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Fiscal Deficits and Inflation

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Abstract

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Macroeconomic theory postulates that fiscal deficits cause inflation. Yet empirical research has had limited success in uncovering this relationship. This paper reexamines the issue in light of broader data and a new modeling approach that incorporates two key features of the theory. Unlike previous studies, we model inflation as nonlinearly related to fiscal deficits through the inflation tax base and estimate this relationship as intrinsically dynamic, using panel techniques that explicitly distinguish between short- and long-run effects of fiscal deficits. Results spanning 107 countries over 1960–2001 show a strong positive association between deficits and inflation among high-inflation and developing country groups, but not among low-inflation advanced economies.

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"A common criticism of this stress on the budget deficit is that the data rarely shows a strong positive association between the size of the budget deficit and the inflation rate."

(Blanchard and Fischer, 1989, p.513)

I. INTRODUCTION

A well-established theory in macroeconomics is that fiscally dominant governments running persistent deficits have sooner or later to finance those deficits with money creation ("seigniorage"), thus producing inflation (Sargent and Wallace, 1981). While this theory does not rule out the importance of other mechanisms through which inflation can be fueled and become persistent, fiscal imbalances have remained central to most models.[2] The "fiscal view" of inflation has been especially prominent in the developing country literature, which has long recognized that less efficient tax collection, political instability, and more limited access to external borrowing tend to lower the relative cost of seigniorage and increase dependence on the inflation tax (Alesina and Drazen, 1991; Cukierman, Edwards, and Tabellini, 1992; Calvo and Végh, 1999).

Yet, as noted in the quote above, empirical work has had little success in uncovering a strong and statistically significant connection between the fiscal deficits and inflation across a broad range of countries and inflation rates. For instance, King and Plosser's (1985) comprehensive analysis of the determinants of seigniorage in the United States and 12 other countries, using both single equation OLS regressions and VARs, indicates no generally significant causality running from fiscal deficits to changes in base money and inflation. In a more restricted sample of high inflation developing countries and using Granger-causality tests and variance decompositions in VARs, Montiel (1989) and Dornbusch, Sturzenegger, and Wolf (1990) find that fiscal deficits tend to accommodate rather than drive inflations—which instead they relate mainly to a combination of exchange rate shocks and inflationary inertia. Employing nonparametric correlation measures for 17 developing countries and dividing them into low- and high-inflation groups, de Haan and Zelhort (1990) find that seigniorage is weakly related to budget deficits except during very high inflation episodes. Click (1998) provides OLS estimates of the determinants of seigniorage in a cross section of 78 (mostly developing) countries and finds that fiscal variables play no significant role. More recently, Fisher, Sahay, and Végh (2002), using fixed effects in a panel of 94 developing and developed economies, conclude that fiscal deficits are main drivers of high inflations (defined in excess of 100 percent a year), and estimate that a 1 percentage point improvement (deterioration) in the ratio of the fiscal balance-to-GDP typically leads to a 4¼ percent decline (rise) in inflation, all else constant. However, they also find that changes in budget balances have no significant inflationary effects in low-inflation countries, or during low-

[2] See Ljungqvist and Sargent (2000), and Fischer, Sahay, and Végh (2002) for recent surveys.

inflation episodes in historically high-inflation countries. Finally, several cross-country studies on the determinants of inflation do not even include fiscal balances in their regressions, implicitly or explicitly assuming that fiscal balances play no role or that their effects are indirectly captured by other variables (Romer, 1993; Lane, 1995; Campillo and Miron, 1997; and Loungani and Swagel, 2001).

This paper takes a new look at this issue. Relative to previous studies, it uses a broad cross-country dataset and proposes a new approach to testing the theory which contains two main novelties. First, a simple intertemporal optimization model is used to show that equilibrium inflation is directly related to the fiscal deficit scaled by narrow money, where the latter stands for the size of the inflation tax base. We show that this distinction between the proposed specification and the standard practice of scaling deficit by GDP is not only theoretically appealing but also empirically relevant, since it introduces a key nonlinearity in the model—namely, it allows a given change in the deficit-to-GDP ratio to have a stronger impact in higher-inflation economies, where inflation tax bases are typically narrower. While previous work has acknowledged the existence of such a nonlinearity and tried to accommodate it through a semi-logarithm specification, the approach we propose is arguably less ad hoc and also empirically superior as shown later on.

Second, and also unlike previous studies, we model the deficit-inflation relationship as intrinsically dynamic, explicitly distinguishing between the short run and long run. Such a distinction is crucial, because fiscal deficits need not lead to higher money creation and inflation in the short run, as governments can temporarily finance their deficits with borrowing. Accordingly, econometric testing of the theory should ideally be capable of uncovering the relevant "equilibrium" or *long-run* parameters amidst a complex (and possibly noncausal) relationship between the two variables in the short run. As discussed in some detail below, 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 estimator" of Pesaran, Shin, and Smith (1999). This method, which explicitly models dynamics, is more germane to the spirit of the theory than the static fixed-effects estimator widely used in the literature, and its country-specific ARDL structure is capable of accommodating cross-country heterogeneity in inflation inertia.

The third main contribution of this paper lies in the use of very broad and up-to-date dataset, spanning 107 countries over 1960–2001 for a total of 3,607 observations. Such a panel is far more comprehensive than those found in previous studies, including Fisher, Sahay, and Végh (2002), which spans 94 countries for a maximum of 2,318 observations. Having a large panel allows us to slice the data into the various groups of interest without issues of sample representativeness or degrees of freedom becoming critical. Also, unlike the existing literature reviewed above, we consider both central and general government balance measures, and test the robustness of the deficit-inflation relationship to the inclusion of several conditioning variables.

The paper is organized as follows. Section II presents the theoretical model. Section III lays out the econometric methodology, while Section IV discusses the data and measurement issues. Estimation results are reported in Section V. Section VI concludes.

II. The Model

A central point of Sargent and Wallace (1981) is that the relationship between fiscal deficit and inflation is dynamic. Under fiscal dominance, deficits determine the *present value* of seigniorage but not necessarily current seigniorage. This is because borrowing allows governments to allocate seigniorage intertemporally, implying that fiscal deficits, seigniorage, and inflation need not be contemporaneously correlated. Moreover, because the short-run dynamics of the deficit-inflation relationship can be very complex (see, e.g., Dornbusch, Sturzenegger, and Wolf, 1990; Calvo and Végh, 1999), its direction and proximate magnitude are not amenable to theoretical predictions.

In contrast, the long-run relationship between the two variables is clearly spelled out by theory. This section shows how a parsimonious and testable specification can be simply derived from a small open economy version of the class of general equilibrium models surveyed by Ljungqvist and Sargent (2000). In this framework, money is assumed to play a role in determining macroeconomic equilibrium through a reduction in transactions costs ("shopping time"), enabling a fiscally dominant government to affect nominal money demand and inflation. The main features of this model economy and its steady-state equilibrium are as follows.

A. Households

The representative household maximizes the following lifetime utility function:

$$\sum_{t=0}^{\infty} \beta^t u(c_t, \iota_t) \tag{1}$$

where β is the subjective discount factor ($0 < \beta < 1$) and where c_t is period-t consumption, and ι_t is period t-leisure. The current-period utility function, u(.,.), is assumed to be strictly increasing and strictly concave in its two arguments.

In each period, the household is endowed with a positive quantity of a good y_t. Out of this endowment, the household pays taxes and can either consume or transfer the after-tax endowment over time through risk-free bond and money holdings. As result, the household is subject to a sequence of budget constraints given by:

$$c_t + \frac{b_{t+1}^p}{R_t^*} + \frac{m_{t+1}}{p_t} = y_t - \tau_t + b_t^p + \frac{m_t}{p_t} \tag{2}$$

where b_t^p is the real value of the household holdings of one-period risk-free bonds that mature at the beginning of period t, these assets are denominated in period t consumption units; m_{t+1} denotes the household's holdings of money balances between t and t+1; τ_t is a lump-sum tax at period t; p_t is the price level; and R_t^* is the international real gross rate of return on one-period bonds. The initial stocks of b_0^p and m_0 are given and $y_t << \infty$.

In each period t, the household has one unit of time which can be allocated to leisure, $\imath_t$, or shopping activities, s_t, so that $\imath_t + s_t = 1$. The amount of time spent on shopping is assumed to be directly related to the level of consumption, c_t, and inversely related to the amount of real balances the household holds between t and t+1 (m_{t+1}/p_t):

$$s_t = S(c_t, \frac{m_{t+1}}{p_t}) \tag{3}$$

where S, S_c, S_{cc}, $S_{m/p,m/p} > 0$ and $S_{m/p}$ and $S_{c,m/p} < 0$. Because transaction costs are negatively related to money holdings, the return on money can be lower than the return in the risk free bond, as in the standard Baumol-Tobin money-demand function.

First order conditions with respect to c_t, $\imath_t$, b_{t+1}, and m_{t+1} yield the following money demand function:

$$\frac{m_{t+1}}{p_t} = M^d(c_t, \frac{1}{R_t^*(1+\pi_t)}) \tag{4}$$

where M^d is increasing on consumption (c_t), and decreasing on the international real interest rate R_t^* as well as on the domestic inflation rate $\pi_t = \frac{p_{t+1}}{p_t} - 1$.

B. Government

In each period t, the government spending g_t is financed with tax collection, the issuance of one-period bonds, or by printing money. So, the respective budget constraint is given by:

$$\frac{b_{t+1}^g}{R_t^*} = \tau_t + b_t^g - g_t + \frac{M_{t+1} - M_t}{p_t} \tag{5}$$

where b_t^g is the real value of the government's net bond holdings denominated in consumption units of period t, and M_t is currency issued by the government at the beginning of the period t. Both b_0^g and M_0 are given. Whenever $b_t^g < 0$, the government is a net borrower in period t.

C. Economy-Wide Budget Constraint and Stationary Equilibrium

With money supply equal to money demand ($m_t = M_t$) and $b_{t+1} = b^p_{t+1} + b^g_{t+1}$ for all t, the economy wide budget constraint is thus:

$$\frac{b_{t+1}}{R^*_t} = y_t - c_t - g_t + b_t \qquad (6)$$

where b_{t+1} is the net holdings of foreign bonds of the economy as a whole and b_0 is given, so that the current account is defined as $b_{t+1} - b_t$.

In the absence of trade restrictions and taxes, both purchasing power parity condition and the uncovered interest rate parity conditions hold, resulting in the equalization of onshore (R_t) and offshore real interest rates (R^*_t). Stationary equilibrium in this small open economy then implies:

$$\begin{aligned} &R = R^* = \beta^{-1} \\ &\frac{M}{p} = M^d(c, \frac{1}{R(1+\pi)}) = \vartheta(\pi) \end{aligned} \qquad (7)$$

Substituting (7) into (5) yields:

$$\frac{\pi}{1+\pi} = \frac{p[g - \tau + b^g \frac{(R-1)}{R}]}{M} \qquad (8)$$

which is the long-run relationship we shall examine in the remainder of the paper. It states that the rate of inflation is proportional to the ratio of gross-of-interest government deficit to the *average* stock of transaction or "narrow" money during the period; or equivalently, that inflation is proportional to the product of the ratio of gross-of-interest fiscal deficit to GDP by the inverse of the ratio of narrow money to GDP. With the demand for transaction money being negatively related to inflation, the size of the inflation tax base (M/GDP) will be lower (higher) as inflation is higher (lower). This implies that fiscal consolidation will be a more powerful instrument of price stabilization the higher the inflation rate.

III. Estimation Methodology

Allowing for generality and making use of the approximation $\pi \approx \pi/(1+\pi)$, we consider the following empirical counterpart of equation (8):

$$\pi = \psi \frac{(G-T)}{M} \qquad (9)$$

where $G-T \approx p(g-\tau+b^{g}\frac{(R-1)}{R})$ is the nominal equivalent of the real budget deficit concept underlying the theoretical model, and ψ is the semi-elasticity parameter to be estimated. The empirical rationale for the nominal deficit approximation and other measurement issues are discussed in detail in Section 4.

To allow for rich dynamics in the way inflation adjusts to changes in the fiscal deficit or to any other variable, we nest equation (9) in an auto-regressive distributed lag (ARDL) structure where dependent and independent variables enter the right-hand side with lags of order p and q, respectively:

$$\pi_{i,t} = \mu_i + \sum_{j=1}^{p} \lambda_{i,j}\pi_{i,t-j} + \sum_{l=0}^{q} \delta_{i,l}^{'}\boldsymbol{x}_{i,t-l} + \varepsilon_{i,t} \qquad (10)$$

where $\pi_{i,t}$ stands for the observed inflation rate in group i at time t; μ_i represents fixed effects; and $\boldsymbol{x}_{i,t}$ is a (kx1) vector of explanatory variables which includes the expression on the right-hand side of (9), i.e., $\boldsymbol{x}_{i,t} = \begin{bmatrix} \frac{(G_{i,t}-T_{i,t})}{M_{i,t}} \\ \boldsymbol{x}_{i,t}^{*} \end{bmatrix}$, and $\boldsymbol{x}^{*}_{i,t}$ is a $(k-1,1)$ vector which includes all other explanatory variables; $\lambda_{i,j}$ are scalars and $\delta_{i,l}$ are (kx1) coefficient vectors. One well-known advantage of working with this ARDL specification, where all right-hand side variables enter the equation with a lag, is to mitigate any contemporaneous causation from the dependent to the independent variable(s) which might bias the estimates.[3] This is an important consideration in the present context due to the presence of money on the right-hand side of (10) and the tight connection between money demand and inflation underlying the theoretical model.

Equation (10) can be re-parameterized and written in terms of a linear combination of variables in levels and first-differences:

$$\Delta\pi_{i,t} = \mu_i + \phi_i\pi_{i,t-1} + \varphi_i^{'}\boldsymbol{x}_{i,t} + \sum_{j=1}^{p-1} \lambda_{i,j}^{*}\Delta\pi_{i,t-j} + \sum_{l=0}^{q-1} \delta_{i,l}^{*'}\Delta\boldsymbol{x}_{i,t-l} + \varepsilon_{i,t}$$

[3] An extensive survey of ARDL models is provided in Banerjee, Dolado, Galbraith, and Hendry (1993). The time-series properties of ARDL models in the estimation of long-run cointegrating relationships are discussed in Pesaran and Shin (1998).

where $\phi_i = -\left(1-\sum_{j=1}^{p}\lambda_{i,j}\right)$, $\varphi_i = \sum_{j=0}^{p}\delta_{i,j}$, $\lambda^*_{i,j} = -\sum_{m=j+1}^{p}\lambda_{i,m}$, $\delta^*_{i,l} = -\sum_{m=l+1}^{q}\delta'_{i,m}$, with $j = 1,2,...,p-1$, and $l = 1,2,...,q-1$. By grouping the variables in levels, this can be re-written as:

$$\Delta\pi_{i,t} = \mu_i + \phi_i[\pi_{i,t-1} - \theta_i' x_{i,t}] + \sum_{j=1}^{p-1}\lambda^*_{i,j}\Delta\pi_{i,t-j} + \sum_{l=0}^{q-1}\delta^{*'}_{i,l}\Delta x^*_{i,t-l} + \varepsilon_{i,t} \qquad (11)$$

where $\theta_i = -\phi_i^{-1}\varphi_i$ defines the long-run equilibrium relationship between the variables involved (i.e. ψ_i, the coefficient on $\frac{(G_{i,t}-T_{i,t})}{M_{i,t}}$, is the first element of this vector) and ϕ_i the speed with which inflation adjusts toward its long-run equilibrium following a given change in $x_{i,t}$.

The econometric literature suggests two approaches to consistent estimation of those parameters in dynamic panels with considerable heterogeneity across the distinct i's and where T is large enough so that (11) can be estimated separately for each country. One is the so-called mean group (MG) estimator. It consists of estimating separate ARDL models for each country and derive θ and ϕ as simples averages of individual country coefficients θ_i and ϕ_i. This produces consistent estimates of the average of the parameters in heterogeneous panels provided that group specific parameters are independently distributed and the regressors are exogenous – Monte Carlo experiments confirming that this result holds generally, provided that N or T are not too small (Hsiao et al., 1999). However, it has also been shown that MG estimates will be inefficient if θ_i and ϕ_i are the same across groups, i.e., if the long-run slope homogeneity restriction holds (Pesaran, Shin, and Smith, 1999). In this case, Pesaran, Shin, and Smith (1999) propose a maximum likelihood-based "pooled mean group" (PMG) estimator which combines pooling and averaging of the individual regression coefficients in (11). This is shown to yield not only consistent but also considerably more efficient estimates that the MGE when the slope homogeneity restriction holds. By allowing the researcher to impose cross-sectionally long-run homogeneity restrictions of the form of $\theta_i = \theta$, $\forall$ i =1,2,..N, the PMG estimator also has the attractive feature of enabling one to test this restriction via standard Hausman-type tests.

Both the MG and PMG estimators have two key advantages over other estimators commonly used in the literature. Unlike the static fixed estimator, they allow for dynamics which is a well-known feature of inflationary processes. Relative to dynamic fixed effects (DFE) estimator, the MGE and the PMGE also have the advantage of allowing the short-run dynamic specification and error variances to differ across countries—a clear benefit since those variances may be quite different reflecting wide international disparities in historical inflation rates. Finally, the underlying ARDL structure dispenses with unit root pre-testing of the variables—a procedure which is marred by the low power of unit root tests and the controversy about their small sample properties in panels (O'Connell, 1998). Provided that there is a unique vector defining the long-run relationship among the variables involved, MG and PMG estimates of an ARDL specification such as in equation (11) yield consistent estimates of that vector—no matter whether the variables involved are I(1) or I(0)—once p and q are suitably chosen.[4]

IV. Data

The theoretical model of Section 2 indicates that the effects of budget deficits on inflation should vary across countries with significantly different inflation rates and levels of financial development, since both have a direct bearing on the size of the inflation tax base. So, sufficient heterogeneity in the country composition of the dataset is an important requirement for rigorous testing of the theory. Moreover, since the theory is mainly concerned with long-run equilibrium relationships and the proposed econometric methodology requires sufficiently long and uninterrupted time series, this is another important data requirement.

The dataset we have put together takes both requirements into account. It comprises *all* countries reported in the IMF's International Financial Statistics (IFS) for which there exist no less than 20 years of *continuous* annual observations for the four variables featuring the theoretical model, i.e., inflation, the budget balance, GDP, and narrow money. Spanning 107 countries over the period 1960–2001, this dataset is the broadest and the most up-to-date we are aware of in the inflation literature.[5] Table 1 describes some of the main features of the data, reporting averages of the relevant ratios by country groups and decade-long sub periods.

As in other studies, inflation is measured by the annual percent change in the consumer price index. The average money stock balance featuring on the right-hand side of (8) is the mean between the current year's end-December stock and the preceding year's end December stock of domestic M1. Since the latter is arguably the closest empirical equivalent for the transactions money concept in the theoretical model and is also a previously used

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[4] If the variables are I(1), the superconsistent property of OLS estimates holds and reverse causality becomes a non-issue (Stock, 1988). If the variables are I(0), the fact that left-hand side variable enter the regression in lagged form helps mitigate endogeneity biases. Moveover, reverse causality in fiscal deficit-inflation relationship seems to be more of an issue only in very high inflation episodes or during hyperinflations (Sargent, 1982; Franco, 1990; Dornbusch, Sturzenegger, and Wolf, 1990).

[5] While the IFS was the main data source, some gaps in the series were filled with data from IMF's country desks and World Economic Outlook databases, and Mitchell (1998a,b,c).

Table 1. Selected Variable Averages, by Country Groups

(Percent)

	Inflation (CPI) [1]	M1/GDP	Central Gov. Balance/ GDP	General Gov. Balance/ GDP	Openness[2]	Oil Prices[3]
All countries						
1961-1970	8.38	18.98	-2.12	...	25.79	-1.43
1971-1980	16.40	17.93	-3.68	-4.14	33.53	46.31
1981-1990	76.35	16.77	-4.50	-4.15	35.09	-1.73
1991-2001	48.85	16.66	-2.65	-2.70	38.99	3.18
Advanced countries						
1961-1970	4.17	27.03	-1.21	...	24.30	-1.43
1971-1980	11.28	24.35	-3.42	-2.02	28.68	46.31
1981-1990	8.36	21.48	-4.36	-3.47	32.00	-1.73
1991-2001	2.85	25.76	-2.81	-2.31	34.15	3.18
Developing countries						
1961-1970	10.18	15.76	-2.58	...	26.39	-1.43
1971-1980	17.86	16.06	-3.76	-4.99	34.98	46.31
1981-1990	94.97	15.48	-4.54	-4.39	35.94	-1.73
1991-2001	61.45	14.17	-2.60	-2.84	40.31	3.18
o/w: Emerging markets						
1961-1970	20.40	16.27	-3.06	...	23.33	-1.43
1971-1980	28.54	16.88	-4.19	-4.47	29.09	46.31
1981-1990	124.05	15.39	-3.92	-4.08	31.42	-1.73
1991-2001	36.82	14.94	-2.03	-2.73	36.31	3.18
Top 25 inflaters						
1961-1970	21.54	12.52	-3.08	...	22.58	-1.43
1971-1980	33.60	13.31	-4.71	-4.23	23.79	46.31
1981-1990	276.37	11.75	-6.43	-5.31	22.98	-1.73
1991-2001	173.84	8.52	-3.27	-3.35	29.12	3.18
Bottom 25 inflaters						
1961-1970	2.89	24.02	-1.41	...	33.96	-1.43
1971-1980	7.98	20.64	-2.84	-2.82	43.52	46.31
1981-1990	4.36	19.52	-3.28	-2.68	48.19	-1.73
1991-2001	2.55	22.27	-2.02	-1.66	49.98	3.18

Sources: International Finance Statistics, IMF's WEO and country desk databases, and Mitchell (1998a,b,c).

1/ Average annual percent change.

2/ ½*(Exports plus Imports)/GDP.

3/ Average annual percent change of the US dollar spot price.

measure of the inflation tax base (e.g., de Haan and Zelhorst, 1990; Rodrick, 1990; Metin, 1998), it seems preferable to other monetary aggregates.[6]

The main fiscal balance measure is the nominal deficit of the central government as reported in the IFS, i.e., including transfers and net interest payments and measured on a cash basis. One issue with this measure concerns its mapping to equation (8), where the term $[g-\tau+b^g\frac{(R-1)}{R}]$ strictly speaking measures changes in the *real* value of government debt. A well-known problem with the nominal deficit measure, as typically reported in the IFS and other statistical sources, is that it can be a misleading indicator of changes in real government debt during high and hyper inflations (regardless of whether nominal deficit is scaled by GDP or current money stock). This has led some practitioners to work with the "operational deficit" concept, defined as the primary balance plus real interest payments on the current debt stock (see, e.g., Tanzi, Blejer, and Teijero, 1993). However, this measure has not only the drawback of requiring certain assumptions and high-frequency data to be calculated with reasonable precision, but also is unavailable for most countries. Another potential criticism of the central government balance measure is that it fails to incorporate local governments, public enterprises, and central bank losses—entities deemed to play a considerable role in inflationary episodes in some countries, especially those where fiscal federalism prevails. In other words, broader deficit measures would be desirable. One problem is that sufficiently long series on public sector aggregates comprising local governments, public enterprises, and central bank losses are unavailable for all 107 countries. However, general government balances for a subset of 87 countries and spanning a shorter-time horizon (usually starting sometime in the mid to late 1970s) are available from the IMF's World Economic Outlook and country desk databases. We use those series to test the robustness of the results, as described below.

V. Results

The second column of Table 2 reports MG and PMG estimates for the 107 country panel with the ARDL lag structure for each country being optimally chosen by the Schwartz

[6] While the change in high-powered money is also a widely used measure of seigniorage, it is less germane to theoretical concept of demand for transactions money in the model. Moreover, high powered money as a measure of the inflation tax base is not unproblematic: it overestimates the inflation tax base when reserve requirements held at the central bank are remunerated (as is the case in some countries in our panel, and underestimates it when the government finds a way of extracting from banks the gains yielded by negative real interest rates paid on sight deposits.

Table 2. Dynamic Panel Estimates of Inflation on Central Government Deficit over M1 [1/]

A. By level of financial development

	All Countries		Advanced Countries		Developing Countries			
					All		Emerging Markets	
	MG	PMG	MG	PMG	MG	PMG	MG	PMG
LR Elasticity (ψ)	1.43	0.02	1.69	-0.29	1.40	0.02	2.26	0.38
	(5.60)	(2.74)	(1.08)	(-4.02)	(2.16)	(3.76)	(1.63)	(31.5)
EC Coefficient (ϕ)	-0.46	-0.49	-0.18	-0.14	-0.53	-0.52	-0.52	-0.40
	(-14.55)	(-15.26)	(-6.76)	(-8.44)	(-16.9)	(-15.4)	(-7.75)	(-5.55)
h-statistic		5.60		1.62		4.51		1.83
		[0.02]		[0.20]		[0.03]		[0.18]
No. of Observations	*3607*	*3607*	*882*	*882*	*2725*	*2725*	*905*	*905*

B. By level of inflation

	All Countries		Top 25		Mid 50		Bottom 25	
	MG	PMG	MG	PMG	MG	PMG	MG	PMG
LR Elasticity (ψ)	1.43	0.02	4.46	0.40	0.18	0.03	0.89	0.00
	(5.60)	(2.74)	(2.35)	(32.78)	(0.44)	(3.52)	(0.86)	(-0.42)
EC Coefficient (ϕ)	-0.46	-0.49	-0.49	-0.43	-0.53	-0.52	-0.53	-0.52
	(-14.55)	(-15.26)	(-7.54)	(-5.95)	(-11.27)	(-11.38)	(-11.58)	(-11.56)
h-statistic		5.60		4.57		0.13		0.75
		[0.02]		[0.03]		[0.72]		[0.39]
No. of Observations	*3607*	*3607*	*924*	*924*	*1765*	*1765*	*918*	*918*

1/ t-ratios in parenthesis and p-values in brackets. The h-statistic refers to the Hausman test on the long-run homogeneity restriction.

Bayesian criterion (SBC).[7] For the vast majority of countries (86 out of 107), specifications with no lagged dependent variables are rejected at conventional levels of statistical significance, indicating that dynamics is important and so that the static fixed-effects method is clearly inadequate to for the task at hand.[8] As the null hypothesis throughout is that of no long-run relation between budget deficit and inflation, t-statistics yielded by both estimators reject the null at 5 or even 1 percent. The estimated error correction coefficient of just under 0.5 indicates that the adjustment of inflation to a given change in the fiscal balance has an average half-life of just over two years. Yet, the dramatic difference between the MG and the PMG estimates of the long-run elasticity parameter ψ points to considerable sample heterogeneity. This is clear from the Hausman h-statistic of 5.6, which rejects the slope homogeneity restriction at 2 percent. Rejection of the homogeneity restriction implies that consistency of the PMG estimate is not warranted and so the MG estimate should be preferred.

To gain some insight into the nature of this heterogeneity, we divide the panel into groups by level of financial development and inflation performance. Clearly, these two broad categories are not unrelated—more financially developed countries do typically display historically lower inflation rates—and not surprisingly about half of the countries we classify as low inflation economies are also classified as developed.[9] However, since the overlapping between the two groups is not perfect (a few countries have both underdeveloped financial systems and stellar inflation performance), and given that other studies have considered low inflation countries a relevant sub-group in its own right, it seems important to consider those two ways of sub-dividing the panel. The respective MG and PMG estimates are reported in Table 2. They indicate that budget deficits are significant drivers of inflation in most groups, with the exceptions of low inflation economies and advanced countries for which one of the

[7] To allow for reasonably rich dynamics without wasting too many degrees of freedom, we impose the condition that $p, q \leq 3$. A shorter lag structure ($p, q \leq 2$) does not change qualitatively the results but lowers the t-ratios in several cases.

[8] The same conclusion holds if other standard model selection criterion—the Akaike Information Criterion—is used.

[9] See the Appendix for the list of countries comprising each group. The developed v. developing country breakdown is based on the IMF's World Economic Outlook classification, whereas the definition of "emerging markets" follows IMF (2001). High inflation countries comprise those in the upper quartile of the inflation distribution (averaged over the 1960–2001 period) and low inflation countries comprise the bottom 25 percent. Using a breakdown by quintiles rather than by quartiles does not change the thrust of the results. Period averages of the inflation rate and of other relevant variables for those various groups are provided in Table 1.

estimates actually yields the "wrong" sign.[10] The estimated effect of changes of budget balance on inflation is very strong for developing countries in general, as the h-test indicates that the MG estimate of 1.40 should be preferred to the PMG estimate of 0.02. For an average M1/GDP ratio of 15.5 percent for these countries (See Table 1), this implies that a 1 percent reduction (increase) in the ratio of budget deficit to GDP lowers (raises) inflation by 8 ¾ percentage points on average, all else constant. For the more homogeneous emerging market group, the impact is less dramatic but still far from negligible. Given that the h-statistic cannot reject the cross-country slope homogeneity restriction and the PMG estimate of 0.38 ought to be preferred, a percentage point change in the ratio of budget balance to GDP is estimated to change inflation by some 2¼ percentage points for historical values of the M1/GDP ratio for this group of countries (see Table 1). Moreover, the t-ratio of 31.5 underlying the estimated coefficient indicates that this elasticity is very precisely calculated implying that the inferences for this group of countries are especially robust.

Breaking down by high vs. lower inflation groups, changes in the budget balance have a very strong effect in high inflation economies which, in our dataset, comprise several countries with average inflation rates above 100 percent during the 1980s and 1990s (see Table 1 and the Appendix). Far less strong but still statistically significant is the effect on "moderate" inflation countries. As the latter category comprises nine advanced countries out of 54 countries comprising the whole group, this raises the question of whether a significant relationship between budget deficits and inflation is also observed for that sub-group. Re-running the separate regressions for this sub-group, we also cannot reject the existence of a positive relationship between deficits and inflation at 5 or 1 percent level of statistical significance.[11] So, even among advanced countries with moderate levels of inflation one can still conclude that budget deficits matter for long-term inflation performance. Only for countries at the very bottom of the inflation distribution is there no evidence that ψ is statistically significant.

This raises the question of why previous studies did not uncover such a significant positive effect of fiscal deficits on inflation across most country groups. Is it because of differences in econometric techniques, model specification, sampling, or a combination of all of the above? To shed light on this question, we re-estimate equation (9) using the static

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[10] Another distinctive feature of the advanced country estimates is the substantially lower coefficient on the error correction term, suggesting that the dynamics of the inflation-deficit relationship takes much longer to unravel.

[11] Specifically, the PMG estimator yields a coefficient of 0.1 with a t-ratio of 3.05 and a h-statistic of 0.29, which clearly does not reject the cross-country slope homogeneity assumption. Full details of these estimates are not reported to conserve on space but are available from the authors upon request.

fixed effects estimator widely used in the literature.[12] As discussed in Section 3, application of the method to cross-country inflation data is bound to be problematic not only due to its neglecting of inflation dynamics, but also because of the assumption of constant error variance across groups. The latter, in particular, is grossly violated in the present panel.[13] Bearing these reservations in mind, the fixed effects estimates reported in panel A of Table 3 basically reinstate the broad inferences obtained with the MG and PMG estimators. The only exception is the developed country group for which the coefficient ψ is now statistically significant and positive, albeit very small. As with the MG and PMG estimators, fiscal deficits do not seem to matter for inflation performance in economies with historically low inflation.

Table 3 also reports fixed effects estimates of the more standard specification in which the budget deficit is scaled by GDP rather than by narrow money. They indicate that the inference that fiscal deficits generally matter can also be obtained with the more standard specification albeit with important quantitative differences relative to the results of Table 2. When inflation is *not* calculated by approximation $\log(1+\pi)$ but simply as $\pi_t = 100(p_t / p_{t-1} - 1)$ where p_t is current CPI, the model does not allow for the inflation-budget deficit elasticity to change across inflation levels; accordingly, panel B estimates are very disparate for the different sub-panels. By taking such non-linearities into account and flattening outlier observations, the use of the log approximation in panel C yields more sensible estimates.[14] Yet, the estimated magnitude of the fiscal effect is generally much lower than those using the (G-T)/M1 specification. For instance, a 1 percentage reduction (increase) in the deficit/GDP ratio is estimated to lower inflation by 1.83 percent—a figure about five times as low as that obtained with the MG estimate and the specification of equation (9). Likewise, the estimated semi-elasticity of 4.05 for the high inflation group, although virtually identical to that reported in Fischer, Sahay and Végh (2002) using a different dataset, is much

[12] We have also estimated (9) using pooled OLS without country-specific fixed effects but standard hausman tests clearly favored the fixed-effect specification relative to pooled OLS at any conventional level of statistical significance.

[13] Standard errors of individual country regressions vary from as low as 1 (Austria) to as high as over 300 as in high inflation economies such as Argentina, Brazil, and Turkey. Such a dispersion of error variances is reflected in the disparate R^2's between the distinct groups as shown in Table 3.

[14] The way the log approximation accomodates non-linearities in the data can be readily seen by taking the derivative of inflation with respect to the deficit in $\ln(1+\pi) = \mu + \psi_y (G-T)/GDP + \varepsilon$. This yields: $\frac{\partial \pi}{\partial[(G-T)/GDP]} = \psi_y (1+\pi)$ which states that, for a given estimate of ψ_y, the effect of a percentage change in the ratio of fiscal deficit to GDP will be higher as π increases.

Table 3. Fixed-Effects Panel Regressions [1/]

A. Inflation on Central Government Deficit/M1

	All Countries	Advanced Countries	Developing Countries: All	Developing Countries: Emerging Markets	By inflation rates: Top 25	By inflation rates: Mid 50	By inflation rates: Bottom 25
Slope	1.35	0.05	1.41	1.51	1.71	0.03	0.01
	(2.93)	(4.65)	(2.55)	(3.13)	(2.46)	(5.05)	(1.14)
R^2	0.11	0.34	0.10	0.14	0.10	0.13	0.07
No. of Obs.:	3623	889	2734	911	925	1773	925

B. Inflation on Central Government Deficit/GDP

	All Countries	Advanced Countries	Developing Countries: All	Developing Countries: Emerging Markets	By inflation rates: Top 25	By inflation rates: Mid 50	By inflation rates: Bottom 25
Slope	19.10	0.37	21.97	12.73	45.68	0.37	0.01
	(2.46)	(5.49)	(2.46)	(3.11)	(2.51)	(7.61)	(0.78)
R^2	0.11	0.19	0.11	0.10	0.13	0.14	0.08
No. of Obs.	3623	889	2734	911	925	1773	925

C. Log(1+Inflation) on Central Government Deficit/GDP

	All Countries	Advanced Countries	Developing Countries: All	Developing Countries: Emerging Markets	By inflation rates: Top 25	By inflation rates: Mid 50	By inflation rates: Bottom 25
Slope	1.83	0.33	2.06	2.95	4.05	0.33	0.03
	(6.36)	(5.59)	(6.27)	(6.15)	(6.45)	(7.77)	(0.91)
R^2	0.35	0.34	0.35	0.33	0.31	0.14	0.14
No. of Obs.:	3623	889	2734	911	925	1773	925

1/ Heteroscedasticity corrected t-ratios in parenthesis.

lower than that previously using the dynamic panel methods and the (G-T)/M1 specification. Only when dynamic panel estimators are used instead of static fixed effects, can the combination of log approximation and the deficit/GDP specification yield significantly higher elasticities, as shown in Table 4. Yet, magnitude of the overall effect is still smaller than that yielded by the combination of the (G-T)/M1 specification and the dynamic panel estimators. This is especially the case regarding emerging market countries, for which the specification of Table 4 yields a semi-elasticity of inflation to the deficit-GDP ratio of 0.38, as opposed to 2¼ (=0.38/16.75) in Table 2.

It follows from these results that difficulties in uncovering a statistically significant and strong relationship between budget deficits and inflation stem from two main factors. One is the use of data samples with a disproportionately high weight on advanced countries and/or low inflation economies, since those would bias the evidence towards rejecting the theory. Second, from inadequate statistical methods such as the fixed effects estimator which, combined with specifications that do not account for differences in the size of the inflation tax base, impart a downward bias on the relevant cross-country estimates.

Before fully embracing the dynamic panel estimates of Table 2, however, it is important to test their robustness to the inclusion of other explanatory variables considered in the literature. One such a variable is oil price—a well-known source of inflationary pressures in the world economy (IMF, 2001; Loungani and Swagel, 2001; Hamilton and Herrera, 2002). As shown in Table 5, the statistical significance of oil price inflation as an explanatory variable varies somewhat across MG and PMG estimates and across country groups, but on the whole holds across much of the panel.[15] Consistent with the findings of the above mentioned studies, the impact of oil prices on domestic long-term inflation is generally modest in developing countries but is quite strong among advanced countries, with a 1 percentage point increase in oil price inflation estimated to raise domestic inflation by some 0.2 percentage points. This role for oil price inflation in the regressions does not detract, however, from the strength of previous estimates on the fiscal deficit variable. With or without including oil price inflation, fiscal deficits continue to display a powerful effect on inflation in developing countries, emerging markets, and high-inflation economies and a much smaller effect amongst moderate inflation countries. Also as before, the estimated coefficient for advanced countries continues to yield an opposite sign as that predicted by theory, but the coefficient for low inflation countries now turns out to be statistically significant, albeit remaining very small.

The significance of the deficit-inflation relationship in developing and high inflation countries also appears to be robust to the potential omission of two other explanatory

[15] The average fit of individual country regressions also improves significantly with the inclusion of oil prices. The full panel country averages of the adjusted R^2 for the model with oil is 0.38, as opposed to 0.27 without oil. The full panel average unadjusted R^2 with oil is 0.45.

Table 4. Dynamic Panel Estimates of Log (1+Inflation) on Central Government Deficit over GDP 1/

A. By level of financial development

	All Countries		Advanced Countries		Developing Countries			
					All		Emerging Markets	
	MG	PMG	MG	PMG	MG	PMG	MG	PMG
LR Semi-Elasticity	4.43	0.14	-1.15	-1.28	5.87	0.21	1.92	0.38
	(2.15)	(3.36)	(-1.54)	(-4.50)	(2.25)	(4.60)	(0.26)	(2.90)
EC Coefficient (ϕ)	-0.45	-0.43	-0.18	-0.14	-0.52	0.50	-0.39	-0.36
	(-14.67)	(-14.52)	(-5.50)	(-9.76)	(-15.10)	(-14.42)	(-6.56)	(-7.08)
h-statistic		4.32		0.04		4.72		0.04
		[0.04]		[0.84]		[0.03]		[0.84]
No. of Observations	*3612*	*3612*	*884*	*884*	*2728*	*2725*	*911*	*911*

B. By level of inflation

	All Countries		Top 25		Mid 50		Bottom 25	
	MG	PMG	MG	PMG	MG	PMG	MG	PMG
LR Semi Elasticity	1.43	0.02	14.83	10.76	1.81	0.30	-0.11	-0.91
	(5.60)	(2.74)	(2.08)	(12.67)	(1.06)	(4.93)	(1.66)	(-1.92)
EC Coefficient (ϕ)	-0.46	-0.49	-0.37	-0.22	-0.53	-0.53	-0.35	-0.36
	(-14.55)	(-15.26)	(-7.08)	(-5.02)	(-11.07)	(-11.40)	(-7.84)	(-7.00)
h-statistic		5.60		0.33		0.77		2.93
		[0.02]		[0.57]		[0.38]		[0.09]
No. of Observations	*3604*	*3604*	*922*	*922*	*1765*	*1765*	*917*	*917*

1/ t-ratios in parenthesis and p-values in brackets. The h-statistic refers to the Hausman test on the long-run homogeneity restriction.

Table 5. Dynamic Panel Estimates of Inflation on Central Government Deficit over M1 and Oil Price Inflation 1/

A. By level of financial development

	All Countries			Advanced Countries			All Developing Countries			Emerging Markets		
	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]
LR Elasticities:												
(G-T)/M1	1.14	0.03	4.05	-0.20	0.07	10.1	1.51	0.03	4.46	2.47	0.38	1.75
	(2.06)	(4.97)	[0.04]	(-2.29)	(7.20)	[0.00]	(2.15)	(4.78)	[0.03]	(1.57)	(30.67)	[0.19]
Oil Price Inflation	-0.02	0.07	0.08	0.21	0.17	0.4	-0.08	0.05	0.1	-0.27	0.08	0.18
	(-0.05)	(20.18)	[0.78]	(3.25)	(15.99)	[0.53]	(-0.19)	(13.33)	[0.75]	(-0.32)	(6.46)	[0.67]
EC Coefficient (ϕ)	-0.50	-0.47		-0.29	-0.24		-0.56	-0.53		-0.39	-0.41	
	(-15.39)	(-15.88)		(-7.22)	(-5.65)		(-14.8)	(-14.25)		(-5.94)	(-6.68)	
No. of Observations	*3607*	*3607*		*848*	*848*		*2725*	*2725*		*905*	*905*	

B. By level of inflation

	All Countries			Top 25			Mid-50			Bottom 25		
	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]
LR Elasticities:												
(G-T)/M1	1.14	0.03	4.05	4.83	0.40	4.61	0.11	0.06	2.6	0.07	0.03	4.56
	(2.06)	(4.97)	[0.04]	(2.34)	(33.30)	[0.03]	(1.05)	(7.79)	[0.11]	(2.02)	(0.38)	[0.03]
Oil Price Inflation	-0.02	0.07	0.08	-0.35	0.05	0.09	0.11	0.08	1.28	0.062	0.078	1.58
	(-0.05)	(20.18)	[0.78]	(-0.27)	(2.09)	[0.76]	(3.26)	(14.07)	[0.26]	(4.59)	(14.54)	[0.21]
EC Coefficient (ϕ)	-0.50	-0.47		-0.45	-0.40		-0.56	-0.56		-0.41	-0.44	
	(-15.39)	(-15.88)		(-7.75)	(-5.96)		(-10.9)	(-12.09)		(-8.99)	(-8.84)	
No. of Observations	*3607*	*3607*		*924*	*924*		*1765*	*1765*		*918*	*918*	

1/ t-ratios in parenthesis and p-values in brackets. The h-statistic refers to the Hausman test on the long-run homogeneity restriction.

variables. One is openness to foreign trade. As argued in Romer (1993) and Lane (1995), the benefits of an expansionary monetary policy tend to be smaller in an economy with a larger share of trade in GDP because: (i) the weight of the home goods sector will be smaller implying that the impact of monetary expansion on domestic employment will be reduced; and (ii) the currency depreciation resulting from the monetary expansion will raise domestic inflation by more than in a closed economy. Hence, the more open the economy the less time-inconsistent the monetary policy, implying a negative relationship between openness and inflation, once other factors are held constant.[16]

Yet, Table 6 results lend limited support to the view. When openness (measured as the ratio of exports plus imports to GDP) enters the regression, MG estimates for the full panel yield the expected negative coefficient but this is statistically insignificant at 5 or 10 percent levels, whereas the PMG estimator yields a coefficient with the opposite sign as that predicated by theory. Only for the developed country group does the openness variable yield a statistically significant coefficient with the predicted sign, but respective effect is relatively small – a one percentage point increase in openness leading to 0.09 percentage point drop in the inflation rate. On the one hand, this suggests that the Romer (1993) and Lane (1995) results are sample specific, as argued in Terra (1998) and Bleaney (1999). On the other hand, Table 6 results indicate that fiscal deficit-inflation relationship is robust to the inclusion of openness among the regressors.

The other explanatory variable we consider is the exchange rate regime. By tying domestic inflation to that of a low inflation country and being more conducive to fiscal and monetary discipline, fixed exchange rate regimes can arguably contribute to lower inflation (Gosh, Gulde, Ostry, and Wolf, 1997). Yet, it has also been argued that in allowing policy markers to lower temporarily inflation without a concomitant fiscal adjustment, fixed exchange rates can actually detract from fiscal discipline and give rise to a peso problem (Tornell and Velasco, 2000; Fatás and Rose, 2001). If so, no positive relationship between flexible exchange rates and inflation should be expected, at least in the medium- to long-run. We consider these hypotheses by including in the regressions the Reinhart-Rogoff (2002) *de facto* index of exchange rate flexibility, which is defined as ranging from 0 (complete inflexibility) to 15 (extreme floating). Our estimates show no evidence of a statistically

[16] The other widely studied hypothesis derived from the time inconsistency theory of monetary policy is that inflation should be lower in countries with more independent central banks or with central banks which are credibly committed to a low inflation mandate (Cuckierman and others, 1992; de Haan and Kooi, 2000). For evidence that central bank behavior helps explain historical swings in inflation rates in the United States, see Goodfriend (1997) and Ireland (1999). Lack of long time series on central bank independence measures for most countries in our panel unfortunately prevents us from evaluating this hypothesis on a broad cross-country basis.

Table 6. Dynamic Panel Estimates of Inflation on Central Government Deficit over M1, Oil Price Inflation, and Openness

A. By Level of Financial Development

	All Countries			Advanced Countries			Developing Countries			Emerging Markets		
	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]
LR Elasticities:												
(G-T)/M1	0.96	0.02	6.21	-0.08	0.02	2.04	1.24	0.03	6.54	1.99	0.07	3.58
	(2.55)	(4.90)	[0.01]	(-1.12)	(1.30)	[0.15]	(2.62)	(6.64)	[0.01]	(1.96)	(4.77)	[0.06]
Oil Price Inflation	-2.00	0.07	0.97	0.13	0.17	10.74	-2.60	0.05	0.97	0.40	0.05	4.30
	(-0.95)	(19.74)	[0.32]	(6.61)	(13.01)	[0.00]	(-0.97)	(13.49)	[0.32]	(2.36)	(8.22)	[0.04]
Openness	-10.90	0.01	1.15	0.06	-0.09	0.22	-13.89	0.02	1.15	3.29	0.03	0.15
	(-1.07)	(1.25)	[0.28]	(0.19)	(-2.12)	[0.64]	(-1.07)	(2.31)	[0.28]	(0.39)	(1.51)	[0.70]
EC Coefficient (ϕ)	-0.56	-0.51		-0.29	-0.20		-0.64	-0.58		-0.60	-0.54	
	(-18.39)	(-16.67)		(-7.17)	(-8.12)		(-19.24)	(-16.41)		(-9.45)	(-8.63)	
No. of Observations	3581	3581		882	882		2699	2699		899	899	

B. By Level of Inflation

	All Countries			Top 25			Mid 50			Bottom 25		
	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]
LR Elasticities:												
(G-T)/M1	0.96	0.02	6.21	3.68	0.36	5.84	0.09	0.03	1.65	-0.80	0.01	4.85
	(2.55)	(4.90)	[0.01]	(2.68)	(21.19)	[0.02]	(2.01)	(4.53)	[0.20]	(-2.15)	(0.11)	[0.03]
Oil Price Inflation	-2.00	0.07	0.97	-8.11	0.03	0.96	0.07	0.06	0.89	0.49	0.08	1.88
	(-0.95)	(19.74)	[0.32]	(-0.98)	(1.53)	[0.33]	(4..98)	(12.53)	[0.35]	(2.51)	(14.39)	[0.17]
Openness	-10.90	0.01	1.15	-43.07	0.23	1.16	-0.02	0.09	0.94	-0.06	-0.01	0.15
	(-1.07)	(1.25)	[0.28]	(-1.07)	(1.64)	[0.28]	(-0.20)	(3.88)	[0.33]	(-0.48)	(-0.94)	[0.70]
EC Coefficient (ϕ)	-0.56	-0.51		-0.56	-0.49		-0.6	-0.56		-0.47	-0.41	
	(-18.39)	(-16.67)		(-10.31)	(-7.72		(-12.72)	(-12.15)		(-9.17)	(-9.04)	
No. of Observations	3581	3581		915	915		1754	1754		912	912	

1/ t-ratios in parenthesis and p-values in brackets. The h-statistic refers to the Hausman test on the long-run homogeneity restriction.

significant relationship between exchange rate flexibility and inflation.[17] As shown at the top of Table 7, the coefficient on exchange rate flexibility for the whole sample yields the expected positive sign but the associated t-ratio is well below usual levels of statistical significance for the whole panel as well as for most country, with notable the exception of emerging markets. But even for the latter, the respective MG and PMG estimates fall short of being statistically significant at 5 percent. Moreover, the estimated magnitude of the effect varies widely across different country groups and, in some of them, the estimates yield the opposite sign as expected by proponents of fixed exchange rates as instruments of price stabilization.

Finally, we test the robustness of our results to the use of a broader measure of the fiscal balance. As noted in Section 3, general government data is only available for a subset of countries (86 out of the 107 countries) and for many of them the respective series do not start until the mid to late 1970s, so the panel size is now drastically reduced. Notwithstanding these changes in panel coverage, Table 8 shows that budget deficits remain a statistically significant driver of inflation for the whole sample as well as for developing and high inflation country groups. Moreover, estimates are on the whole similar to those previously obtained using the central government measure. The only noteworthy differences lie in the advanced country group (for which both MG and PMG estimates of ψ are statistically insignificant) and the low inflation group (for which the h-statistic cannot reject the cross-country homogeneity restriction and the corresponding PMG estimate of 0.10 is statistically significant at 1 percent). Thus, if anything the use of general government balances yields even more favorable results to the theory than the central government measure.

VI. Conclusions

Economic theory postulates a clear causal connection between fiscal deficits and inflation. However, the strength of this relationship is not easy to measure. One reason stressed by Sargent and Wallace (1981) is that deficits need not cause inflation in the short run, since under fiscal dominance deficits determine the *present value* of seigniorage but not necessarily current seigniorage. This implies that proper empirical assessment of the theory requires sufficiently long time series and econometric techniques that can capture the dynamic dimension of this relationship. Second, theory also suggests that the inflation-deficit semielasticity may vary across countries with disparate inflation levels and distinct inflation tax bases. So, a suitable model should be also capable of accommodating such non-linearities. Third, the strength of the effect is likely to depend on the country's level of financial development and central bank commitment to low inflation—deeper financial markets and more credible central banks in advanced economies tend to facilitate continuous

[17] As the index is not available for all countries in the data set, the panel size drops to 78 countries.

Table 7. Dynamic Panel Estimates of Inflation on Central Government Deficit over M1, Oil Price Inflation, and Exchange Rate Regime 1/

1. By level of financial development

	All Countries			Advanced Countries			Developing Countries			Emerging Markets		
	MG	PMG	*h-stat.*	MG	PMG	*h-stat.*	MG	PMG	*h-stat.*	MG	PMG	*h-stat.*
LR Elasticities:												
(G-T)/M1	0.68	0.04	4.3	-0.11	0.33	205.6	1.08	0.06	4.62	1.66	0.38	1.62
	(2.21)	(5.80)	[0.04]	(-1.87)	(6.29)	[0.00]	(2.26)	(5.99)	[0.03]	(1.65)	(27.82)	[0.20]
Oil Price Inflation	-0.02	0.07	0.52	-0.14	-0.35	0.65	0.07	0.05	0.02	-0.45	0.08	0.44
	(-0.12)	(18.04)	[0.47]	(-0.51)	(-4.99)	[0.42]	(0.55)	(10.31)	[0.89]	(-0.56)	(6.47)	[0.51]
Exchange Rate Regime[1]	1.80	2.93		-0.06	-0.22		3.02	-0.65		3.61	3.30	
	(0.96)	(1.18)		(-1.36)	(-1.20)		(1.05)	(-0.47)		(1.75)	(1.84)	
EC Coefficient (ϕ)	-0.50	-0.45		-0.35	0.00		-0.59	-0.52		-0.61	-0.44	
	(-12.57)	(-12.65)		(-7.69)	(-0.01)		(-12.55)	(-11.64)		(-7.16)	(-5.70)	
No. of Observations	*2820*	*2820*		*852*	*852*		*1968*	*1968*		*872*	*872*	

2. By level of inflation

	All Countries			Top 25			Mid-50			Bottom 25		
	MG	PMG	*h-stat.*	MG	PMG	*h-stat.*	MG	PMG	*h-stat.*	MG	PMG	*h-stat.*
LR Elasticities:												
(G-T)/M1	0.68	0.04	4.3	1.97	0.39	1.98	0.00	0.08	2.76	-0.08	0.00	2.79
	(2.21)	(5.80)	[0.04]	(1.75)	(31.33)	[0.16]	(0.37)	(8.30)	[0.10]	(1.63)	(0.06)	[0.09]
Oil Price Inflation	-0.02	0.07	0.52	-0.55	0.02	0.42	0.12	0.11	0.19	0.072	0.08	0.44
	(-0.12)	(18.04)	[0.47]	(-0.63)	(0.69)	[0.52]	(4.31)	(15.88)	[0.67]	(5.70)	(12.26)	[0.51]
Exchange Rate Regime[1]	1.80	2.93		-7.04	-6.89		0.11	0.15		3.61	3.30	
	(0.96)	(1.18)		(-1.24)	(-1.24)		(0.91)	(1.15)		(1.75)	(1.84)	
EC Coefficient (ϕ)	-0.50	-0.45		-0.49	-0.37		-0.5	-0.47		-0.61	-0.44	
	(-12.57)	(-12.65)		(-5.17)	(-4.42)		(-8.56)	(-9.29)		(-7.16)	(-5.70)	
No. of Observations	*2820*	*2820*		*812*	*812*		*1311*	*1311*		*697*	*697*	

1/ t-ratios in parenthesis and p-values in brackets. The h-statistic refers to the Hausman statistic to test the cross-country long-run homogeneity restriction. Exchange rate regime measured by the Reinhart-Rogoff (2002) index and treated as a fixed regressor.

Table 8. Dynamic Panel Estimates of Inflation on General Government Deficit over M1 and Oil Price Inflation.

A. By level of financial development

	All Countries			Advanced Countries			Developing Countries			Emerging Markets		
	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]
LR Elasticities:												
(G-T)/M1	1.41	0.02	4.03	-0.42	-0.04	1.23	2.08	0.04	4.90	2.47	0.37	1.58
	(2.03)	(2.96)	[0.04]	(-1.31)	(-1.49)	[0.27]	(2.23)	(5.66)	[0.03]	(1.55)	(18.36)	[0.21]
Oil Price Inflation	0.143	0.08	1.32	0.55	0.20	1.57	-0.89	0.05	1.82	0.57	0.07	0.28
	(1.20)	(16.35)	[0.25]	(1.94)	(9.32)	[0.21]	(-1.27)	(11.42)	[0.18]	(0.47)	(3.86)	[0.60]
EC Coefficient (ϕ)	-0.50	-0.55		-0.22	-0.16		-0.66	-0.52		-0.53	-0.37	
	(-13.40)	(-12.92)		(-8.32)	(-8.36)		(-11.04)	(-11.59)		(-7.34)	(-5.23)	
No. of Observations	*2368*	*2368*		*670*	*670*		*1698*	*1698*		*684*	*684*	

B. By level of inflation

	All Countries			Top 25			Mid-50			Bottom 25		
	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]	MG	PMG	*h-stat.*[1]
LR Elasticities:												
(G-T)/M1	1.41	0.02	4.03	7.03	0.40	4.89	0.12	0.04	0.74	0.04	0.10	1.02
	(2.03)	(2.96)	[0.04]	(2.34)	(25.63)	[0.03]	(1.37)	(3.91)	[0.39]	(0.65)	(2.64)	[0.31]
Oil Price Inflation	0.143	0.08	1.32	-3.12	-0.03	1.65	0.07	0.09	0.06	-0.11	-0.35	1.97
	(1.20)	(16.35)	[0.25]	(-1.29)	(-0.74)	[0.20]	(0.97)	(10.36)	[0.80]	(-0.60)	(-4.89)	[0.16]
EC Coefficient (ϕ)	-0.50	-0.55		-0.68	-0.38		-0.54	-0.47		-0.45	-0.45	
	(-13.40)	(-12.92)		(-4.17)	(-3.87)		(-8.81)	(-9.28)		(-0.41)	(-7.21)	
No. of Observations	*2368*	*2368*		*504*	*504*		*1140*	*1140*		*724*	*724*	

1/ t-ratios in parenthesis and p-values in brackets. The h-statistic refers to the Hausman test on the long-run homogeneity restriction.

rolling over of sizable debt stocks, obviating the need for seigniorage. So, it is important that empirical assessments of the theory take these differences into account.

This paper has sought to address each of these issues. By using dynamic panel data techniques, we modeled the deficit-inflation relationship as intrinsically dynamic. Also unlike previous studies, this relationship was modeled as nonlinear in the inflation tax base, leading to a distinct specification in which the fiscal deficit is scaled by narrow money rather than by GDP—a distinction shown to be empirically important in a panel comprising low- and high-inflation countries. Finally, working with a panel spanning 107 countries and 42 years of data has enabled us to test the theory over a reasonably lengthy horizon and across distinct institutional settings and inflation thresholds, without sample representativeness issues or degree of freedom limitations becoming critical.

On the other hand, the results are much more favorable to fiscal-based theories of inflation than previous research had found. Fiscal deficits have been shown to matter not only during high and hyperinflations but also under moderate inflation ranges, even though effects are substantially weaker in the latter case. Disaggregating by country groups, the deficit-inflation relationship comes out as surprisingly strong over a broad range of developing countries, also in contrast with the earlier literature. Moreover, the theory is also supported by the finding that none of the alternative explanatory variables considered undermine the strength of the effect or prove to be statistically significant across much of the panel, with the sole exception of oil prices. In particular, trade openness was found to matter for the developed country group but not for all countries; and little evidence was found that fixed exchange rate regimes help lower inflation on a systematic basis.

On the other hand, and similarly to previous studies, we did not detect any positive and strong connection between deficits and inflation in advanced economies and low-inflation country groups. This begs the question of why the theory seems to be violated in those two cases. Regarding the low-inflation country group, since half of its constituents consist mostly of very small, open economies with longstanding hard pegs or those that have given up their national currencies altogether (see the Appendix), the assumption of fiscal dominance underlying the theory is either severely weakened or nonexistent. The other half (13 out of 26) comprises advanced countries, taking us straight back to the question as to why the theory does not seem to hold for this group. Since we find evidence of a statistically significant relationship between budget deficits and inflation among advanced countries in the middle inflation ranges, violation of the theory appears to be more narrowly confined to the subgroup of low-inflation advanced countries. As noted above, greater monetary policy credibility, institutional constraints to fiscal dominance, and deeper financial markets seem to be at least part of the answer.

Last but not least, this paper has also shown that the *statistical significance* of the fiscal deficit-inflation relationship in most countries is relatively robust to alternative specifications and the use of standard panel data techniques. At the same time, we have also shown that the measured *strength* of the effect is not: the dynamic panel estimators and the econometric specification proposed in this paper yield considerably higher elasticities and seem to fit the data better than the standard specification using static fixed effects. This suggest that previous failures in uncovering a strong relationship between budget deficits and inflation stem, at least in part, from using a model specification and/or econometric techniques which do not accommodate key features of the theory. In sum, this paper's findings indicate that the continuing prominence of fiscal-based theories of inflation does not seem unjustified, after all.

Appendix: List of Countries and Country Groups

A. By level of financial development			B. By Inflation level		
Advanced	*Emerging Markets*	*Other Developing*	*Top 25*	*Mid-50*	*Bottom 25*
Australia	Argentina	Bahamas	Argentina	Barbados	Australia
Austria	Brazil	Bahrain	Bolivia	Belize	Austria
Belgium	Chile	Barbados	Brazil	Bhutan	Bahamas
Canada	China	Belize	Chile	Botswana	Bahrain
Cyprus	Colombia	Bhutan	Colombia	Burundi	Belgium
Denmark	Czech Republic	Bolivia	Congo, Dem. Rep.	Cameroon	Burkina Faso
Finland	Egypt	Botswana	Ecuador	Chad	Canada
France	Hungary	Burkina Faso	Ghana	Costa Rica	China
Germany	India	Burundi	Guyana	Czech Republic	Cyprus
Greece	Indonesia	Cameroon	Iceland	Denmark	St. Vincent & Gren.
Iceland	Israel	Chad	Indonesia	Dominican Republic	France
Ireland	Jordan	Congo, Dem. Rep. of	Israel	Egypt	Germany
Italy	Korea	Costa Rica	Malawi	El Salvador	Japan
Japan	Malaysia	Dominican Republic	Mexico	Ethiopia	Malaysia
Netherlands	Mexico	Ecuador	Nicaragua	Fiji	Malta
New Zealand	Pakistan	El Salvador	Nigeria	Finland	Morocco
Norway	Peru	St. Vincent & Gren.	Peru	Gabon	Netherlands
Portugal	Philippines	Ethiopia	Poland	Gambia	Norway
Spain	Poland	Fiji	Romania	Greece	Oman
Sweden	Singapore	Gabon	Sierra Leone	Guatemala	Panama
Switzerland	South Africa	Gambia, The	Tanzania	Haiti	Singapore
United Kingdom	Thailand	Ghana	Turkey	Honduras	St. Kitts and Nevis
United States	Turkey	Guatemala	Uganda	Hungary	Sweden
	Uruguay	Guyana	Uruguay	India	Switzerland
	Venezuela	Haiti	Venezuela	Iran	Thailand
	Zimbabwe	Honduras	Zambia	Ireland	United States
		Iran	Zimbabwe	Italy	
		Kenya		Jordan	
		Lesotho		Kenya	
		Madagascar		Korea	
		Malawi		Lesotho	
		Maldives		Madagascar	
		Malta		Maldives	
		Mauritius		Mauritius	
		Morocco		Myanmar	
		Myanmar		Nepal	
		Nepal		New Zealand	
		Nicaragua		Pakistan	
		Nigeria		Papua New Guinea	
		Oman		Paraguay	
		Panama		Philippines	
		Papua New Guinea		Portugal	
		Paraguay		Rwanda	
		Romania		Solomon Islands	
		Rwanda		South Africa	
		Sierra Leone		Spain	
		Solomon Islands		Sri Lanka	
		Sri Lanka		St. Lucia	
		St. Kitts and Nevis		Swaziland	
		St. Lucia		Syrian Arab Republic	
		Swaziland		Tonga	
		Syrian Arab Republic		Trinidad and Tobago	
		Tanzania		Tunisia	
		Tonga		United Kingdom	
		Trinidad and Tobago			

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