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Journal of Monetary Economics 52 (2005) 529–554

Journal of MONETARY ECONOMICS

www.elsevier.com/locate/jme

Fiscal deficits and inflation ☆

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Received 26 June 2001; received in revised form 14 February 2003; accepted 2 June 2004
Available online 1 April 2005

Abstract

Macroeconomic theory postulates that persistent fiscal deficits are inflationary. 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 non-linearly 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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JEL classification: E31; E60

Keywords: Fiscal policy; Inflation; Macroeconomic stabilization; Developing countries

☆We thank Tam Bayoumi, Arnold Harberger, Ayhan Kose, Jonathan Ostry, Kenneth Rogoff, an anonymous referee as well as the editor, Robert King, for extensive comments on earlier drafts. The paper also benefited from comments by seminar participants at the IMF, UCLA, and the LACEA 2003 meetings. Emily Conover and Ben Sutton provided excellent research assistance. The views expressed in this paper are the authors' alone and do not necessarily reflect those of the International Monetary Fund.
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doi:10.1016/j.jmoneco.2004.06.003

"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)

1. Introduction

A well-established theory in macroeconomics is that 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.[1] 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 et al., 1992; Calvo and Végh, 1999).

Yet, as highlighted in the above quote from Blanchard and Fischer's (1989) textbook, 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 et al. (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 Zelhorst (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, Fischer et al. (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

[1]See Ljungqvist and Sargent (2000, chapter 17), and Fischer et al. (2002) for concise surveys. On a recent class of models associated with the fiscal theory of the price level, which views the government intertemporal budget constraint as an equilibrium condition wherein the price level adjusts so as to accommodate changes in fiscal performance, see Sims (1997), Cochrane (1998), and Woodford (2001). A critical discussion of the underpinnings of these models is provided in Canzoneri et al. (2001) and Buiter (2002).

ratio of the fiscal balance-to-GDP typically leads to a $4\frac{1}{4}$ 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-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, 1997; Campillo and Miron, 1997; Loungani and Swagel, 2001).

This paper takes a new look at this issue. Relative to previous studies, it uses a broad cross-country data set 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 non-linearity 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 and the 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 non-causal) relationship between the two variables in the short run. As discussed 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 et al. (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 data set, spanning 107 countries over 1960–2001 for a total of 3607 observations. Such a panel is far more comprehensive than those found in previous studies, including Fischer et al. (2002), which spans 94 countries for a maximum of 2318 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 core finding of the paper is that fiscal deficits are inflationary in most countries. We find that this relationship is especially strong for developing economies and for countries with average inflation rates at the top quartile of cross-country distribution over the period 1960–2001. Furthermore, we also find that fiscal deficits have a significant bearing on long-run inflation among countries within a "moderate" inflation range (defined as those with upper single digit and lower double-digit annual rates)—another novel result relative to earlier studies. In contrast and corroborating the findings of previous work, fiscal deficits appear to have no significant positive effect on long-run inflation among developed countries with a long history of low single-digit inflation. These various results are robust to alternative partitions of the sample and conditioning to other variables.

The remainder of the paper is organized as follows. Section 2 presents the theoretical model. Section 3 lays out the econometric methodology, while Section 4 discusses the data and measurement issues. Estimation results are reported in Section 5. Section 6 concludes. Specifics of sample breakdowns and country list are provided in Appendix A.

2. The model

A central point of Sargent and Wallace (1981) is that the relationship between fiscal deficit and inflation is dynamic. Under an independently set fiscal policy ("fiscal dominance"), deficits determine the *present value* of the necessary money creation ("seigniorage") to finance them, but do not necessarily determine *current* seigniorage and hence current inflation. This is because borrowing allows governments to allocate seigniorage intertemporally, implying that fiscal deficits 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 et al., 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.

2.1. Households

The representative household maximizes the following lifetime utility function:

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

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 a result, the household is subject to a sequence of budget constraints given by

$$c_t + \frac{b^{\mathrm{p}}_{t+1}}{R^*_t} + \frac{m_{t+1}}{p_t} = y_t - \tau_t + b^{\mathrm{p}}_t + \frac{m_t}{p_t}, \tag{2}$$

where b^{p}_t 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^{p}_0 and m_0 are given and $y_t \ll \infty$.

In each period t, the household has one unit of time which can be allocated to leisure, ι_t, or shopping activities, s_t, so that $\iota_t + s_t = 1$. 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 (m_{t+1}/p_t) the household holds between t and $t+1$:

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

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, ι_t, b_{t+1}, and m_{t+1} yield the following money demand function:

$$\frac{m_{t+1}}{p_t} = M^{\mathrm{d}}\left(c_t, \frac{1}{R^*_t(1+\pi_t)}\right) \tag{4}$$

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 $\pi_t = (p_{t+1}/p_t) - 1$.

2.2. Government

In each period t, the government spending g_t is financed with tax collection, the issuance of one-period bonds, the reduction of its international asset holdings (if any), or by printing money. So, the respective budget constraint is given by

$$\frac{b^{\mathrm{g}}_{t+1}}{R^*_t} = \tau_t + b^{\mathrm{g}}_t - g_t + \frac{M_{t+1} - M_t}{p_t}, \tag{5}$$

where b_t^g is the real value of the government's net asset 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.

2.3. Economy-wide budget constraint and stationary equilibrium

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

$$\frac{b_{t+1}}{R_t^*} = y_t - c_t - g_t + b_t, \tag{6}$$

where b_{t+1} is the net holdings of foreign assets 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 domestic (R_t) and international real interest rates (R_t^*). Stationary equilibrium in this small open economy then implies:

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

Substituting (7) into (5) yields in stationary equilibrium:

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

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 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.

3. Estimation methodology

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

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

where $G - T \approx p(g - \tau + b^g((R-1)/R))$ 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 5.

To allow rich dynamics in the way inflation adjusts to changes in the fiscal deficit or to any other variable, we nest Eq. (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:

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

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 $(k \times 1)$ 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 $\boldsymbol{\delta}_{i,l}$ are $(k \times 1)$ 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.[2] 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.

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

$$\begin{aligned}\Delta\pi_{i,t} &= \mu_i + \phi_i\pi_{i,t-1} + \boldsymbol{\varphi}'_i\boldsymbol{x}_{i,t} \\ &\quad + \sum_{j=1}^{p-1} \lambda^*_{i,j}\,\Delta\pi_{i,t-j} + \sum_{l=0}^{q-1} \boldsymbol{\delta}^{*'}_{i,l}\,\Delta\boldsymbol{x}_{i,t-l} + \varepsilon_{i,t},\end{aligned}$$

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

$$\begin{aligned}\Delta\pi_{i,t} &= \mu_i + \phi_i[\pi_{i,t-1} - \boldsymbol{\theta}'_i\boldsymbol{x}_{i,t}] \\ &\quad + \sum_{j=1}^{p-1} \lambda^*_{i,j}\,\Delta\pi_{i,t-j} + \sum_{l=0}^{q-1} \boldsymbol{\delta}^{*'}_{i,l}\,\Delta\boldsymbol{x}^*_{i,t-l} + \varepsilon_{i,t},\end{aligned} \tag{11}$$

where $\boldsymbol{\theta}_i = -\phi_i^{-1}\boldsymbol{\varphi}_i$ defines the long-run equilibrium relationship between the variables involved (i.e. ψ_i, the coefficient on $(G_{i,t} - T_{i,t})/M_{i,t}$, is the first element of

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

this vector) and ϕ_i the speed with which inflation adjusts toward its long-run equilibrium following a given change in $\boldsymbol{x}_{i,t}$.

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 $\boldsymbol{\theta}$ and ϕ as simple averages of individual country coefficients $\boldsymbol{\theta}_i$ 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. However, it has also been shown that MG estimates will be inefficient if $\boldsymbol{\theta}_i$ is the same across groups, i.e., if the long-run slope homogeneity restriction holds (Pesaran et al., 1999). In this case, Pesaran et al. (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 $\boldsymbol{\theta}_i = \boldsymbol{\theta}$, $\forall i = 1, 2, \ldots, N$, 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 Eq. (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.[3]

4. 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

[3]If the variables are $I(1)$, the superconsistent property of OLS estimates holds and reverse causality becomes a non-issue (Stock, 1987). 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 et al., 1990).

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 data set 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 data set we have put together satisfies both requirements. 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 data set is the broadest and the most up-to-date we are aware of in the inflation literature.[4] 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 with inflation, fiscal deficits tend to be persistent. This can be more formally gauged by individual country regressions of the current annual deficit (scaled by either money or GDP) on its lag. With deficit scaled by money, the estimated autoregressive coefficients lie betweeen 0.5 and 1.0 and are statistically significant at 1 percent for 73 out of the 107 countries, averaging 0.80 for the developed country group, 0.62 for emerging markets, and 0.53 for all developing countries. Allowing for a higher lag order in the specification and/or scaling the fiscal balance by GDP yield similar estimates. Full results are available from the authors upon request.

On other measurement specifics, throughout this paper we measure inflation by the annual percent change in the consumer price index. The money stock variable featuring in Table 1 as well as in all regressions 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 the closest empirical equivalent for the transactions money concept in the theoretical model and is also a previously used measure of the inflation tax base (e.g., de Haan and Zelhorst, 1990; Rodrik, 1991; Metin, 1998), it is therefore preferable to other monetary aggregates.[5]

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 Eq. (8), where the term $[g - \tau + b^{g}((R-1)/R)]$ 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

[4]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–c).

[5]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 1
Selected variable averages by country groups[a] (percent)

	Inflation (CPI)[b]	M1/GDP	Central gov. balance/GDP	General gov. balance/GDP	Openness[c]	Oil prices[d]
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–c).

[a]For the group classification criterion, see the text. A list of their constituent countries is provided in the appendix.

[b]Average annual percent change.

[c]1/2 *(exports plus imports)/GDP.

[d]Average annual percent change of the US dollar spot price.

hyperinflations (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 et al., 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 85 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.

5. Results

Table 2 reports MG and PMG estimates for the 107 country panel. Throughout we choose the optimal ARDL lag structure for each country by the Schwartz Bayesian criterion (SBC) except in the few instances where the SBC yields full panel statistics ambiguously close to the 5 percent threshold, in which cases the Akaike information criterion (AIC) is used as an alternative.[6] 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.[7] 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 one year. 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 insight into the nature of this heterogeneity, we divide the panel into groups by level of financial development and inflation performance. The developed

[6] To allow for reasonably rich dynamics without loosing too many degrees of freedom, we generally impose the condition that $p, q \leqslant 3$. All PMG estimates in multivariate regressions have been computed by the Newton–Raphson (NR) algorithm, which takes into account the first and second derivatives of the likelihood function. The bivariate regressions of Tables 2 and 4 alternate the NR with the back-substitution algorithm (which uses only the first derivative) since the latter yields lower standard errors in some cases. Reassuringly, the two algorithms yield nearly identical estimates in the vast majority of regressions. All PMG numerical computations use the mean group estimates as starting values.

[7] Individual country regressions are not reported for space reasons but are available from the authors upon request.

Table 2
Dynamic panel estimates of inflation on central government deficit over M1[a]

A. By level of financial development

	All countries		Advanced countries		Developing countries			
	MG	PMG	MG	PMG	All		Emerging markets	
					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)	−0.49	−0.46	−0.18	−0.14	−0.53	−0.52	−0.52	−0.40
	(−15.26)	(−14.55)	(−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

	Top 25				Mid-50		Bottom 25	
	All countries		Excl. hyperinflaters		MG	PMG	MG	PMG
	MG	PMG	MG	PMG				
LR elasticity (ψ)	4.46	0.40	0.49	0.40	0.18	0.03	0.89	0.00
	(2.35)	(32.80)	(3.46)	(32.79)	(0.44)	(3.52)	(0.86)	(−0.42)
EC coefficient (ϕ)	−0.49	−0.43	−0.47	−0.43	−0.53	−0.52	−0.40	−0.39
	(−7.54)	(−5.95)	(−6.60)	(−5.75)	(−11.27)	(−11.38)	(−7.21)	(−7.70)
h-statistic		4.57		0.42		0.13		0.75
		[0.03]		[0.52]		[0.72]		[0.39]
No. of observations	924	924	713	713	1765	1765	918	918

[a] *t*-ratios in parenthesis and *p*-values in brackets. The *h*-statistic refers to the Hausman test on the long-run homogeneity restriction.

vs. 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 (characterized as having average annual inflation higher than $15\frac{1}{2}$ percent over 1960–2001), while low-inflation countries comprise the bottom 25 percent, i.e., those with average annual inflation rates lower than $15\frac{3}{4}$ percent over 1960–2001.[8] Clearly, these broad categories are not unrelated—for instance, 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 (see Appendix A.). However, since the

[8] A breakdown by quintiles rather than by quartiles does not change the thrust of the results. See Appendix A for the list of countries comprising each group.

overlapping between such groups is far from perfect, and given that other studies have considered high- and low-inflation countries as relevant sub-groups in their own right, it seems important to consider such a breakdown of 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. The exceptions are low-inflation economies and advanced countries for which of the PMG estimate actually yields the "wrong" sign, whereas lower ϕ's suggest greater adjustment inertia. The 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 1.4 should be preferred to the PMG estimate of 0.02. For an average M1/GDP ratio of $15\frac{1}{2}$ 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 $9\frac{1}{4}$ percentage points on average, all else constant. For the more homogeneous emerging market group, the impact is less dramatic but still sizeable. 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 $2\frac{1}{4}$ 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 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 data set, comprise several countries with average inflation rates above 100 percent during the 1980s and 1990s. To evaluate the extent to which this result is being influenced by a handful of countries that experienced very high and hyperinflations (Argentina, Brazil, Bolivia, Congo, Nicaragua, and Peru), we also report results once they are excluded from the group of 25 top inflaters. The respective long-run elasticity yielded by the PMG estimator remains sizeable and precisely estimated, thus indicating that fiscal deficits also have powerful inflationary effects in the sub-group that excludes extreme inflations. 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 9-country advanced country 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.[9] 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 the effect is present.

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

[9]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.

Table 3
Fixed effects panel regressions[a]

	All countries	Advanced countries	Developing countries		By inflation rates		
			All	Emerging markets	Top 25	Mid-50	Bottom 25
A. Inflation on central government deficit/M1							
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							
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							
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

[a]Heteroscedasticity corrected *t*-ratios in parenthesis.

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 Eq. (9) using the static fixed effects estimator widely used in the literature.[10] 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.[11] 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.

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 the

[10]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.

[11]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 shown in Table 3.

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.[12] 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 Eq. (9).

Likewise, the estimated semi-elasticity of 4.05 for the high-inflation group, although virtually identical to that reported in Fischer et al. (2002) using a different data set, is much 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, the magnitudes are still smaller on the whole than those yielded by the combination of the (*G*–*T*)/M1 specification and the dynamic panel estimators. For the emerging market group in particular, the specification of Table 4 yields a semi-elasticity of inflation to the deficit-GDP ratio of 0.38, as opposed to $2\frac{1}{4}$ ($= 0.38/16$) in Table 2.

In short, previous difficulties in uncovering a statistically significant and strong relationship between budget deficits and inflation seem to stem from two main factors. One is the use of data samples with a disproportionately high weight on advanced countries or economies with historically low inflation. One reason is that countries which have well-established institutions that curb fiscal profligacy, central banks that are credibly committed to low inflation, and deep financial markets, arguably have great latitude in managing their intertemporal budget constraints: as noted in Canzoneri et al. (2001), the necessary turn around in the primary balance to ensure fiscal solvency can then take much longer—possibly decades or even over a century. Coupled with evidence that the primary surplus to GDP ratio does respond more elastically to the government debt ratio in advanced countries than in developing ones (Bohn, 1998; IMF, 2003)—hence suggesting much lower fiscal dominance in the former group—it is not surprising that the inflationary effects of fiscal deficits in advanced countries can go undetected. Among developing countries, a main obstacle to uncovering the fiscal deficit-inflation relationship is inadequate modeling choice: as shown above, the fixed effects estimator combined with specifications that do not account for differences in the size of the inflation tax base imparts a downward bias on the relevant cross-country estimates.

[12] The way the log approximation accommodates non-linearities in the data can be readily seen by taking the derivative of inflation with respect to the deficit in $\ln(1+\pi) = \mu + \psi_y(G-T)/\mathrm{GDP} + \varepsilon$. This yields $\partial\pi/\partial[(G-T)/\mathrm{GDP}] = \psi_y(1+\pi)$ 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 4
Dynamic panel estimates of log(1 + inflation) on central government deficit over GDP[a]

	All countries		Advanced countries		Developing countries			
	MG	PMG	MG	PMG	All		Emerging markets	
					MG	PMG	MG	PMG
A. By level of financial development								
LR elasticity	4.43	0.14	−0.84	−1.15	5.87	0.21	1.92	0.38
	(2.15)	(3.36)	(−1.67)	(−4.52)	(2.25)	(4.60)	(0.26)	(2.90)
EC coefficient (ϕ)	−0.45	−0.43	−0.20	−0.15	−0.52	0.50	−0.39	−0.36
	(−14.67)	(−14.52)	(−6.29)	(−9.31)	(−15.10)	(−14.42)	(−6.56)	(−7.08)
h-statistic		4.32		0.50		4.72		0.04
		[0.04]		[0.48]		[0.03]		[0.84]
No. of observations	3612	3612	884	884	2728	2728	911	911

B. By level of inflation

	Top 25				Mid-50		Bottom 25	
	All countries		Excl. hyperinflation		MG	PMG	MG	PMG
	MG	PMG	MG	PMG				
LR elasticity	14.83	10.76	5.39	3.48	1.81	0.30	−0.91	−0.11
	(2.08)	(12.67)	(2.60)	(7.69)	(1.06)	(4.93)	(1.92)	(−1.66)
EC coefficient (ϕ)	−0.37	−0.22	−0.34	−0.29	−0.53	−0.53	−0.36	−0.35
	(−7.08)	(−5.02)	(−6.15)	(−5.76)	(−11.07)	(−11.40)	(−7.00)	(−7.89)
h-statistic		0.33		0.90		0.77		2.93
		[0.57]		[0.34]		[0.38]		[0.09]
No. of observations	922	922	711	711	1773	1773	917	917

[a] *t*-ratios in parenthesis and *p*-values in brackets. The *h*-statistic refers to the Hausman test on the long-run homogeneity restriction.

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. Because the addition of other explanatory variables create the possibility that the restriction of a homogeneous long-run coefficient across countries may hold for only for a subset of these variables, it is important to allow for this possibility. Accordingly, all the long-run coefficients reported henceforth correspond to regressions where such a cross-country restriction is only applied to those variables for which *h*-statistic cannot reject the homogeneity restriction at 5 percent or less. A first control variable considered is oil price—a well-known source of inflationary pressures in the world economy (Loungani and Swagel, 2001; Barsky and Kilian, 2002; Hamilton and Herrera, 2004). As shown in Table 5, world oil price inflation is

Table 5
Dynamic panel estimates of inflation on central government deficit over M1 and oil price inflation[a]

	All countries	By level of financial development			By level of inflation		
		Advanced	All developing	Emerging markets	Top 25	Mid-50	Bottom 25
Long run elasticities:							
$(G–T)$/M1	1.13	−0.52	1.45	0.39	4.52	0.07	0.00
	(2.18)	(−1.89)	(2.20)	(31.34)	(2.33)	(9.65)	(−0.71)
Oil price inflation	0.08	0.16	0.06	0.09	0.05	0.12	0.08
	(22.15)	(15.70)	(14.83)	(7.22)	(2.37)	(20.53)	(14.29)
EC coefficient (ϕ)	−0.50	−0.25	−0.56	−0.39	−0.47	−0.55	−0.41
	(−15.78)	(−5.65)	(−15.58)	(−5.91)	(−7.53)	(−12.54)	(−9.35)
h-statistics:							
$(G–T)$/M1	4.15	5.92	4.50	1.75	4.61	2.98	2.75
	[0.04]	[0.01]	[0.03]	[0.19]	[0.03]	[0.08]	[0.10]
Oil price inflation	0.10	0.20	0.12	0.18	0.10	0.06	0.83
	[0.75]	[0.66]	[0.73]	[0.67]	[0.76]	[0.81]	[0.36]
No. of observations	3607	882	2725	905	924	1765	918

[a] *t*-ratios in parenthesis and *p*-values in brackets. The *h*-statistic refers to the Hausman test on the restriction that the respective long-run coefficient is the same across groups in the panel. The reported long-run elasticities refer to specifications which apply the long-run homogeneity restriction only to those variables for which the restriction is not rejected by the *h*-statistic.

a significant explanatory variable of inflation across the panel.[13] Consistent with the findings of the above-mentioned studies, the impact of oil prices on domestic long-term inflation is stronger among advanced countries than among developing countries, with a 1 percentage point increase in oil price inflation estimated to raise advanced country inflation by near 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 smaller effect amongst moderate inflation countries.

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

[13]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. Also, since similar inferences obtain and given space constraints, we do not report the results for the high-inflation group excluding countries with very high-inflation and hyperinflation episodes as in Tables 2 and 4. These estimates are available from the authors upon request.

explanatory variables. One is openness to foreign trade. As argued in Romer (1993) and Lane (1997), 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, all else constant.[14]

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, estimates for the full panel yield a coefficient with the opposite sign as that postulated by theory and which is statistically insignificant at conventional levels. Only for the developed country group does the openness variable yield a statistically significant coefficient at 5 percent and with the predicted sign, indicating that a 1 percentage point increase in openness leading to 0.10 percentage point drop in the inflation rate. On the one hand, this suggests that the Romer (1993) and Lane (1997) results are group 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 (Ghosh et al., 1997). Yet, it has also been argued that in allowing policy makers 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 (2004) 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 significant relationship between exchange rate flexibility and inflation.[15] As shown in Table 7, the coefficient on exchange rate flexibility for the whole panel yields the expected positive sign but the associated *t*-ratio is well below usual levels of statistical significance. The coefficient is also statistically insignificant (and sometimes with

[14]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 (Cukierman et al., 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.

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

Table 6
Dynamic panel estimates of inflation on central government deficit over M1, oil price inflation, and trade openness[a]

		By level of financial development			By level of inflation		
	All countries	Advanced	All developing	Emerging markets	Top 25	Mid-50	Bottom 25
Long run elasticities:							
$(G–T)$/M1	0.75	0.00	1.03	0.32	3.42	0.05	−0.11
	(2.13)	(0.21)	(2.42)	(18.27)	(2.83)	(6.13)	(−2.37)
Oil price inflation	0.07	0.16	0.05	0.32	0.01	0.08	0.07
	(20.97)	(6.50)	(14.05)	(2.19)	(0.86)	(16.15)	(14.41)
Trade openness	0.01	−0.10	0.02	−0.11	0.17	0.04	0.002
	(0.62)	(−3.24)	(1.65)	(−1.79)	(1.23)	(1.65)	(0.18)
EC coefficient (ϕ)	−0.52	−0.27	−0.60	−0.47	−0.51	−0.56	−0.42
	(−16.82)	(−10.86)	(−17.27)	(−6.69)	(−8.51)	(−12.01)	(−8.89)
h-statistics:							
$(G–T)$/M1	6.30	2.04	6.57	2.73	15.59	0.95	4.93
	[0.01]	[0.15]	[0.01]	[0.10]	[0.00]	[0.33]	[0.03]
Oil price inflation	0.98	10.8	0.97	3.85	1.26	0.74	1.94
	[0.32]	[0.00]	[0.32]	[0.05]	[0.26]	[0.39]	[0.16]
Trade openness	1.15	0.22	1.15	0.16	0.27	0.29	0.15
	[0.28]	[0.64]	[0.28]	[0.69]	[0.60]	[0.59]	[0.70]
No. of observations	3581	882	2699	899	915	1754	912

[a] *t*-ratios in parenthesis and *p*-values in brackets. The h-statistic refers to the Hausman test on the restriction that the respective long-run coefficient is the same across groups in the panel. The reported long-run elasticities refer to specifications which apply the long-run homogeneity restriction only to those variables for which the restriction is not rejected by the *h*-statistic.

point estimates of the opposite sign) for all country groups with the exception of emerging markets; but even then statistical significance falls short of the 5 percent level.

Finally, we test the robustness of our results to the use of a broader measure of the fiscal balance. As noted in Section 4, general government data is only available for a subset of countries (85 out of the 107 countries) and for many of them the respective series do not start until the mid to late 1970s, so that the time series dimension of much of the panel is also reduced. While such shorter time series tends to sacrifice precision in the estimation of long run coefficients, it does provide an opportunity to gauge the existence of mid-period structural breaks in the relevant relationships. Table 8 shows that budget deficits not only remain a statistically significant driver of inflation among most country

Table 7
Dynamic panel estimates of inflation on central government deficit over M1, oil price inflation, and exchange rate regime[a]

		By level of financial development			By level of inflation		
	All countries	Advanced	All developing	Emerging markets	Top 25	Mid-50	Bottom 25
Long run elasticities:							
$(G–T)$/M1	0.69	−0.34	1.07	0.40	0.39	0.09	0.00
	(1.99)	(−2.16)	(2.32)	(30.10)	(31.33)	(8.56)	(0.06)
Oil price inflation	0.08	0.18	0.05	0.09	0.02	0.13	0.08
	(20.70)	(16.01)	(11.49)	(6.45)	(0.69)	(18.56)	(12.26)
Exchange rate regime[b]	1.38	−0.07	3.06	3.3	7.04	0.11	0.03
	(0.84)	(−1.49)	(1.13)	(1.84)	(−1.24)	(0.90)	(0.66)
EC coefficient (ϕ)	−0.48	−0.24	−0.57	−0.44	−0.37	−0.45	−0.39
	(−13.00)	(−5.02)	(−12.35)	(−5.54)	(−4.42)	(−9.61)	(−8.38)
h-statistics:							
$(G–T)$/M1	4.62	8.04	4.69	1.58	1.98	0.37	2.79
	[0.03]	[0.00]	[0.03]	[0.21]	[0.16]	[0.54]	[0.09]
Oil price inflation	0.54	1.51	0.01	0.45	0.42	1.30	0.19
	[0.46]	[0.20]	[0.92]	[0.50]	[0.52]	[0.25]	[0.66]
No. of observations	2820	852	1968	837	812	1311	697

[a] *t*-ratios in parenthesis and *p*-values in brackets. The *h*-statistic refers to the Hausman test on the restriction that the respective long-run coefficient is the same across groups in the panel. The reported long-run elasticities refer to specifications which apply the long-run homogeneity restriction only to those variables for which the restriction is not rejected by the *h*-statistic.

[b] Exchange rate regime measured by the Reinhart and Rogoff (2004) index and treated as a fixed regressor. Hence no *h*-statistics on long-run restrictions is reported for that coefficient.

groups, but also that the respective estimates are on the whole similar to those previously obtained using the central government measure and longer time series for most countries. This clearly underscores the robustness of our previous results to the use of a broader fiscal measure and to the possibility of mid-period structural breaks.

6. Conclusion

Economic theory postulates a 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 persistent deficits cause inflation in the long run but not necessarily in the short run. This implies that proper

Table 8
Dynamic panel estimates of inflation on general government deficit over M1, oil price inflation[a]

		By level of financial development			By level of inflation		
	All countries	Advanced	All developing	Emerging markets	Top 25	Mid-50	Bottom 25
Long run elasticities:							
$(G–T)$/M1	1.16	0.00	1.83	0.38	7.03	0.07	0.00
	(2.16)	(0.11)	(2.07)	(19.75)	(2.34)	(6.91)	(−0.35)
Oil price inflation	0.11	0.23	0.06	0.08	−3.12	0.17	0.10
	(17.18)	(9.63)	(12.59)	(4.88)	(−1.29)	(16.51)	(12.22)
EC coeficient (ϕ)	−0.43	−0.16	−0.59	−0.41	−0.68	−0.49	−0.37
	(−12.48)	(−7.85)	(−13.83)	(−5.72)	(−4.17)	(−9.00)	(−8.01)
h-statistics:							
$(G–T)$/M1	4.23	2.11	4.34	1.68	32.7	0.51	1.59
	[0.04]	[0.15]	[0.04]	[0.20]	[0.00]	[0.74]	[0.21]
Oil price inflation	0.02	0.58	0.66	0.52	6.04	0.16	1.24
	[0.88]	[0.45]	[0.42]	[0.47]	[0.01]	[0.69]	[0.27]
No. of observations	2300	602	1698	684	504	1114	682

[a] *t*-ratios in parenthesis and *p*-values in brackets. The *h*-statistic refers to the Hausman test on the restriction that the respective long-run coefficient is the same across groups in the panel. The reported long-run elasticities refer to specifications which apply the long-run homogeneity restriction only to those variables for which the restriction is not rejected by the *h*-statistic.

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 elasticity 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 credible policy commitment to low inflation—deeper financial markets and more credible central banks in advanced economies tend to facilitate continuous rolling over of sizable debt stocks and obviate the need for inflating the debt away. So, it is important that empirical assessments of the theory take these differences into account.

This paper has addressed 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 non-linear 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 lengthy horizon and across distinct institutional settings and inflation thresholds.

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. Disaggregating by country groups, the deficit-inflation relationship is especially strong over a broad range of developing countries, also in contrast with the earlier literature. Moreover, none of the alternative explanatory variables previously considered in previous studies undermine the strength of this effect or prove to be statistically significant across the panel, with the sole exception of world oil price changes. In particular, trade openness was found to matter for the developed country group but not for all countries; and there is no evidence that fixed exchange rate regimes help lower inflation on a systematic basis.

As previous researchers have found, however, we do not detect any positive and strong connection between deficits and inflation among low inflation advanced economies and other low-inflation country groups. This begs the question of why the theory seems to be violated in those cases. Since half of the constituents of the low-inflation group consist mostly of very small, open economies with longstanding hard pegs or those that have given up their national currencies altogether (see Appendix A), the assumption of fiscal dominance underlying the theory is either severely weakened or non-existent. 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 range, violation of the theory appears to be more narrowly confined to the subgroup of low-inflation advanced countries. As discussed earlier on, the answer seems to lie at least in part on greater monetary policy autonomy and credibility, as well as other institutional constraints that make public borrowing more closely related to tax and spending smoothing, rather than a systematic source of financing given the exhaustion of other, non-inflationary sources.

Finally, 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 to the use of standard panel data techniques when combined with sufficiently long data series and a broad country panel. But we have also shown that the measured *strength* of the effect is not: the dynamic panel estimators and the econometric specification employed in this paper yield considerably higher elasticities and fit the data much better than the standard specification using static fixed effects. Hence, previous failures in uncovering a strong relationship between budget deficits and inflation partly stem not only from sample selection biases but also from using a model specification and econometric techniques which do not accommodate key features of the theory. Once these limitations are overcome, support for the view that persistent fiscal deficits are inflationary is strong.

Appendix A. 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		
Tunisia		
Uganda		
Zambia		

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