized Phillips curves across a sample of 11 OECD countries using panel data regressions, and test whether growth during recovery from financial crises tends to be more inflationary than during recovery from other economic downturns which are not associated with crisis events (henceforth referred to as "normal downturns"). Several recent papers have investigated the issue of whether financial crises constitute adverse supply shocks, by estimating their effects on long run or trend output (Barrell et al., 2010; Benati, 2012; Cecchetti et al., 2009; Cerra and Saxena, 2008; Furceri and Mourougane, 2009). We pursue a novel approach in terms of detecting an impact on inflation.

2. Definitions of financial crises, recessions and recoveries

We construct our sample using data on 11 countries (G7 plus Finland, Norway, Spain and Sweden), thereby including all major OECD economies plus several other countries which have experienced systemic banking crises in recent history. We begin our sample in 1990, in order to avoid the possibility of structural change and potential for model misspecification in a longer time series (for example, recent improvements in the efficiency of monetary policy (Cecchetti et al., 2006) and shifts in the slope of the Phillips curve (Borio and Filardo, 2007)).

Our definition of financial crises is derived from Laeven and Valencia (2010), who define a systemic banking crisis as "significant signs of financial distress in the banking system", accompanied by "significant banking policy intervention measures

* Tel.: +44 2082237798; fax: +44 2082233395.
E-mail address: m.bijapur@btinternet.com.

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in response to significant losses". Using their comprehensive and very current crisis resolution database we identify the following crises (start dates in brackets): Finland (1991), Norway (1991), Sweden (1991), Japan (1997), US (1988) and (2007), UK (2007), Germany (2008) and Spain (2008). We exclude those crises which the authors classify as "borderline" as it is less certain that these led to significant tightening in credit availability, with the following two exceptions: (1) Spain (2008), in light of the recent acceleration of the Spanish crisis culminating in the 2011 government restructuring plan (note also that Reinhart and Rogoff (2009) classify Spain as a systemic banking crisis), and (2) US (1988)—Bernanke and Lown (1991) and Bijapur (2010) both present evidence that the aftermath of the Savings and Loans crisis of the late 1980's was associated with a significant tightening in the availability of credit.

We characterize recessions by the adopting the familiar technical definition of 2 consecutive quarters of decline in real GDP (derived from the NBER's more comprehensive definition). On this basis, the following constitutes our sample of normal downturns (start dates in brackets): Canada (1990) and (2008), Finland (2008), France (1992) and (2008), Germany (1991) and (2000), Italy (1992), (2001) and (2008), Japan (1992), (2001) and (2008), Norway (2008), Spain (1992), Sweden (2008), UK (1990).

Finally, the recovery period starts from the first occurrence of either two consecutive quarters of positive growth, or average growth over 2 quarters turns positive, and lasts for 10 quarters. (In the case of "double-dip" recessions, the recovery period starts after the second dip ends). We believe this is a reasonable length of time to estimate the effects of interest, and extending it further risks the results becoming contaminated by other events. Our reasons for excluding the contraction period itself from the estimation subsample is that credit shortages experienced during crises are likely to have lagged effects (due to "time to build" investment lags and also lags with which stricter lending criteria transmit to actual lending volumes).

3. Data issues and model estimation

We estimate stylized Phillips curves, specified as the relationship between the change in the rate of inflation and growth in real output, using quarterly data. In line with much of the existing literature, we also include lagged commodity prices given their power in explaining future inflation. Given that the limited number of financial crisis episodes renders country-by-country estimation infeasible, we pool the data and use panel data methods in order to increase estimation efficiency. We use a variety of estimators in order to test for unobserved country-specific heterogeneity which may be correlated with the regressors and thus lead to biased estimates: (1) Pooled OLS (all parameters are assumed fixed and homogeneous across countries), (2) one-way Fixed Effects (pooled OLS with country-specific intercepts) and (3) Swamy Random Co-Efficients Model (RCM) (all parameters are random, i.e. realizations from a probability distribution with fixed mean and variance, and distributed independently of the regressors).

Our model specification is:

$$d\pi_{it} = \alpha_i + \sum_{n=1}^{N}[(\beta_n + u_{in}) + \delta_n^y F_{it} + \lambda_n^y R_{it}] * dy_{it-n} + \sum_{n=1}^{N}[(\gamma_n + v_{in}) + \delta_n^c F_{it} + \lambda_n^c R_{it}] * dc_{t-n} + \varepsilon_{it}$$

where $d\pi_{it}$ denotes second differenced log Consumer Prices Index (CPI) for country i on date t, dy_{it-n} and dc_{t-n} denote lagged first differenced log GDP (chained volume measures) and log commodity prices respectively, α_i denotes a fixed effect and u_{in} and v_{in} denote random effects for country i (included in some specifications). F_{it} and R_{it} are interactive dummies, set equal to 1 during the recovery periods following financial crises and normal downturns respectively (intercept dummies are also included). Finally, in order to isolate the impact of financial crises on inflation and control for heterogeneity in the size of the output drop preceding recovery across the two subsamples, we also include a control variable. This is set equal to the (absolute value of the) peak to trough decline in GDP (in per cent) preceding the recovery period.

CPI and GDP data are taken from the SourceOECD database, and commodity prices are measured using the IMF's All Primary Commodities spot price index (taken from the IFS database — prior to 1992q1 we use oil prices (world average crude petroleum) due to data unavailability). We estimate over sample period 1990q1–2011q3 (for Germany, we begin in 1992q2 in order to exclude the re-unification break, and for Finland in 1991q2 due to data unavailability; for Japan we end in 2010q4 in order to exclude the effects of the March 2011 earthquake and tsunami).

We test for erosion of potential output in the aftermath of financial crises as follows. Under normal economic conditions, faster growth leads to accelerating inflation, thus the predicted signs of the lag co-efficients β_n are positive. However, following a recession a large amount of spare capacity opens up and hence the economy can grow faster without inflation taking off. Thus, the predicted signs of the interactive dummy co-efficients λ_n^y are negative. However, if a financial crisis leads to erosion of potential output, then the margin of spare capacity would be smaller, and hence a percentage point increase in the growth rate would lead to a larger increase in inflation relative to a normal downturn. Thus, in the presence of this effect, the dummy co-efficients δ_n^y would be greater (i.e. less negative) than λ_n^y. Formally, we test the null $H0 : \sum_{n=1}^{N} \delta_n^y - \sum_{n=1}^{N} \lambda_n^y \leq 0$ against the alternative $H1 : \sum_{n=1}^{N} \delta_n^y - \sum_{n=1}^{N} \lambda_n^y > 0$.

Augmented Dickey–Fuller (ADF) tests for unit roots indicated that all variables were stationary. We included 4 lags of the GDP and commodity price variables (Wooldridge (2002)'s panel autocorrelation test indicated this was sufficient to avoid dynamic misspecification). The White test for heteroscedasticity was significant, hence White-robust standard errors were used wherever possible. Time dummies were inserted in order to exclude outliers in the CPI data due to indirect tax changes (Canada: 1991q1–q2, Germany: 1993q1–q2, Japan: 1997q2-q3, Sweden: 1991q1–q2, 1993q1–q2, UK: 1991q2–q3, 1993q2–q3, 2011q1), seasonality (Norway, Spain (post-2001), Sweden, UK), and other erratic factors (all countries: spike in fruit and vegetable prices (2001q2–q3)).

4. Results

Table 1 reports the main results of the panel regressions. In all three specifications the results show supportive evidence that growth is more inflationary during the aftermath of financial crises compared to other downturns. Consistent with economic theory, the sum of co-efficient estimates on the output growth lags for the whole sample is positive. The sum of growth dummy co-efficient estimates in the normal downturn subsample displays the predicted negative sign (although not statistically significant), indicating that the large amount of spare capacity created by recessions enables growth to occur with a more muted inflation response. However, in the financial crisis subsample, the equivalent estimate is actually positive (although not statistically significant), suggesting that inflation is more responsive to growth during recovery from financial crises, holding constant other factors. The difference in magnitude of the dummy co-efficients between financial crises and normal downturns is statistically

Table 1
Estimation results. We report the co-efficient estimates for four model specifications. Columns (1)–(3) report the results of the pooled OLS, Fixed Effects and pooled OLS assuming heterogeneity in β_4 respectively. (4) reports the results in which the German crisis is excluded.

Explanatory variable	(1)	(2)	(3)	(4)
Constant	−.001	−.001***	−.001	−.001***
GDP(−1)	.039	.041**	.043	.041**
GDP(−2)	−.001	.003	.010	.003
GDP(−3)	.061	.071**	.074	.071**
GDP(−4)	−.022	−.013	−.040[a]	−.014
COMMOD(−1)	−.002	−.003	−.002	−.003
COMMOD(−2)	−.014***	−.014***	−.014***	−.014***
COMMOD(−3)	.001	.001	.000	.001
COMMOD(−4)	−.006*	−.006*	−.006	−.006*
Output Drop	.011	.015	.013	.014
Financial crisis dummies:				
Intercept	−.001	−.001**	−.002	−.002***
GDP(−1)	.048	.048	.056	.086**
GDP(−2)	.033	.033	.029	.019
GDP(−3)	−.063	−.072	−.063	−.098
GDP(−4)	.078	.065	.109	.110
COMMOD(−1)	−.008	−.005	−.009	−.009
COMMOD(−2)	.008	.007	.007	.012
COMMOD(−3)	−.007	−.005	−.007	−.008
COMMOD(−4)	.009*	.008	.009	.011
Normal downturn dummies:				
Intercept	.001	.000	.001	.000
GDP(−1)	−.065	−.050	−.075	−.051
GDP(−2)	.023	.030	.010	.029
GDP(−3)	−.069	−.061	−.070	−.062
GDP(−4)	−.024	−.026	−.018	−.025
COMMOD(−1)	−.004	−.004	−.001	−.004
COMMOD(−2)	.002	−.000	.001	−.000
COMMOD(−3)	.007	.006	.008	.007
COMMOD(−4)	.003	.000	.001	.000
Adjusted R^2	.425	.443	.428	.438
(A) Sum of financial crisis GDP dummies	.096	.075	.131	.118**
(B) Sum of normal downturn GDP dummies	−.135	−.107	−.153	−.109
(A)–(B)	.231**	.182**	.284**	.226**

[a] Mean estimate across all countries.
* Indicates significance of a Wald test (one-tailed for the bottom row only) at the 10% level.
** Indicates significance of a Wald test (one-tailed for the bottom row only) at the 5% level.
*** Indicates significance of a Wald test (one-tailed for the bottom row only) at the 1% level.

significant in all three specifications (reported in the bottom row of Table 1).

These results indicate that the margin of spare capacity tends to be smaller following financial crises relative to other downturns. Inclusion of the variable controlling for the size of the output drop suggests that this is caused by an erosion of potential output, as opposed to the preceding recession being shallower in crisis episodes compared to other downturns (in fact, in our sample the opposite is true). The positive sign on the co-efficient estimate is somewhat puzzling (although it is not statistically significant), as one would expect a larger output drop to lead to weaker inflation. However, this could be explained by the fact that the output drop variable is significantly correlated with the financial crisis variables (financial crises tend to be associated with deeper recessions), thus potentially giving rise to a problem of multicollinearity.

The parameter estimates remain fairly stable across all three panel model specifications, with the signs of the dummy co-efficients unchanged for the crisis and non-crisis subsamples, and the difference in magnitude remaining similar. There is little evidence of fixed effects — although the restriction of homogeneity on country-specific intercepts is rejected at the 1% significance level, the Hausman test strongly rejects fixed effects in favour of random effects. In order to investigate the possibility that our results might instead be reflecting heterogeneity in the output-inflation trade-offs across countries, we also estimate using the Swamy RCM specification. However, the estimated random effects show very little dispersion around the fixed means for all regressors, with the exception of the fourth lag of GDP growth. Re-estimating with (fixed and heterogeneous) country-specific parameters for this variable, we find the results are largely unaffected, although the difference in magnitude of the dummy co-efficients actually increases in significance.

We also investigate whether our results are being disproportionately driven by an isolated crisis episode, by dropping each crisis in turn and re-estimating (for brevity, we report results only for the Fixed Effects specification although all three specifications yield similar results). The dummy co-efficients remain relatively stable, with the difference in magnitude ranging from 0.153 to 0.226. An interesting finding is that unlike the other recent crisis episodes, the estimates actually increase in significance when the German banking crisis is omitted, tentatively indicating that Germany did not experience a significant impact on productive potential.

5. Conclusions

This paper sheds light on the widely debated issue of whether financial crises constitute adverse supply shocks and thus lead to impairment in an economy's productive potential. We follow a novel approach by investigating whether inflationary pressures tend to be stronger in the aftermath of financial crises compared to non-crisis economic downturns. By estimating stylized Phillips curves in an OECD panel, we find that growth experienced during the aftermath of financial crises generates a significantly

larger increase in inflation relative to normal downturns, holding constant other relevant factors, thus pointing to impairment in productive potential.

It needs mentioning that the evidence we find for a stronger inflation response might be driven in part by other factors which are characteristic of financial crises, such as the inflationary impact of central bank balance sheet expansion (as pointed out in Cecchetti et al., 2009). However, an attempt to disentangle the specific mechanisms at work is beyond the scope of the present paper and is left for future research. Nevertheless, the policy implications are clear — as the recovery gathers pace, policy might need to be tightened more quickly and more aggressively than would otherwise be the case.

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