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# DYNAMIC RELATIONSHIP BETWEEN INFLATION AND FINANCIAL DEVELOPMENT

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This paper studies the long- and short-run relationships between inflation and financial development. Applying the Pooled Mean Group estimator of Pesaran, Shin, and Smith (1999, *Journal of the American Statistical Association* 94, 621–634) to unbalanced panel data for 87 countries over the period 1960–2005, we find that a negative long-run relationship between inflation and financial development coexists with a positive short-run relationship. However, when the data are split into different income or inflation groups, these results can be observed only in low-income or low-inflation economies.

**Keywords:** Inflation, Financial Development, Pooled Mean Group Estimator

## 1. INTRODUCTION

An extensive literature in the field of economic development concludes that financial development has not only a significant growth-enhancing effect,[1] but also an inequality-reducing effect.[2] By alleviating informational asymmetries and transactions costs and disproportionally relaxing financial constraints on the poor, financial development encourages more productive investment in physical and human capital that substantially facilitates economic growth, narrowing the gap between the poor and the rich. Moreover, because credit market frictions can propagate and amplify business cycle fluctuations, financial development helps mitigate macroeconomic volatility.[3] Understanding what causes financial development is important because of the welfare improvement effects on these important macroeconomic issues.

On the other hand, a substantial theoretical and empirical literature recognizes that high and sustained inflation is detrimental to economic growth[4] and worsens

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wealth inequality.[5] As argued in Smith (2003), a possible candidate through which inflation brings about real consequences is the effects on the financial system, especially those from damaging financial markets or impeding their operations. Therefore, unlike existing literature emphasizing legal environments or globalization as the driving force of a well-functioning financial system,[6] this paper examines whether financial development is associated with the level of inflation. The finding of significant impacts of inflation on financial development would have important policy implications for real economic activity. In particular, it would suggest that inflation exerts its influence on economic performance through a financial mechanism.

Despite considerable studies, the exact link between inflation and financial development is still far from being uncontroversial. According to Mundell (1963) and Tobin (1965), inflation causes portfolio allocations away from money into capital that leads to lower real returns on capital and higher investment, with positive effects on economic growth. English (1999) takes a step further and argues that higher inflation leads households to substitute purchased transaction services for money balances, which increases the production of financial services and boosts the size of the financial services sector. On the other hand, theoretical models based on imperfect credit markets postulate that when there are information-type credit market frictions whose severity is endogenous, higher rates of inflation create greater credit rationing and distort the flows of information, thereby exacerbating credit market frictions. Furthermore, high inflation can repress financial intermediation by eroding the usefulness of money assets and by leading to policy decisions that distort the financial structure. Thus, an increase in inflation may interfere with the ability of financial sectors to allocate resources, thereby reducing capital accumulation and economic growth. These models also emphasize that only when inflation exceeds some threshold levels do informational frictions necessarily play a substantial role.[7]

On the empirical side, English (1999) provides cross-country evidence in support of a positive effect of inflation on the size of the financial sector, whereas Haslag and Koo (1999), Boyd et al. (2001), and Khan et al. (2006) show a negative and even nonlinear relationship between inflation and financial development. Haslag and Koo (1999) argue that inflation is associated with financial repression and find a negative relationship between inflation and financial development, but the relationship disappears with increases in the inflation rate above a threshold. Boyd et al. (2001) obtain similar results and show that such an inflation threshold occurs at 15% per year. By contrast, Khan et al. (2006) find that the threshold level of inflation is about 3%–6% and that, for rates above the threshold level, increases in inflation have strongly negative effects on financial development.

In addition to long-run effects, short-run considerations may also play a role in the relationship. Mankiw (1989) points out that inflation tends to rise in booms and fall in recessions in the absence of identifiable real shocks such as oil price changes. Moreover, because financial development is characterized not only by long-run financial deepening but also by short-run financial instability [Loayza

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and Ranciere (2006)], and because the risk of bank crises tends to be higher in high-inflation environments, higher inflation may foster financial fragility that has short-run implications for real activity [Boyd and Champ (2003)].

In particular, the literature on debt deflation suggests that unanticipated deflation increases the burden of nominal debt by redistributing wealth from debtors to creditors,[8] worsens borrowers' net worth, and causes debtors to take on risk and default on loans, which lead to a decline in the nominal value of assets on bank balance sheets. This deterioration in bank assets threatens insolvency and hence restricts the credit supply [Fisher (1933); Bernanke (1983)].[9] Alternatively, according to the hypothesis of credit channels, that is, mechanisms of propagation of primitive shocks such as technology, taste, and monetary shocks, a negative wealth shock (such as an unanticipated drop in asset prices) damages the borrowers' balance sheet and reduces the amount of credit that borrowers can get, as a firm's net worth plays a role similar to that of collateral [Bernanke and Gertler (1989, 1990, 1995); Kiyotaki and Moore (1997)]. Because the amount of redistribution that can be triggered even by moderate inflation is sizable [Doepke and Schneider (2006)], and because small shocks can induce large effects on credit and hence business cycles [Bernanke et al. (1999)], the debt-deflation/balance sheet channel implies that there may be a positive association between inflation and financial development in the short run.

Accordingly, econometric assessments of the relationship between inflation and financial development should ideally be capable of uncovering the relevant long-run parameters as well as the short-run link between the two variables. To account for theoretically conflicting contributions and to advance previous research, this paper models the inflation–finance nexus using a panel technique that explicitly distinguishes between short- and long-run effects of inflation on financial development. This can be accomplished by specifying an autoregressive distributed lag (ARDL) model for each country, pooling them together in a panel, and then testing the cross-equation restriction of a common long-run relationship between the two variables using the Pooled Mean Group (PMG) estimator of Pesaran et al. (1999). Such a country-specific ARDL structure allows not only accommodating cross-country heterogeneity in the degree of credit market imperfections and in policy regimes, but also capturing certain interesting time-series relations that cross-sectional analysis cannot deal with. Moreover, this methodology can be applied to either stationary or nonstationary variables and hence does not require pretesting of unit roots. This circumvents some problems in cointegration analysis that focuses only on the estimation of long-run relationships among nonstationary variables, as well as the low power of unit-root tests against plausible alternatives. Further, instead of averaging the data per country to isolate trend effects,[10] both long- and short-run relationships are estimated using a panel of data that pool time-series and cross-sectional effects.[11]

Using panel data pooled from 87 developed and developing countries for the period 1960–2005, we find evidence of a strong link between inflation and financial development, irrespective of alternative financial development measures, control

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variables, and inflation uncertainty proxies. Specifically, higher inflation appears to stymie financial development in the long run but stimulate financial activity in the short run. In splitting the data into different income or inflation groups, we find a very robust negative long-run effect of inflation on financial development, whereas the positive short-run finance effect of inflation can be observed only in low-income or low-inflation economies. In addition, the long-run impact is generally larger than the short-run effect.

The remainder of the paper is organized as follows. Section 2 introduces the PMG estimator proposed by Pesaran et al. (1999). Section 3 describes the data and source, and Section 4 reports empirical results and robustness tests. Section 5 concludes the analysis.

## 2. THE AUTOREGRESSIVE DISTRIBUTED LAG APPROACH

To examine the long-run effect of inflation on financial development, it is common to estimate the following cross-sectional regression:

$$\text{finance}_i = \alpha + \beta\,\text{inflation}_i + \omega\,\text{controls}_i + \varepsilon_i, \tag{1}$$

where finance is the financial development indicators, inflation is the inflation index, controls is a set of control variables, $i = 1, 2, \ldots, N$ is the country index, and $\varepsilon$ is the error term.

To allow rich dynamic heterogeneity in the inflation–finance nexus over time and across countries, we nest equation (1) in an ARDL specification where the dependent and independent variables enter the right-hand side with lags of order $p$ and $q$, respectively:

$$y_{it} = \mu_i + \sum_{j=1}^{p} \lambda_{ij} y_{it-j} + \sum_{j=0}^{q} \delta_{ij}^{'} x_{it-j} + \varepsilon_{it}, \tag{2}$$

where $t = 1, 2. \ldots, T$, $y_{it} = \text{finance}_{it}$, $x_{it} = (\text{inflation}_{it}, \text{controls}_{it})$, and $\mu_i$ is the fixed effects.

By reparameterization, equation (2) can be written as

$$\Delta y_{it} = \mu_i + \phi_i y_{it-1} + \beta_i^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta x_{it-j} + \varepsilon_{it}, \tag{3}$$

where

$$\phi_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right), \quad \beta_i = \sum_{j=0}^{q} \delta_{ij},$$

$$\lambda_{ij}^{*} = -\sum_{m=j+1}^{p} \lambda_{im}, \quad j = 1, 2, \ldots, p-1,$$

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and

$$\delta_{ij}^{*} = -\sum_{m=j+1}^{q} \delta_{im}, \quad j = 1, 2, \ldots, q-1.$$

By further grouping the variables in levels, equation (3) can be rewritten as

$$\Delta y_{it} = \mu_i + \phi_i(y_{it-1} - \theta_i^{'} x_{it}) + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta x_{it-j} + \varepsilon_{it}, \qquad (4)$$

where $\theta_i = -(\beta_i/\phi_i)$ defines the long-run or equilibrium relationship among $y_{it}$ and $x_{it}$. $\lambda_{ij}^{*}$ and $\delta_{ij}^{*}$ are the short-run coefficients relating financial development to its determinants $x_{it}$. Finally, $\phi_i$ measures the speed of adjustment of $y_{it}$ toward its long-run equilibrium following a change in $x_{it}$. $\phi_i < 0$ ensures that such a long-run relationship exists. As a result, a significant and negative value of $\phi_i$ is treated as evidence in support of cointegration between $y_{it}$ and $x_{it}$.

As argued in Catao and Solomou (2005) and Catao and Terrones (2005), the ARDL specification in equation (4) not only mitigates the contemporaneous feedback running from financial development to inflation, but also accommodates the substantial persistence of finance adjustments and captures potentially rich inflation adjustment dynamics. In addition, the model allows for heterogeneity in the relationship between financial development and inflation across countries because the various parameters in equation (4) are not restricted to be the same across countries. Moreover, the dynamic panel approach allows us to estimate an empirical model that encompasses the long- and short-run effects of inflation on financial development using a data field composed of a relatively large sample of countries and annual observations.

There are a few existing procedures for estimation of the above model. At one extreme, the pooled estimator assumes the fully homogeneous–coefficient model in which all slope and intercept parameters are restricted to be identical across countries. At the other extreme, the fully heterogeneous–coefficient model imposes no cross-country coefficients constraints and can be estimated on a country-by-country basis. This is the mean group (MG) estimator introduced by Pesaran and Smith (1995). The approach amounts to estimating separate ARDL regressions for each group and obtaining $\theta$ and $\phi$ as simple averages of individual group coefficients $\theta_i$ and $\phi_i$. In particular, Pesaran and Smith (1995) show that the MG estimator will provide consistent estimates of the average of parameters interested.

Between these extremes, the dynamic fixed-effect (DFE) method allows the intercepts to differ across groups but imposes homogeneity of all slope coefficients and error variances. Alternatively, Pesaran et al. (1999) propose the PMG estimator, which restricts the long-run parameters to be identical over the cross section, but allows the intercepts, short-run coefficients (including the speed of adjustment), and error variances to differ across groups in the cross section. If the long-run homogeneity restrictions are valid, it is known that MG estimates are inefficient. In this case, the maximum likelihood–based PMG approach proposed

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by Pesaran et al. (1999) yields a more efficient estimator.[12] As shown in Pesaran et al. (1999), the validity of a cross-sectional, long-run homogeneity restriction of the form $\theta_i = \theta, i = 1, 2, \ldots, N$, and hence the suitability of the PMG estimator can be tested by a standard Hausman-type statistic.

In terms of the relationship between inflation and financial development, the PMG estimator offers the best available compromise in the search for consistency and efficiency. This estimator is particularly useful when the long run is given by conditions expected to be homogeneous across countries,[13] whereas the short-run adjustment depends on country characteristics such as monetary and fiscal adjustment mechanisms, capital market imperfections, average rate of inflation, and relative price and wage flexibility [e.g., Loayza and Ranciere (2006)]. Therefore, we use the PMG approach to estimate a long-run relationship that is common across countries while allowing for unrestricted country heterogeneity in the adjustment dynamics.

## 3. DATA DESCRIPTION AND SOURCES

Our dataset consists of a panel of 87 countries over the period 1960–2005 and is mainly taken from the World Development Indicator (2006) published by the World Bank. Data on financial development are obtained from the Financial Structure Database originally compiled by Beck et al. (2000). Inflation (*inf*) is calculated as the percentage change in the Consumer Price Index. We use three bank-based financial development indicators: Private Credit (*lprivo*), Liquidity Liabilities (*llly*), and Bank Assets (*ldby*).[14] Private Credit is the value of credits by financial intermediaries to the private sector, divided by GDP. It is Beck et al.'s (2000) preferred measure because it excludes credit granted to the public sector and credit issued by the central bank and development banks. Liquidity Liabilities is equal to the sum of currency and demand and interest-bearing liabilities of banks and nonbank financial intermediaries, divided by GDP. This is a common measure of financial depth, although it might involve double-counting and it includes liabilities backed by credits to the public sector. And Bank Assets is defined as the domestic assets of deposit money bank as a share of GDP. Thus, Bank Assets measures the degree to which domestic banks allocate society's savings.

To strengthen our empirical results, we control for conditioning variables in the relation between inflation and financial development. The variables include the initial real per capita GDP (*initial*) to control for a causal link from the income level to financial development, the ratio of government expenditure to GDP (*lgov*) to measure macroeconomic stability, and the sum of exports and imports as a share of GDP (*ltrade*) to account for external shocks.[15]

Table 1 displays the list of countries in the sample, and Table 2 provides descriptive statistics and correlations of the variables over the period 1960–2005. It is noticed that inflation and each of three financial intermediary development measures are negatively correlated, and that the correlations between any pair of three financial development measures are positive.

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**TABLE 1.** List of sample countries by inflation

| High-inflation countries | Medium-inflation countries | Low-inflation countries |
|---|---|---|
| Bolivia | Burundi | Australia |
| Chile | Costa Rica | Austria |
| Ecuador | Côte d'Ivoire | Belgium |
| Ghana | Dominican Republic | Belize |
| Iceland | Egypt | Burkina Faso |
| Indonesia | Ethiopia | Canada |
| Israel | Fiji | Central African |
| Jamaica | Gambia, The | Chad |
| Malawi | Greece | Cyprus |
| Mexico | Grenada | Denmark |
| Nigeria | Guatemala | Dominica |
| Sierra Leone | Haiti | Finland |
| Suriname | Honduras | France |
| Uganda | Hungary | Germany |
| Uruguay | India | Japan |
| Venezuela, RB | Iran | Luxembourg |
| Zambia | Ireland | Malaysia |
| | Italy | Morocco |
| | Jordan | Netherlands |
| | Kenya | Niger |
| | Korea, Rep. | Norway |
| | Madagascar | Panama |
| | Mauritius | Senegal |
| | Nepal | St. Kitts and Nevis |
| | New Zealand | St. Vincent and the Grenadines |
| | Pakistan | Sweden |
| | Paraguay | Switzerland |
| | Philippines | Thailand |
| | Portugal | United States |
| | Rwanda | |
| | Seychelles | |
| | South Africa | |
| | Spain | |
| | Sri Lanka | |
| | St. Lucia | |
| | Swaziland | |
| | Syrian Arab Republic | |
| | Togo | |
| | Tonga | |
| | Trinidad and Tobago | |
| | United Kingdom | |

*Note*: Countries with inflation of 15% and above averaged over the sample period are classified as high-inflation countries, those with inflation between 6% and 15% as medium-inflation countries, and those with inflation of 6% and below as low-inflation countries.

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**TABLE 2.** Descriptive statistics (1960–2005)

| | *inf* | *lprivo* | *ldby* | *llly* | *initial* | *lgov* | *ltrade* |
|---|---|---|---|---|---|---|---|
| | | | Panel A: Summary statistics | | | | |
| Mean | 15.6337 | 3.3272 | 3.4400 | 3.6154 | 7.6207 | 2.6466 | 4.0382 |
| Std. | 199.9059 | 0.9436 | 0.8497 | 0.6337 | 1.6359 | 0.3799 | 0.5712 |
| Min. | −13.0566 | −0.3071 | 0.3447 | 1.5327 | 4.4242 | 0.9502 | 1.8438 |
| Max. | 11,749.6400 | 5.8438 | 5.5546 | 5.8236 | 10.8305 | 3.9985 | 5.6832 |
| | | | Panel B: Correlation matrix | | | | |
| *inf* | 1.0000 | | | | | | |
| *lprivo* | −0.0628 | 1.0000 | | | | | |
| *ldby* | −0.0782 | 0.9155 | 1.0000 | | | | |
| *llly* | −0.0873 | 0.8364 | 0.8983 | 1.0000 | | | |
| *initial* | −0.0209 | 0.7551 | 0.7405 | 0.6859 | 1.0000 | | |
| *lgov* | −0.0184 | 0.4129 | 0.4859 | 0.4345 | 0.4496 | 1.0000 | |
| *ltrade* | −0.0163 | 0.2784 | 0.3439 | 0.3246 | 0.2384 | 0.3717 | 1.0000 |

## 4. EMPIRICAL RESULTS

### 4.1. Basic Results

Table 3 displays the results on specification tests and the estimation of long- and short-run parameters linking inflation and financial development.[16] We emphasize the results from using the PMG estimator, considering its gains in consistency and efficiency over other panel error-correction estimators. For comparison, we also present the results obtained by the MG and DFE estimators.

For a long-run relationship to exist (dynamic stability), the coefficient on the error-correction term should be negative and within the unit circle. As can be seen in Table 3, the pooled error-correction coefficient estimates are significantly negative and fall within the dynamically stable range for PMG, MG, and DFE estimators. This indicates a mean reversion to a nonspurious long-run relationship and therefore stationary residuals, meaning that inflation and financial development are cointegrated. In addition, the Hausman test does not reject long-run homogeneity restriction, implying that the PMG estimator is more suitable than MG for the analysis. These results hold for alternative financial development measures. Accordingly, the following analysis focuses on the PMG approach.

Regarding the estimated parameters of primary interest, we find that the long-run coefficient of inflation is negative and significant. This suggests that inflation tends to hinder financial development in the long run. However, the short-run coefficients on inflation tell a different story. Because the price regimes and capital market frictions vary across countries in the short run, the short-run coefficients are not restricted to be the same across countries, so we do not have a single pooled estimate for each coefficient. Nevertheless, we can still analyze the average short-run effect by considering the mean of the corresponding coefficients across countries. As Table 3 shows, the average short-run relationship between financial

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**TABLE 3.** The effect of inflation on financial development

| | PMG | MG | DFE | Hausman test |
|---|---|---|---|---|
| | | Panel A: *lprivo* | | |
| Long-run coefficients | | | | |
| Inflation | −0.0256*** | −0.0153 | 0.0004 | 0.1512 |
| | (0.0026) | (0.0267) | (0.0003) | [0.6974] |
| Error correction | | | | |
| Phi | −0.0637*** | −0.0763*** | −0.0546*** | |
| | (0.0073) | (0.0099) | (0.0046) | |
| Short-run coefficients | | | | |
| △ Inflation | 0.0015*** | 0.0015*** | 0.0000*** | |
| | (0.0005) | (0.0005) | (0.0000) | |
| Constant | 0.2435*** | 0.2806*** | | |
| | (0.0276) | (0.0353) | | |
| | | Panel B: *ldby* | | |
| Long-run coefficients | | | | |
| Inflation | −0.0237*** | 0.0120 | 0.0007** | 0.6020 |
| | (0.0024) | (0.0460) | (0.0003) | [0.4378] |
| Error correction | | | | |
| Phi | −0.0553*** | −0.0706*** | −0.0487*** | |
| | (0.0075) | (0.0108) | (0.0045) | |
| Short-run coefficients | | | | |
| △ Inflation | 0.0009** | 0.0008* | −0.0001*** | |
| | (0.0004) | (0.0005) | (0.0000) | |
| Constant | 0.2197*** | 0.2765*** | | |
| | (0.0306) | (0.0440) | | |
| | | Panel C: *llly* | | |
| Long-run coefficients | | | | |
| Inflation | −0.0122*** | −0.0060 | −0.0004* | 0.1192 |
| | (0.0015) | (0.0181) | (0.0002) | [0.7299] |
| Error correction | | | | |
| Phi | −0.0763*** | −0.0961*** | −0.0513*** | |
| | (0.0089) | (0.0120) | (0.0052) | |
| Short-run coefficients | | | | |
| △ Inflation | 0.0010* | 0.0011* | 0.0000*** | |
| | (0.0006) | (0.0006) | (0.0000) | |
| Constant | 0.3057*** | 0.3859*** | | |
| | (0.0368) | (0.0469) | | |

*Note*: The values in the parentheses (brackets) are the standard errors ($p$-value) of corresponding coefficient estimates. ***, ** and * Significant at 1%, 5%, and 10% level, respectively.

development and inflation appears to be significantly positive. That is, on the average, inflation is a significant driver of financial development.

The findings hold for alternative measures of financial development, and indicate that the inflation–finance relationship depends on whether their movements

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**TABLE 4.** The PMG results (robustness tests)

| | Financial development indicator | | |
|---|---|---|---|
| | *lprivo* | *ldby* | *llly* |
| Long-run coefficients | | | |
| Inflation | −0.0271*** | −0.0308*** | −0.0130*** |
| | (0.0031) | (0.0034) | (0.0014) |
| Income | 1.8578*** | 2.0460*** | 0.6547*** |
| | (0.1194) | (0.1229) | (0.0170) |
| Government | 0.6142*** | 0.4106*** | 0.5007*** |
| | (0.0943) | (0.0957) | (0.0278) |
| Trade | 0.3328*** | 0.3639*** | 0.4797*** |
| | (0.1053) | (0.0998) | (0.0353) |
| Error correction | | | |
| Phi | 0.0607*** | 0.0516*** | 0.1053*** |
| | (0.0080) | (0.0081) | (0.0125) |
| Short-run coefficients | | | |
| △ Inflation | 0.0013** | 0.0015*** | 0.0018** |
| | (0.0006) | (0.0005) | (0.0008) |
| △Income | 0.4505*** | 0.1689** | 0.0520 |
| | (0.0772) | (0.0781) | (0.0918) |
| △Government | 0.1698*** | 0.1916*** | 0.1331*** |
| | (0.0379) | (0.0360) | (0.0408) |
| △Trade | 0.0462** | 0.0539** | 0.0684*** |
| | (0.0254) | (0.0260) | (0.0221) |
| Constant | 0.7617*** | 0.5926*** | 0.4497*** |
| | (0.1130) | (0.1054) | (0.0602) |

*Note*: The values in the parentheses are the standard errors of corresponding coefficient estimates.
***, ** and * Significant at 1%, 5%, and 10% level, respectively.

are temporary or for a long haul. The association of inflation with financial development is found to be larger in the long run than in the short run in terms of magnitude. More importantly, the findings of coexistence of positive short-run effects and negative long-run effects imply that although the arguments of imperfect credit markets and financial repression tend to dominate in the long run, the debt-deflation/balance sheet mechanism can account for the positive association of inflation and financial development in the short run. A devaluation of existing debt resulting from surprise inflation may lower financial constraints of existing borrowers, which frees up collateral and makes new borrowing and lending possible.

To consider if the results are sensitive to model specification, we add three control variables into the models: income, government size, and trade openness. Table 4 reports the results. The estimation outcome is qualitatively similar to that in Table 3. The signs and statistical significance of both long- and short-run coefficients remain unchanged. The pooled error-correction coefficients continue

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to be significantly negative and within the unit circle, indicating that there is a long-run equilibrium relationship among financial development, inflation, and three control variables. Consequently, our findings that inflation has significantly negative effects on financial development in the long run but significantly positive effects in the short run do not seem to suffer severely from common omitted factors. An interesting finding is that although income and government size seem to have a positive and significant influence on financial deepening in both time horizons, trade openness appears to have a significant positive long-run impact but a negative short-run effect on financial development. And, for all three control variables, their long-run impacts are stronger than those in the short run.

### 4.2. The Effects of Inflation Uncertainty

Recently, Dotsey and Sarte (2000) argued that because inflation and inflation uncertainty are highly correlated, the presence of uncertainty tends to attenuate the negative long-run relationship between inflation and real growth. In terms of the inflation–financial development relationship, as claimed by Lucas (1990) and Fuerst (1992), inflation uncertainty resulting from high and variable inflation affects nominal interest rates and so affects decisions to use money or transaction services to make purchases. In Lucas and Stokey (1987) and English (1999), however, it is mean inflation, not inflation uncertainty, that affects transactions decisions. Thus, as another robustness check, this section tests whether there is an independent effect of inflation uncertainty on the financial sector and whether the addition of inflation uncertainty changes the relationship between inflation and financial development.

Following convention, we use conditional variances derived from exponential GARCH (EGARCH) and component GARCH (CGARCH) to proxy for inflation uncertainty. They are denoted as $h_{it}^{eg}$ and $h_{it}^{cg}$, respectively.[17] The PMG estimation results controlling for inflation uncertainty are summarized in Table 5. As indicated, the inclusion of inflation uncertainty does not change our previous exercises. The pooled error-correction coefficient remains significantly negative and falls within the unit circle, supporting a long-run equilibrium relationship among financial development, inflation, inflation uncertainty, and control variables. Moreover, the significantly negative long-run impacts coexist with significantly positive short-run effects, meaning that inflation has direct impacts on financial development in both the short run and the long run. Also, as expected, inflation appears to have a stronger influence in the long run than in the short run.

Note that the long-run coefficient estimate of inflation uncertainty is significant and positive, whereas the short-run estimate is not significant, except for $\Delta h_{it}^{eg}$ on *llly*. This implies that inflation uncertainty affects financial development in a positive fashion in the long run but tends to have an insignificant effect in the short run. The evidence supports the hypothesis that increasing inflation risk encourages precautionary saving that is beneficial for financial sector expansion. This helps

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**TABLE 5.** The PMG results when considering inflation uncertainty

| | Financial development indicator | | |
|---|---|---|---|
| | *lprivo* | *ldby* | *llly* |
| Panel A: Inflation uncertainty measured by EGARCH model | | | |
| Long-run coefficients | | | |
| Inflation | −0.0427*** | −0.0414*** | −0.0262*** |
| | (0.0048) | (0.0044) | (0.0024) |
| Inflation uncertainty, $h_{it}^{eg}$ | 0.1891*** | 0.2915*** | 0.1731*** |
| | (0.0413) | (0.0491) | (0.0294) |
| Error correction | | | |
| Phi | −0.0593*** | −0.0682*** | −0.0930*** |
| | (0.0078) | (0.0070) | (0.0089) |
| Short-run coefficients | | | |
| △ Inflation | 0.0021*** | 0.0026*** | 0.0021*** |
| | (0.0006) | (0.0005) | (0.0006) |
| △ Inflation uncertainty, $\Delta h_{it}^{eg}$ | 0.0001 | 0.0014 | −0.0101** |
| | (0.0076) | (0.0122) | (0.0040) |
| Panel B: Inflation uncertainty measured by CGARCH model | | | |
| Long-run coefficients | | | |
| Inflation | −0.0219*** | −0.0207*** | −0.0171*** |
| | (0.0032) | (0.0023) | (0.0018) |
| Inflation uncertainty, $h_{it}^{cg}$ | 0.0928** | 0.0473 | 0.0547* |
| | (0.0415) | (0.0334) | (0.0294) |
| Error correction | | | |
| Phi | −0.0820*** | −0.0919*** | −0.1061*** |
| | (0.0097) | (0.0091) | (0.0099) |
| Short-run coefficients | | | |
| △ Inflation | 0.0014** | 0.0015*** | 0.0017*** |
| | (0.0007) | (0.0006) | (0.0006) |
| △ Inflation uncertainty, $\Delta h_{it}^{cg}$ | 0.0073 | 0.0086 | −0.0051 |
| | (0.0200) | (0.0136) | (0.0099) |

*Note*: The estimates on control variables are omitted for brevity. The values in parentheses are the standard errors of corresponding coefficient estimates.
***, ** and * Significant at 1%, 5%, and 10% level, respectively.

explain that the Dotsey and Sarte (2000) finding of a positive growth-improving effect of inflation uncertainty possibly works through the financial channel.

### 4.3. The Effects of Economic Development

Literature on the finance–growth nexus postulates that the growth-enhancing effects of finance are nonlinear and even nonmonotonic, depending on the stage of economic development. De Gregorio and Guidotti (1995) and Rioja and Valev (2004), for example, report that positive effects between banking development and economic growth are particularly strong in medium- and high-income countries.

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Deidda and Fattouh (2002) also obtain similar results. By contrast, Calderon and Liu (2003), Wachtel (2003), and Masten et al. (2008) provide strong evidence that the bank–growth link is not as strong among developed countries as it is among less developed ones. It is thus interesting to explore whether such nonlinearity can be attributed to differential responses of finance to inflation. On this point, English (1999) notes that because countries with higher per capita income generally have larger financial sectors, it seems likely that the effect of inflation on financial sectors is larger in high-income countries as well. His cross-country investigation supports this idea. By contrast, Dotsey and Sarte (2000) assert that sustained inflation should not be expected to yield significant growth effects for industrialized countries that appear to have a high degree of financial sophistication. Because our data show that higher initial real GDP per capita seems to have a positive impact on financial depth, this section examines whether the inflation–financial development relationship differs along with economic development.

To test the empirical relevance of real development in the relationship, we divide countries into three equal-sized country subsamples (high-, medium-, and low-income groups), according to real income per capita in the middle of the sample period, redoing the estimation for each country subsample. As claimed in Rioja and Valev (2004), separating countries into three roughly equal-sized groups is fairly mechanical and may leave the positioning of some countries open to skepticism. However, it has the advantage of avoiding subjective judgments on grouping countries. The estimated heterogeneous responses of financial development to inflation are reported in Table 6.

As can be seen, the pooled error-correction coefficients continue to be significantly negative and within the unit circle for all income groups, indicating a long-run cointegrating relationship among financial development, inflation, and control variables. However, the finding of a positive short-run and a negative long-run effect is observed only in the low-income group. It shows that the relationship between inflation and financial development indeed varies with economic development. In particular, the short-run coefficient estimates of inflation tend to be nonsignificant for the medium- and high-income countries but significantly positive for the low-income countries. By contrast, the long-run coefficient estimate of inflation turns out to be significantly negative for all subsamples (except for the case of *ldby* in the high-income group). And the data reveal that the long-run effect of inflation tends to increase with economic development. Finally, the long-run impact is generally larger than the short-run effect.

### 4.4. The Effects of the Inflation Level

Theory suggests that inflation matters because it affects the severity of credit market frictions. And because credit market rationing may not occur in the environment with low rates of inflation and because endogenous rationing of credit with higher rates of inflation worsens information frictions, the relationship between inflation and financial development should be better characterized by threshold nonlinearity

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**TABLE 6.** The PMG results for different income country subsamples

| | Financial development indicator | | |
|---|---|---|---|
| | *lprivo* | *ldby* | *llly* |
| Panel A: High-income countries | | | |
| Long-run coefficients | | | |
| Inflation | −0.0914*** | 0.0000 | −0.0353*** |
| | (0.0168) | (0.0008) | (0.0064) |
| Error correction | | | |
| Phi | −0.0236*** | −0.0749*** | −0.0455*** |
| | (0.0083) | (0.0147) | (0.0111) |
| Short-run coefficients | | | |
| △ Inflation | 0.0010 | 0.0008 | 0.0016 |
| | (0.0012) | (0.0012) | (0.0022) |
| Panel B: Medium-income countries | | | |
| Long-run coefficients | | | |
| Inflation | −0.0213*** | −0.0123*** | −0.0114*** |
| | (0.0041) | (0.0012) | (0.0026) |
| Error correction | | | |
| Phi | −0.0942*** | −0.0967*** | −0.0599*** |
| | (0.0162) | (0.0244) | (0.0221) |
| Short-run coefficients | | | |
| △ Inflation | 0.0011 | 0.0010 | 0.0012** |
| | (0.0010) | (0.0006) | (0.0006) |
| Panel C: Low-income countries | | | |
| Long-run coefficients | | | |
| Inflation | −0.0083*** | −0.0131*** | −0.0085*** |
| | (0.0011) | (0.0025) | (0.0012) |
| Error correction | | | |
| Phi | −0.1347*** | −0.1170*** | −0.1412*** |
| | (0.0210) | (0.0122) | (0.0211) |
| Short-run coefficients | | | |
| △ Inflation | 0.0022*** | 0.0021*** | 0.0021*** |
| | (0.0006) | (0.0005) | (0.0004) |

*Note*: The estimates on control variables are omitted for brevity. The values in parentheses are the standard errors of corresponding coefficient estimates.
***, ** and * Significant at 1%, 5%, and 10% level, respectively.

in the level of inflation. The empirical literature indeed provides evidence of the nonlinearity. Khan et al. (2006) find the potential threshold level of inflation at about 3 to 6%, below which inflation has positive effect on financial development but above which the effect turns negative. By contrast, Boyd et al. (2001) find that the inflation threshold occurs at 15%, above which inflation has limited effects on financial activity. These observations suggest that there might be two thresholds for three regimes in the inflation–financial development link.

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**TABLE 7.** The PMG results for different inflation country subsamples

| | Financial development indicator | | |
|---|---|---|---|
| | *lprivo* | *ldby* | *llly* |
| Panel A: High-inflation countries with inflation >15% (17 countries) | | | |
| Long-run coefficients | | | |
| Inflation | −0.0123*** | −0.0145*** | −0.0105*** |
| | (0.0017) | (0.0016) | (0.0012) |
| Error correction | | | |
| Phi | −0.1087*** | −0.0945*** | −0.1433*** |
| | (0.0272) | (0.0292) | (0.0204) |
| Short-run coefficients | | | |
| △ Inflation | 0.0007 | 0.0003 | 0.0008*** |
| | (0.0006) | (0.0003) | (0.0003) |
| Panel B: Medium-inflation countries with inflation about 6%–15% (41 countries) | | | |
| Long-run coefficients | | | |
| Inflation | −0.0256*** | −0.0245*** | −0.0149*** |
| | (0.0044) | (0.0038) | (0.0039) |
| Error correction | | | |
| Phi | −0.0860*** | −0.0916*** | −0.0498*** |
| | (0.0112) | (0.0103) | (0.0150) |
| Short-run coefficients | | | |
| △ Inflation | 0.0011 | 0.0018*** | 0.0009 |
| | (0.0008) | (0.0006) | (0.0007) |
| Panel C: Low-inflation countries with inflation <6% (29 countries) | | | |
| Long-run coefficients | | | |
| Inflation | −0.0101* | −0.0008 | −0.0149** |
| | (0.0055) | (0.0045) | (0.0064) |
| Error correction | | | |
| Phi | −0.0712*** | −0.0360*** | −0.0524** |
| | (0.0153) | (0.0121) | (0.0247) |
| Short-run coefficients | | | |
| △ Inflation | 0.0023** | 0.0022** | 0.0029** |
| | (0.0010) | (0.0009) | (0.0015) |

*Note*: The estimates on control variables are omitted for brevity. The values in parentheses are the standard errors of corresponding coefficient estimates.
***, ** and * Significant at 1%, 5%, and 10% level, respectively.

Accordingly, in this section, we reinvestigate the issue by dividing countries into high-, medium-, and low-inflation country groups. In particular, countries with annual inflation rates above 15% are classified as high-inflation countries, whereas those with inflation rates below 6% are grouped as low-inflation countries. The rest are medium-inflation countries. The results of differential effects of inflation on financial development are reported in Table 7.

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As indicated, the pooled error-correction coefficient estimate remains significantly negative and lies inside the dynamically stable range for all subsamples, implying long-run equilibrium relationships among financial development, inflation, and control variables. However, the coexistence of positive short-run and negative long-run effects can only be observed in the low-inflation group. Specifically, the short-run coefficient estimate of inflation is significantly positive in low-inflation countries but tends to be nonsignificant in the medium- and high-inflation groups. And the positive short-run effect tends to decrease as inflation rises. This is in line with the debt-deflation/balance-sheet hypothesis. In countries with very high inflation, borrowing and lending are generally for short terms and often in foreign currencies such as the U.S. dollar, so there may not be a significant link between inflation and finance. However, in countries with relatively low inflation rate, the use of long-maturity nominal assets and liabilities is widespread and hence there may be a significant positive short-run association between inflation and financial development.

On the other hand, the long-run estimate is significant and negative in all three subsamples, except for that on *ldby* in the low-inflation group. In addition, we observe a U-shaped long-run link between inflation and financial development in that the magnitude of the negative influence of inflation first increases and then decreases as inflation heightens. This also evidences the existence of inflation thresholds in the relationship between inflation and financial development.

## 5. CONCLUSIONS

In their recent important paper, Dotsey and Sarte (2000) postulate that inflation and growth may be negatively related in the long run although positively related at cyclical frequencies. Moreover, financial development is characterized by long-run financial deepening and by short-run financial instability. Thus this paper investigates whether the heterogeneous responses of growth to inflation work possibly through the financial system. Specifically, we assess whether the impacts of inflation on financial development differ in the short and long run.

Using PMG estimation for a panel of data consisting of 87 countries over 1960–2005, we find evidence of the coexistence of negative long-run effects and positive short-run effects of inflation on financial development. The findings are robust to alternative financial development indicators, conditioning variables, and controlling for inflation uncertainty. However, dividing the sample into different income or inflation groups yields some interesting insights. We find that the inflation–finance link indeed varies with the level of economic development. Although a negative long-run effect coexists with a positive short-run effect of inflation in low-income countries, inflation tends to have negative long-run but insignificant short-run effects on financial development in higher-income countries. Finally, the data suggest that the inflation–finance link is nonlinear. Although the positive short-run impacts of inflation on financial depth seem to decrease with inflation,

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the negative long-run link between inflation and financial development appears to be U-shaped in the level of inflation.

## NOTES

1. Please see Levine (1997, 2005) for an excellent, both theoretical and empirical, survey and references.

2. Banerjee and Newman (1993), Galor and Zeira (1993), and Aghion and Bolton (1997) assert that financial intermediary development reduces income inequality by disproportionately boosting the income of the poor. And Li et al. (2001), Honohan (2004), Clarke et al. (2006), and Beck et al. (2007) provide empirical evidence of the assertion.

3. The existing literature implies that finance can affect macroeconomic cycles. More developed financial markets and institutions more efficiently match savers and investors, allowing the economy to absorb shocks more easily. The financial sector also facilitates diversification (at both the microeconomic and macroeconomic levels), which would reduce risk and volatility. In contrast, financial development may be a proxy for the extent of information asymmetries, which may cause an increased volatility. Please see Denizer et al. (2002), Braun and Larrain (2005), and Raddatz (2006) for discussions and references.

4. Please see Gillman (2005) for theoretical discussions, and Barro (1995), Bullard and Keating (1995), Bruno and Easterly (1998), Khan and Senhadji (2001), Rousseau and Wachtel (2002), and Bose and Murshid (2008), to name a few, for empirical investigations.

5. Cysne et al. (2005), Desai et al. (2005), and Albanesi (2007) show that inflation is positively related to inequality, perhaps because higher inflation may worsen capital market frictions or weaken the political bargaining power of the poor.

6. Please see La Porta et al. (1997, 1998), Levine (2002), Rajan and Zingales (2003, 2004), and Djankov et al. (2007). Also see, Mishkin (2007, 2009) for more detailed discussions.

7. Please see discussions in Azariadis and Smith (1996), Schreft and Smith (1997), Huybens and Smith (1998, 1999), Bose (2002), Hung (2003), and Smith (2003).

8. This is particularly true when nominal instruments with a long maturity are involved (such as long-term bonds or mortgages). Short-term assets are less affected, because nominal interest rates adjust to inflation.

9. The debt-deflation mechanism has been used to analyze the Great Depressions of the 1930s, the sluggishness of Japan's economy in the 1990s, and financial crises in the 1990s: the South East Asian currency crisis in 1997–1998, Mexico 1994–1995, Argentina 1995, Korea 1997–1998, and Russia 1998 [see Martin and Wachtel (1993); King (1994); Bernanke et al. (1999); Mishkin (1999, 2007, 2009); Fackler and Parker (2005); and Mendoza and Smith (2006), to name a few].

10. As noted in Loayza and Ranciere (2006), although averaging clearly induces a loss of information, it is not obvious that averaging over fixed-length intervals effectively eliminates business-cycle fluctuations. Averaging eliminates information that may be used to estimate a more flexible model that allows for some parameter heterogeneity across countries. Further, averaging hides the dynamic relationship between inflation and financial development, particularly the presence of opposite effects at different time frequencies.

11. The PMG estimator has recently been applied to measure the effect of exchange rate uncertainty on investment [Byrne and Davis (2005a, 2005b)], to assess the trade effect of real effective exchange rates [Catao and Solomou (2005)], to estimate the impacts of fiscal deficits on inflation [Catao and Terrones (2005)], to estimate the relationship between financial development and economic growth [Loayza and Ranciere (2006)], to examine the relationship between inequality and growth [Frank (2008)], and to test the Fisher hypothesis [Lin (in press)].

12. The underlying ARDL specification dispenses with unit root pretesting of the variables. Provided that there is a unique vector defining the long-run relationship among variables involved, with the lag orders $p$ and $q$ suitably chosen, MG and PMG estimates of an ARDL regression yield consistent estimates of that vector, no matter whether the variables involved are $I(1)$ or $I(0)$.

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13. It is argued that with increasing globalization in both trade and financial accounts, governments may be forced to adopt less distortionary policies and more disciplined types of macroeconomic management and to implement institutional reforms under the pressures of international competition [Sachs and Warner (1995); Obstfeld (1998); Ades and Di Tella (1999); Stiglitz (2000); Rajan and Zingales (2003); Tytell and Wei (2004); Winters (2004); Gourinchas and Jeanne (2006)]. Therefore, given by conditions expected to converge, the long-run relationship between inflation and financial development may be homogeneous across countries.

14. These measures indicate development of financial intermediaries, instead of financial markets. We use them because they are more available, in terms of both the number of countries and the length of time periods. In particular, the PMG methodology requires large $T$ and $N$ to address dynamic features in data.

15. All variables except inflation are in natural logarithms.

16. Loayza and Ranciere (2006) suggest that when the main interest is in the long-run parameters, the lag order of ARDL can be selected using consistent information criteria on a country-by-country basis; however, when the interest is also in analyzing and comparing the short-run parameters, it is recommended to impose a common lag structure across countries. Thus, this paper sets $p = q = 1$. We have also tried different lag orders selected by the Akaike information criterion (AIC), the Schwarz Bayesian criterion (SBC), and Hannan and Quinn (HQ), finding qualitatively and quantitatively similar results.

17. The detailed specifications for EGARCH and CGARCH are presented in the Appendix.

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# APPENDIX: DERIVING INFLATION VOLATILITY

To estimate the effect of inflation uncertainty on financial development, we need a measure for inflation uncertainty. This is obtained via the GARCH-type models to generate conditional variance as a proxy for inflation uncertainty. Specifically, suppose that inflation follows a pure ARIMA model,

$$\pi_t = \alpha_0 + \sum_{i=1}^{p} \alpha_i \pi_{t-i} + \sum_{j=1}^{q} \varphi_j \varepsilon_{t-j}, \qquad \textbf{(A.1)}$$

where $\varepsilon_t$ is white noise and $\pi_t$ is the rate of inflation. Also, to allow for conditional heteroskedasticity, we assume that $\varepsilon_t | \Omega_{t-1} = h_t^{\frac{1}{2}} \eta_t$ and $h_t \sim NID(0, 1)$. In the study, two alternative specifications of the conditional variance $h_t$ are considered for each country. The first is an exponential GARCH (EGARCH) process proposed by Nelson (1991) that takes account of the asymmetric effects of negative and positive shocks. The specification

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can be written as

$$\ln(h_t) = a_0 + a_1 \ln(h_{t-1}) + b_1 \left| \varepsilon_{t-1} \Big/ \sqrt{h_{t-1}} \right| + c_1 \frac{\varepsilon_{t-1}}{\sqrt{h_{t-1}}}. \quad \textbf{(A.2)}$$

The second model for the conditional variance is an extension of the basic GARCH model. Engle and Lee (1999) represent the GARCH(1,1) model as characterized by reversion to a constant mean $\bar{\mu}$, that is,

$$h_t = \bar{\mu} + a_1(h_{t-1} - \bar{\mu}) + b_1(\varepsilon_{t-1}^2 - \bar{\mu}). \quad \textbf{(A.3)}$$

In contrast, their component GARCH (CGARCH) process allowing reversion to a time-varying mean $m_t$ is modeled as

$$\begin{aligned} h_t - m_t &= \bar{\mu} + a_1 (h_{t-1} - \bar{\mu}) + b_1 \left(\varepsilon_{t-1}^2 - \bar{\mu}\right) \\ m_t &= \mu + \rho (m_{t-1} - \mu) + \varsigma \left(\varepsilon_{t-1}^2 - h_{t-1}\right) \end{aligned} \quad \textbf{(A.4)}$$

We follow conventional applications such as Asteriou and Price (2005) and Byrne and Davis (2005a, 2005b) to proxy inflation uncertainty by the logarithm of the fitted (conditional) volatility values from equations (A.2) and (A.4), respectively. The corresponding inflation uncertainty measures are denoted as $h_{it}^{eg}$ and $h_{it}^{cg}$.

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