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Price Stickiness, Inflation, and Output Dynamics: A Cross-Country Analysis

by

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Abstract

The sticky-price model of aggregate fluctuations implies that countries with high trend inflation rates should exhibit less-persistent output fluctuations than countries with low trend inflation. I conduct a cross-country analysis of output persistence and inflation that takes into account the within-country time variation in trend inflation. My results do not support the implication. The results suggest that further research is needed before models based on nominal price stickiness can offer a complete microfoundation for persistent effects of aggregate demand shocks.

JEL classification: E31, E32
Bank classification: Business fluctuations and cycles

Résumé

D'après l'explication que le modèle à prix rigides fournit du cycle économique, les fluctuations de la production devraient être moins persistantes dans les économies où le taux de l'inflation tendancielle est élevé que dans celles où il est bas. L'auteur analyse la persistance de la production et l'inflation dans divers pays tout en tenant compte des variations temporelles que l'inflation tendancielle affiche dans chacun d'eux. Ses résultats ne confirment pas la validité du modèle examiné. De nouvelles recherches semblent nécessaires pour que des modèles reposant sur l'hypothèse de rigidité des prix nominaux parviennent à expliquer entièrement au niveau microéconomique la persistance des effets des chocs de demande globale.

Classification JEL : E31, E32
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1 Introduction

Expansions and contractions in economic activity, as summarized by the deviation of national output from its long-run growth path, typically last several years.[1] A major challenge for macroeconomists is to explain the persistent nature of output fluctuations. A model of business cycles that emphasizes aggregate demand-driven fluctuations (particularly via monetary shocks) gives price stickiness a central role in propagating the effect of shocks.[2] Price stickiness can arise, for example, in the presence of small fixed costs (or menu cost) associated with adjusting nominal prices.[3] Therefore, in response to aggregate demand shocks, firms may keep their prices unchanged and adjust output. When such pricing decisions are staggered across firms, the aggregate price level is slow to adjust and demand shocks can cause output to persistently deviate from its long-run growth path. The duration for which a firm keeps its price unchanged can depend on the inflationary environment described by the trend inflation rate.[4] High trend inflation would rapidly erode the real price of a firm's output. The firm would make frequent revisions in its nominal price and correspondingly less adjustment in output. Thus, the extent to which price stickiness can lead to persistence in output fluctuations depends inversely on trend inflation.

Since nominal price stickiness should be less important in high inflationary environments,

[1] According to NBER, the average duration of expansion in the postwar U.S. economy is 50 months and the average duration of contraction is 11 months. Campbell and Mankiw (1987) document the persistent nature of economic fluctuations in the industrialized countries.

[2] The work of Taylor (1979, 1980), Sheshinski and Weiss (1977, 1983), Blanchard (1991), Calvo (1983), Mankiw (1985), Akerlof and Yellen (1985), and Blanchard and Kiyotaki (1987) has provided the microfoundations for a theory of demand-driven aggregate fluctuations.

[3] Akerlof and Yellen (1985) and Mankiw (1985) show how price stickiness may arise as an outcome of profit-maximizing behaviour by firms.

[4] Romer (1990) formalizes this idea. He combines the two approaches to staggered pricing, namely, the *time-dependent* and the *state-dependent* pricing structures, in a static general-equilibrium model. Dotsey et al. (1998) accomplish this task in a dynamic general-equilibrium framework.

the sticky-price model implies that, *other things being equal, deviations of output from a long-run growth path should be less persistent in high-inflation countries than in low-inflation countries.*

This paper examines the above-noted implication empirically. Kiley (2000) also examines the issue of price-stickiness and business-cycle persistence. However, he ignores the issue of within-country time variation in inflation in the empirical implementation of the hypothesis. As discussed below, his conclusion is very different from mine. The analysis of this paper follows the tradition of Lucas (1973) and Ball et al. (1988), among others, who have demonstrated the usefulness of testing cross-sectional predictions of macroeconomic theories. I use annual data from 51 countries to investigate the relationship between inflation and persistence in output. The empirical implementation focuses on the issue of within-country time variation in inflation. From the viewpoint of empirical implementation, a high time variability of inflation within a country is an undesirable feature. This within-country time variation in inflation may not be captured by the average inflation rate. Two separate methods are used to control for this time variability. First, I identify the inflationary environment within a particular country by the long-run movement in its inflation rate. This is the trend inflation rate. Second, to control for high time variability in inflation, I exclude periods of extreme monetary instability. These periods are called episodes of hyperinflation. In this manner, the empirical implementation can better test the hypothesis. I use a panel approach and a two-stage approach for estimation.

My results do not support the hypothesis that countries with high trend inflation have less-persistent output fluctuations. For both quadratic detrended and H-P filtered international data, I find that the relationship between trend inflation and persistence in the deviations of output from its trend is statistically insignificant.

When I use average inflation (which ignores the within-country time variation in inflation) to characterize the inflationary environment, the data lend weak support to the hypothesis. However, the inverse relationship in the data is likely due to factors other than the price-stickiness mechanism. The exclusion of a few episodes of hyperinflation identified in Argentina, Brazil, and Israel makes the relationship between persistence in output fluctuations and average inflation statistically insignificant. These hyperinflationary episodes constitute less than 1 per cent of the entire panel data.

The empirical results of this paper have important implications. First, the sticky-price model is used as a microfoundation for an environment that can generate persistent effects of aggregate demand shocks (see, for example, Taylor (1980), Jeanne (1998), and Bergin and Feenstra (1998)).[5] The lack of empirical support for a key prediction of the sticky-price model

[5] Chari et al. (1996) implement Taylor's (1980) staggered pricing structure in a dynamic general-equilibrium (DGE) model. They find that, contrary to Taylor's insight, the staggered pricing structure can generate persistent real effects of monetary shocks *only if* one assumes that prices are exogenously sticky for a long period of time (about 2 1/2 years). The intuition behind their result is that if firms do not change the price of their product, then price must not be very sensitive to changes in marginal costs. In their DGE model this insensitivity of price to changes in marginal costs occurs when preferences have zero income effects (so that the labour supply response is not dampened) and the elasticity of labour supply (flat labour-supply curves) is implausibly high.

This result raises questions about the ability of the sticky-price framework to generate persistent effects of monetary shocks from a theoretical standpoint. Several recent papers show that Chari et al.'s result is particular. Bergin and Feenstra (1998) generate persistent output dynamics by making two modifications to the model of Chari et al. (1996): they introduce translog preferences instead of constant elasticity of substitution (CES) preferences, and they allow for an input-output structure where a firm's inputs are other firms' output. The first modification implies that price-setting rules are not a simple markup over a firm's own marginal cost; rather, they are influenced by other firms' prices. This interaction introduces strong linkages in price-setting rules across firms. The second modification implies that output prices of firms become an important component of marginal cost, thereby reducing the importance of labour input.

Likewise, Jeanne (1998) develops a DGE model where there is an interaction between product market nominal rigidity (sticky prices) and labour market real rigidity (efficiency wages). Jeanne builds on the static model of Ball and Romer (1988), who show that the presence of real rigidities can enhance the effect of nominal rigidities. He confirms the result of Ball and Romer (1988) in a DGE model. The period for which prices must remain sticky to generate quantitatively important persistence in output can be reduced substantially to three quarters.

Anderson (1998) and Huang and Liu (1998) emphasize the differences between staggered wage and staggered price contracts. Huang and Liu (1998) show that a key parameter governing the persistence properties suggests that while these models are important environments in which to study the impact effect of aggregate demand shocks, they might be less useful in explaining the persistent effect of those shocks. Second, the results indicate that persistence in output fluctuations is inelastic with respect to the average inflation rate over a very broad range, from 2—3 per cent to 60—70 per cent. Only periods of extreme monetary instability are informative about the prediction of the model examined here. This finding is problematic for the sticky-price model, because the model is used to study relatively stable environments.

Kiley (2000) examines the issue of price stickiness and business-cycle persistence. His conclusions, however, are very different from mine. Using the updated Ball et al. (1988) data set, Kiley finds that price stickiness decreases with average inflation across countries, and interprets this finding as being strongly supportive of the propagation mechanism based on sticky-price (menu-cost) models. He ignores the time variation in inflation across countries. Taking account of this time variation is important for an accurate empirical implementation of the theory. Further, where there is support for the hypothesis, I find a statistically significant inverted-U relationship between persistence and average inflation. To have a positive relationship up to a certain average inflation rate and a negative one thereafter is not consistent with a clear prediction of the sticky-price model.

Previous empirical research has focused on the contemporaneous output-inflation trade-off or the impact effect of aggregate demand shocks on output.[6] In particular, Ball et al. (1988) find that the *impact effect* of aggregate shocks is smaller in countries with high

in staggered wage (price) contract models is the elasticity of relative wage (price). The smaller the value of this parameter, the smaller the response of wages to demand shocks, the less the need to revise price, and the greater the persistence in output. For staggered wage models, this elasticity turns out to be small, whereas for staggered price models, as in Chari et al. (1996), it is large.

[6]Lucas (1973), Alberro (1981), Ball et al. (1988), Defina (1991), and Koelln et al. (1996).

inflation. My estimation results confirm that this is a robust finding. Therefore, my results suggest that there is statistical support for the prediction of the sticky-price model for one aspect of the business cycle: the impact effect of aggregate demand shocks. However, the data do not support the prediction of the model for another key aspect of business cycles: the persistence of output fluctuations.

Section 2 describes the theoretical framework for the empirical analysis, and Section 3 the empirical implementation issues and methodology. Section 4 describes the main results, Section 5 the results for OECD and non-OECD countries, and Section 6 the results from an alternative methodology (two-stage estimation).

2 Theoretical Framework

A typical sticky-price model has a monopolistically competitive market structure. A convenient way to introduce staggered price adjustments in this environment is to use the price-setting structure of Calvo (1983). Under this pricing structure, each firm can adjust its price in a given period with probability $1 - \phi$. Therefore, with probability ϕ a firm charges the predetermined nominal price and satisfies the demand at that price. This probability is assumed to be independent across time and identical across firms. The symmetric market structure implies that all firms that change their price charge the same price. The average duration for which the price quotation of a firm stays unchanged is $\frac{1}{1-\phi}$. At any time, the fraction of firms that adjusted their price k periods ago is given as $\theta_k = (1 - \phi)\phi^k$. The aggregate price level takes the form[7]:

$$P_t = (\sum_{k=0}^{\infty} \theta_k (P_{t-k}^*)^{1-\epsilon})^{\frac{1}{1-\epsilon}}, \tag{1}$$

where P_{t-k}^* is the optimal price chosen by a firm k periods ago and ϵ is the elasticity of demand. Using the expression for θ_k and (1), the aggregate price level can be shown to take the following recursive form:

$$P_t = [(1-\phi)(P_t^*)^{1-\epsilon} + \phi(P_{t-1})^{1-\epsilon}]^{\frac{1}{1-\epsilon}}. \tag{2}$$

This expression shows that the aggregate price level has two components. First, a fraction $(1-\phi)$ of firms charge the optimal price P_t^* at time t. Second, the remaining fraction of firms ϕ charge the predetermined price P_{t-1}. Thus, the aggregate price level is completely characterized by $\{P_t^*, P_{t-1}\}$. This set-up allows for an arbitrary degree of price rigidity: If $\phi = 1$, then prices are completely rigid. If $\phi = 0$, then prices are completely flexible. Therefore, the degree of nominal rigidity is given by the parameter ϕ. The inflation rate is given as $\pi_t = \frac{P_t}{P_{t-1}} - 1$. I use the standard quantity equation to represent the aggregate demand of the economy as

$$y_t = m_t - p_t, \tag{3}$$

[7]Examples of DGE models with staggered pricing are described in Yun (1996), Woodford (1996), Rotemberg (1996), King and Wolman (1996, 1998), and Dotsey et al. (1998).

where $y_t = log(Y_t)$, $m_t = log(M_t)$, and $p_t = log(P_t)$ are real output, nominal demand, and the price level in logs, respectively. This relationship can be obtained from a representative consumer's optimization problem in which the consumer must hold money to finance purchases. The nominal demand, m_t, in equation (1) is equal to the nominal money stock.[8] The driving process is the rate of growth of money stock, m_t, which is assumed to follow:

$$\mu_t = \bar{\mu} + \rho_m \mu_{t-1} + \epsilon_{mt}, \qquad \epsilon_{mt} \sim i.i.d.(0, \sigma_{mt}). \tag{4}$$

In this environment, Jeanne (1998) shows that the deviations of output from trend $y_t^d = y_t - y_t^T$ can be expressed as a process:

$$y_t^d = \Phi(\phi) y_{t-1}^d + \Psi(\phi)\mu_t. \tag{5}$$

The coefficient $\Phi(\phi)$ characterizes persistence in output when the degree of nominal rigidity ϕ is fixed. Following the theoretical discussion, the relationship between the persistence parameter and ϕ is given as

$$\frac{\partial \Phi(\phi)}{\partial \phi} > 0. \tag{6}$$

That is, an increase in the fraction of firms that keep their prices fixed leads to an increase in the persistence of output deviations. The output deviations become more persistent since an increase in ϕ will lead to an increase in the average duration for which a firm's price is

[8] Two assumptions underlie this formulation: money-demand distortions due to a positive nominal interest rate are negligible, and fiscal policy is in the form of lump-sum taxes and transfers.

fixed. Therefore, the firm responds by adjusting its output. In the Calvo (1983) framework, the probability of a price change $1-\phi$ is exogenously fixed. Romer (1990) endogenizes this frequency in terms of the underlying features of the economy as

$$\phi = \phi(\pi^T, \sigma_m^2, F, K), \tag{7}$$

where π^T is the trend inflation rate, σ_m^2 is the variability of nominal shocks, F is the fixed cost of changing prices, and K can be interpreted as the degree of real rigidity in the market.[9] Further,

$$\frac{\partial \phi}{\partial \pi^T} < 0. \tag{8}$$

That is, a rise in trend inflation will lead to a decrease in the average duration $\frac{1}{1-\phi}$ for which prices are fixed.[10]

To summarize, the theoretical framework described above shows that a higher trend inflation will increase the frequency of price adjustments among firms. Therefore, most

[9] In equilibrium, the trend inflation rate is linked to the exogenous trend growth in the money stock. In related work, Gray (1978), Canzoneri (1980), and Gray and Kandil (1991) endogenize the length of labour contracts in terms of variability of nominal demand. The trend inflation level does not influence the length of the contracts, since labour contracts could be indexed with respect to the trend inflation level.

[10] In Romer's (1990) model, each firm chooses the probability of adjusting (or, equivalently, the probability of not adjusting) its price, taking as given the other firms' decisions. In equilibrium, owing to symmetry, all firms choose the same probability of adjusting their nominal price. Factors other than the trend inflation rate influence a firm's choice in the following manner:

$$\frac{\partial \phi}{\partial \sigma_m^2} < 0, \qquad \frac{\partial \phi}{\partial F} > 0, \qquad \frac{\partial \phi}{\partial K} > 0.$$

More recently, Dotsey et al. (1998) have synthesized time-dependent pricing and state-dependent pricing in a DGE framework. The staggered pricing structure in their paper is richer than Romer's (1990), because it allows ϕ to be firm-specific. Thus, at any given point in time, there are different vintages of firms according to when they last adjusted their price. In this environment, the probability of price adjustment rises with an increase in steady-state inflation or the trend inflation rate.

firms in the economy adjust prices in response to demand shocks. Firms that do not change prices make adjustments in output. This proportion decreases with an increase in trend inflation, leading to a less-persistent deviation of output from its long-term growth path. The implication is that a country with a high trend level of inflation should exhibit less-persistent fluctuation in output from its long-term growth path than a country with a low trend level of inflation.

3 Empirical Implementation

I postulate an empirical counterpart of equation (5) where the detrended output follows a general second-order linear process:

$$y_{it}^{d} = \Phi^{1}(\phi(f(\Pi_{it})))y_{it-1}^{d} + \Phi^{2}(\phi(f(\Pi_{it})))y_{it-2}^{d} + \Psi(\phi(f(\Pi_{it})))\mu_{it} + u_{it}, \tag{9}$$

where (i denotes the country) $i = 1, ..., N$ and $t = 1, ...T_i$; y_{it}^{d} is the detrended output; $f(\Pi_{it})$ denotes the inflationary environment within country i; and μ_{it} represents fluctuations in aggregate demand; u_{it} is the error term.

The empirical implementation of the hypothesis that countries with a high trend inflation rate have a low persistence in output fluctuations raises several important issues.

The ideal environment for testing the cross-sectional implication of the theory is to have a set of countries with different levels of steady inflation rates. In other words, a high cross-sectional variation and minimal within-country variation in inflation would be a desirable feature of the data. I could then make the assumption that the same sticky-price model applies to individual countries and, thus, examine how differences in inflation would lead to different outcomes. Unfortunately, the countries that have high average inflation have experienced enormous time variation in inflation. The empirical analysis should take into account this time variation or the changes in the inflationary environment of individual countries. By reducing the effect of the within-country variation in inflation, my empirical implementation better tests the theoretical predictions of the model. One way to reduce this time variation within countries is to consider long-run movements in inflation. A second way is to exclude periods of extreme time variation—that is, episodes of hyperinflation within countries. I discuss the latter approach in Section 3.1.

I consider two representations of the inflationary environment $f(\Pi_{it})$:

$$f(\Pi_{it}) = \begin{cases} \pi_{it}^{T} & : \text{the long-run (trend) inflation rate,} \\ \bar{\pi}_i & : \text{the average inflation rate.} \end{cases} \tag{10}$$

The long-run (trend) inflation rate is time-varying. I obtain the long-run component of inflation in two ways. First, I use the H-P filter to remove the short-run (high-frequency) component of inflation. The low-frequency component represents the long-run or the trend inflation rate. Second, I identify the long-run component of inflation as the quadratic time-fitted value of inflation. I also use average inflation to emphasize the consequences of neglecting time variation in inflation. The long-run (trend) inflation and the average inflation rate characterize the inflationary environment for individual countries.

The data on a monetary aggregate that captures the shifts in the stance of monetary policy—one source of aggregate demand shocks—are not available for most countries. Therefore, following other related studies in this literature, I use nominal output growth to represent exogenous aggregate demand fluctuations.[11] I use the unanticipated nominal output growth as an alternative representation to capture the exogenous aggregate demand fluctuations. I assume a rational forecasting rule to obtain the unanticipated nominal growth for each country. I obtain x_{it}^{UA} as the residual from the regression of x_{it} on a constant, x_{it-1}, and x_{it-2}:

$$\mu_{it} = \begin{cases} x_{it} & : \text{ nominal output growth rate,} \\ x_{it}^{UA} & : \text{ unanticipated nominal output growth.} \end{cases} \quad (11)$$

To obtain the cyclical component of real output (y_{it}^{d}), I use the H-P filter and quadratic time detrending (QTD). These procedures are used extensively in the business-cycle literature.[12]

The functions $\Phi^1(\phi(f(\Pi_{it})))$ and $\Phi^2(\phi(f(\Pi_{it})))$ reflect that the trend inflation affects the dynamic behaviour of output by influencing the frequency of price adjustment. The term $\Psi(\phi(f(\Pi_{it})))$ captures the impact effect of nominal demand shocks. I take a linear approximation of these functions to obtain the following specification:

$$y_{it}^{d} = \alpha + \Phi_0^1 y_{it-1}^{d} + \Phi_\pi^{11} f(\Pi_{it}) y_{it-1}^{d} + \Phi_0^2 y_{it-2}^{d} + \Phi_\pi^{21} f(\Pi_{it}) y_{it-2}^{d} + \Psi_0 \mu_{it} + \Psi_\pi f(\Pi_{it}) \mu_{it} + u_{it}. \quad (12)$$

This specification has the advantage that it nests both the impact effect and the persistence in output fluctuations due to aggregate demand shocks. In my analysis I consider both

[11] See Akerlof et al. (1988) for difficulties that arise with the use of nominal output growth.
[12] See Guay and St-Amant (1996) for a critical evaluation of filtering procedures.

AR(1) and AR(2) output dynamics. I describe the measure of persistence for the AR(1) in Section 3.1.

3.1 AR(1) output dynamics

For AR(1) output dynamics, I restrict the coefficients $\Phi_{\pi}^{21} = 0$ and $\Phi_0^2 = 0$ in equation (12), writing the equation in the stacked form to obtain

$$\mathbf{y_t^d} = \alpha + \Phi_0^1 \mathbf{y_{t-1}^d} + \Phi_{\pi}^{11} \mathbf{f}(\Pi_\mathbf{t}) \mathbf{y_{t-1}^d} + \Psi_0 \mu_\mathbf{t} + \Psi_\pi \mathbf{f}(\Pi_\mathbf{t}) \mu_\mathbf{t} + \mathbf{u_t}. \tag{13}$$

The definition of persistence in output that I consider in this section is: *how closely the deviation of output from its trend in the current period is related to the deviation of output from its trend in the previous period.* That is,

$$\frac{\partial \mathbf{y_t^d}}{\partial \mathbf{y_{t-1}^d}} = \Phi_0^1 + \Phi_\pi^{11} \mathbf{f}(\Pi_\mathbf{t}). \tag{14}$$

This paper focuses on how the inflationary environment affects persistence in output, that is, the cross-partial:

$$\frac{\partial(\partial \mathbf{y_t^d})}{\partial \mathbf{f}(\Pi_\mathbf{t})(\partial \mathbf{y_{t-1}^d})} = \Phi_\pi^{11}. \tag{15}$$

Therefore, the coefficient of the interaction variable of inflationary environment and the lagged deviation of output from its long-run growth path or trend is of central interest. The sticky-price model implies that an increase in the trend inflation rate will make firms adjust their prices more frequently and, therefore, make less adjustment in output. Thus, deviations of output from its trend will be less persistent. The theoretical prediction of this class of models is that Φ_π^{11} should be *negative*.

As stated earlier, the focus of a large body of previous research, with the exception of Kiley (2000), has been on the impact effect of aggregate demand shocks. In my empirical specification, the impact effect is captured by the coefficients Ψ_0 and Ψ_π:

$$\frac{\partial \mathbf{y_t^d}}{\partial \mu_\mathbf{t}} = \Psi_0 + \Psi_\pi \mathbf{f}(\Pi_\mathbf{t}). \tag{16}$$

In particular, the influence of the inflationary environment on the impact effect of aggregate demand shocks is given by:

$$\frac{\partial(\partial \mathbf{y_t^d})}{\partial \mathbf{f}(\Pi_\mathbf{t})(\partial \mu_\mathbf{t})} = \Psi_\pi. \tag{17}$$

Sticky-price models predict that Ψ_π should be negative. That is, the impact effect of aggregate demand shocks should be less in high-inflation countries. Ball et al. (1988) test this prediction.

I estimate the following specification for the panel of 51 countries allowing for cross-sectional heteroscedasticity:[13]

$$y_{it}^d = \alpha + \Phi_0^1 y_{it-1}^d + \Phi_\pi^{11} f(\Pi_{it}) y_{it-1}^d + \Psi_0 \mu_{it} + \Psi_\pi f(\Pi_{it}) \mu_{it} + u_{it}. \quad (18)$$

3.2 Data

The sample consists of annual International Financial Statistics (IFS) data on real and nominal GDP/GNP for 51 countries. The overall sample period is 1950—96. However, several countries have shorter time periods, depending upon the availability of data. The inflation rate for a country is the growth rate of its GDP deflator. Table 1 includes the 43 countries in the Ball et al. (1988) study.[14] I add eight countries: South Korea, New Zealand, India, Pakistan, Paraguay, Thailand, Malaysia, and Kenya.

[13]Table 1 provides information on the time period and the average inflation for individual countries. The presence of lagged dependent variables raises the issue of fixed effects addressed in the recent literature on dynamic panel-data models (see Arellano and Bond (1991) and Baltagi (1995)). Typically, data in (log) levels are first-differenced to eliminate the fixed effect. When examining growth issues in a dynamic panel setting (see, for example, Easterly et al. (1997)), the variables of the first-difference specification have a natural interpretation of current and lagged growth rates. However, in the present case this approach is not useful since I begin the analysis with detrended data. First-differencing the detrended data does not have any meaningful economic interpretation. Recent work by Lancaster (1999) and Lancaster and Aiyar (1999) uses Bayesian techniques, which do not involve first-differencing in the estimation of a dynamic panel-data model. Exploration of Lancaster's technique for an alternative estimation of the empirical specification examined here and to other related business-cycle issues is left for future research.

[14]I thank Michael Kiley for generously providing the updated Ball et al. data set.

Table 1: Data summary

No.	Country	Avg. Inflation (%)
1	Argentina	100.1
2	Australia*	7.5
3	Austria*	4.4
4	Belgium*	5
5	Bolivia	6
6	Brazil	123
7	Canada*	5.5
8	Colombia	21.7
9	Costa Rica	19.3
10	Denmark	6.3
11	Dominican Republic	15.1
12	Ecuador	25.4
13	El Salvador	12.3
14	Finland	7.4
15	France*	6.6
16	Germany*	3.6
17	Greece*	16.8
18	Guatemala	12.9
19	Iceland*	24.8
20	India	8.5
21	Iran	20.3
22	Ireland*	8.1
23	Israel	43.5
24	Italy*	10.9
25	Jamaica	20.6
26	Japan*	3.7
27	Kenya	10.9
28	Malaysia	4.9
29	Mexico	29.3
30	Netherlands*	3.8
31	New Zealand*	8.7
32	Norway*	6.3
33	Pakistan	9.8
34	Panama	4
35	Paraguay	17.5
36	Philippines	11.9
37	Portugal*	15.3
38	Singapore	4.1
39	South Africa	12.9
40	South Korea	11.1
41	Spain*	10.4
42	Sweden*	7.4
43	Switzerland*	3.5
44	Thailand	6.4
45	Tunisia	7.3
46	UK*	8.3
47	US*	5.4
48	Venezuela	20.8
49	Zaire	78.7

* indicates OECD country (22 countries). Mexico and South Korea became OECD members in 1994 and 1996, respectively.

4 Results

I first estimate a restricted version of (18) which serves as a benchmark case. This version is given as

$$y_{it}^{d} = \alpha + \Phi_{0}^{1} y_{it-1}^{d} + \Phi_{\pi}^{11} f(\Pi_{it}) y_{it-1}^{d} + u_{it}. \tag{19}$$

This specification puts minimal structure on the regression, since it does not include the nominal growth terms. The results are for all 51 countries. To be consistent across the two detrending methods, I obtain the trend inflation using the same detrending method as for output.[15] The p-value shows the exact significance level below which the null hypothesis ($H_0 : \Phi_{\pi}^{\mathbf{11}} = 0$) cannot be rejected. The symbol '*' denotes that the coefficient is statistically different from zero at the 5 per cent level of significance. As indicated in (10), I use both the trend inflation (deterministic and stochastic) and average inflation to characterize the inflationary environment of the countries.

The results in Table 2 indicate that the coefficient of interest, $\Phi_{\pi}^{\mathbf{11}}$, is statistically insignificant at the 5 per cent level in the data. This statistical significance holds for both the QTD and H-P cases. For the H-P data, even the sign of the coefficient is positive, which is opposite from the predicted one.

The results from Table 2 indicate that countries with higher trend inflation do not exhibit less-persistent output fluctuations. This evidence does not support the prediction of the

[15] For example, when output is detrended using the QTD method, the long-run trend inflation is also obtained by this method. Similarly, when output is detrended using the H-P filter, trend inflation is obtained as the low-frequency component of the actual inflation rate. The regression assumption is that the trend inflation is proxied by either the deterministic trend or the stochastic trend.

Table 2: Persistence and trend inflation

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.833*	0.017	0.000	0.714*	0.020	0.000
Φ_π^{11}	-0.169	0.110	0.104	0.009	0.109	0.929
α	-0.0002	0.0005	0.723	-0.0001	0.0005	0.977

sticky-price model, which is used to explain persistence in output fluctuations as described in Section 2.

Table 3 lists the results where the average inflation rate is used to characterize the inflationary environment within countries. The QTD data support the hypothesis. The coefficient $\Phi_\pi^{\mathbf{11}}$ is negative and statistically significant. The empirical support for the model in this case is extremely fragile for two reasons. The first regards controlling for within-country time variation in inflation. This point can be seen by comparing results for QTD data in Tables 2 and 3. I investigate this point in detail in Section 4.1. The second reason regards the methodology used to obtain the cyclical component of output. The results for the H-P filtered data in Table 3 indicate that the coefficient $\Phi_\pi^{\mathbf{11}}$ is statistically insignificant.

I now estimate (18) where actual nominal output growth, x_t, is introduced as a regressor to explicitly capture fluctuations in aggregate demand and their impact effect on output.[16] The results for the relationship between trend inflation and persistence in output are reported in Table 4. The estimate of coefficient $\Phi_\pi^{\mathbf{11}}$ is statistically insignificant for both QTD and H-P filtered data. These results are similar to the ones reported in Table 2. They indicate that a higher trend inflation rate does not diminish persistence in output fluctuations.

Table 5 presents the results where average inflation is used. The results are similar to

[16] In using nominal output growth as a regressor, I follow the literature on cross-country studies of output-inflation trade-off, for example, Ball et al. (1988).

Table 3: Persistence and average inflation

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.861*	0.020	0.000	0.737*	0.024	0.000
Φ_π^{11}	-0.434*	0.145	0.003	-0.195	0.159	0.221
α	-0.0002	0.0005	0.704	-0.0001	0.0005	0.983

Table 4: Persistence and trend inflation

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.833*	0.017	0.000	0.722*	0.020	0.000
Φ_π^{11}	-0.169	0.110	0.104	-0.093	0.106	0.377
Ψ_0	0.064*	0.006	0.000	0.077*	0.006	0.000
Ψ_π	-0.046*	0.005	0.000	-0.069*	0.005	0.000
α	-0.007*	0.0009	0.000	-0.008*	0.0008	0.000

those in Table 4. For QTD data, the null hypothesis $(H_0 : \Phi_\pi^{\mathbf{11}} = 0)$ is rejected at the 5 per cent level. That is, a higher average inflation inversely affects the persistence in deviations of output from its trend. This support for the hypothesis, as found without x_t, is fragile with respect to controlling for within-country time variation in inflation and with respect to the particular detrending method. For instance, in H-P filtered data the coefficient $\Phi_\pi^{\mathbf{11}}$ is statistically insignificant.

The coefficient ψ_π, which captures the impact effect of aggregate demand shocks, is

Table 5: Persistence and average inflation

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.843*	0.020	0.000	0.722*	0.024	0.000
Φ_π^{11}	-0.305*	0.152	0.045	-0.069	0.166	0.677
Ψ_0	0.054*	0.006	0.000	0.047*	0.006	0.000
Ψ_π	-0.064*	0.009	0.000	-0.061*	0.009	0.000
α	-0.005*	0.0009	0.000	-0.005*	0.0008	0.000

negative and statistically significant in Tables 4 and 5.[17] Therefore, countries with higher trend inflation do have a lower impact effect of aggregate demand shocks. This result is consistent with the finding of Ball et al. (1988).[18]

To summarize, the results from Tables 2—5 indicate that when the within-country time variation in inflation is accounted for, a key prediction of the sticky-price model is not supported in the international data. When I do not control for the within-country time variation in inflation by examining the relationship between average inflation and output persistence, there is one case for QTD output where the hypothesis is supported. However, this support is extremely fragile. In Section 4.1, I investigate an alternative way of controlling for the within-country time variation in inflation. This method enables me to further establish the fragility of support for the hypothesis found in the QTD data.

4.1 Episodes of hyperinflation

As stated earlier, a key assumption implicit in the hypothesis is that countries are at different levels of steady inflation rates. In the data set, however, high-inflation countries have enormous time variation (or instability) in inflation. For those countries, in particular, average inflation may not capture the movement in the inflationary experience. The importance of this aspect is shown in the results reported in Tables 2—5, especially for the QTD data. Given the feature of existing data described above, the question is when average inflation can be a more appropriate measure of the inflationary environment and, therefore, more

[17] The intercept term is significant because x_{it} is a growth rate, while y_{it}^d and y_{it-1}^d are detrended variables.

[18] The results for the case, when I restrict attention to the countries in the Ball et al. (1988) study, are qualitatively similar. As found in Defina (1991), the impact effect is not sensitive to controlling for within-country time variation in inflation. Furthermore, similar results hold when I use unanticipated nominal output growth to capture fluctuation in aggregate demand and its impact effect.

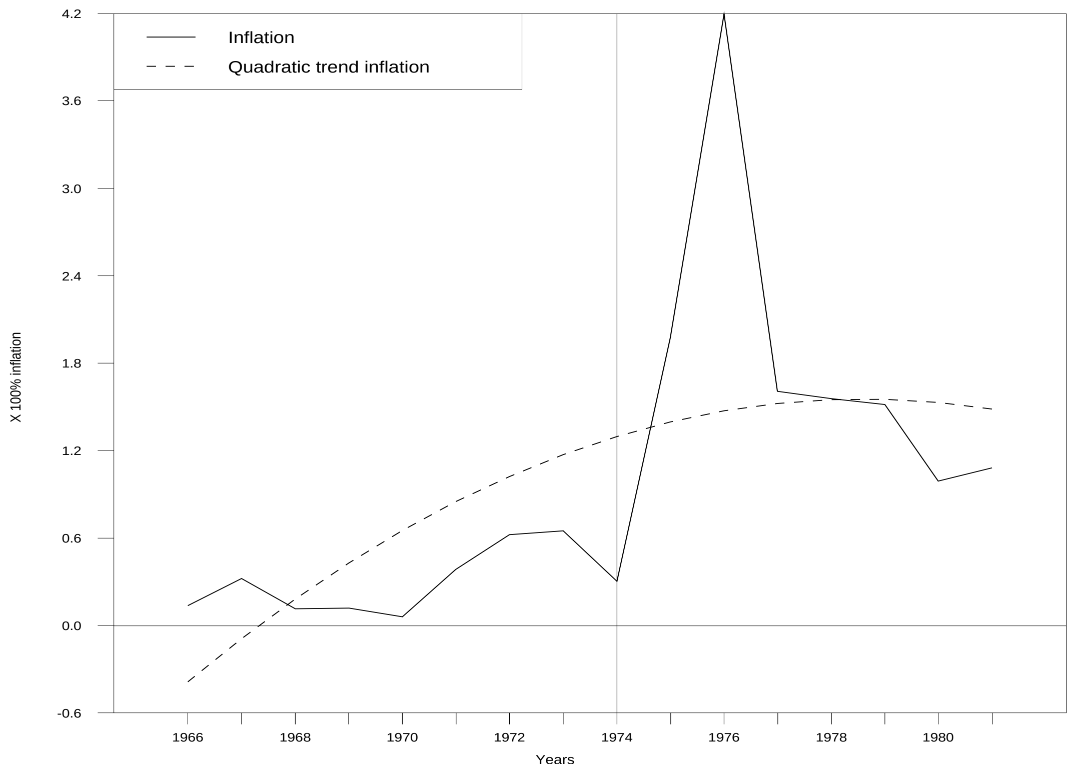

Figure 1: Argentina

appropriate for testing the hypothesis. In this section I pursue the strategy of reducing the within-country time variation by excluding only episodes of hyperinflation. Episodes of hyperinflation are typically periods of extreme monetary instability. They are driven by forces that are not part of any sticky-price model that is designed to explain the persistent nature of output fluctuations.

Figures 1—3 show that the inflationary experiences of Argentina, Brazil, and Israel are marked by episodes of hyperinflation. For those countries, in particular, average inflation may not be a reasonable summary statistic. Moreover, the causes of the hyperinflationary episodes may be linked to drastic shifts in monetary policy or sudden institutional changes.

I identify periods of relatively stable inflation within the high-inflation countries by ex-

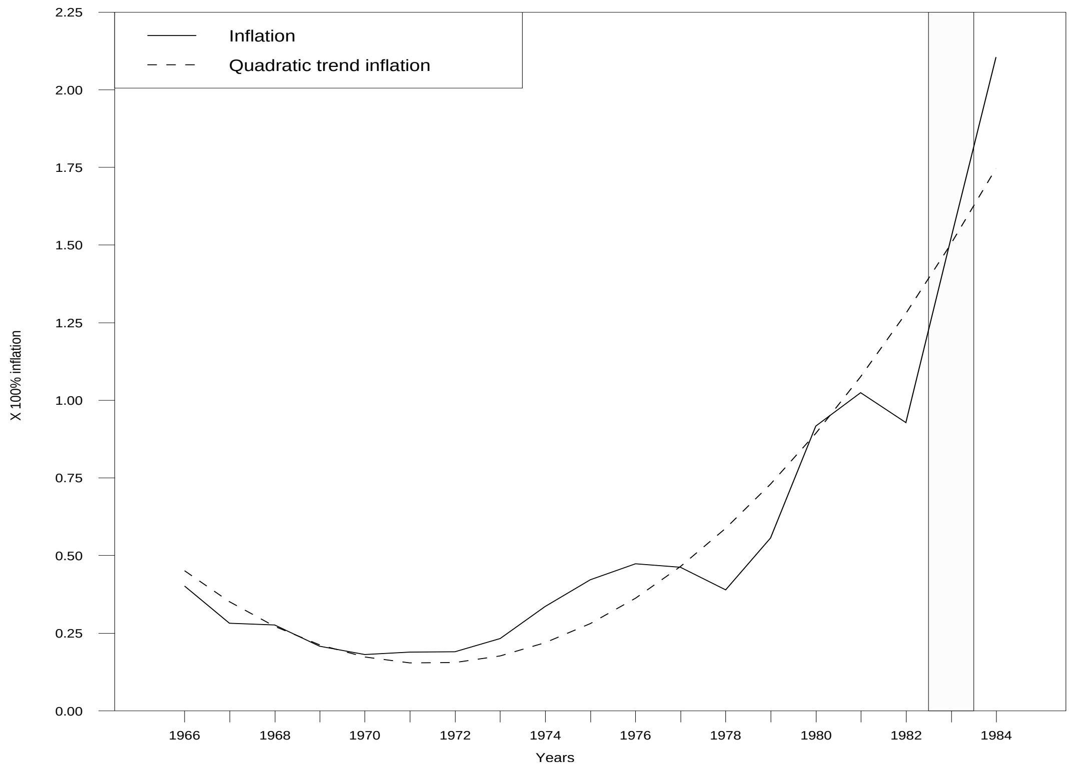

Figure 2: Brazil

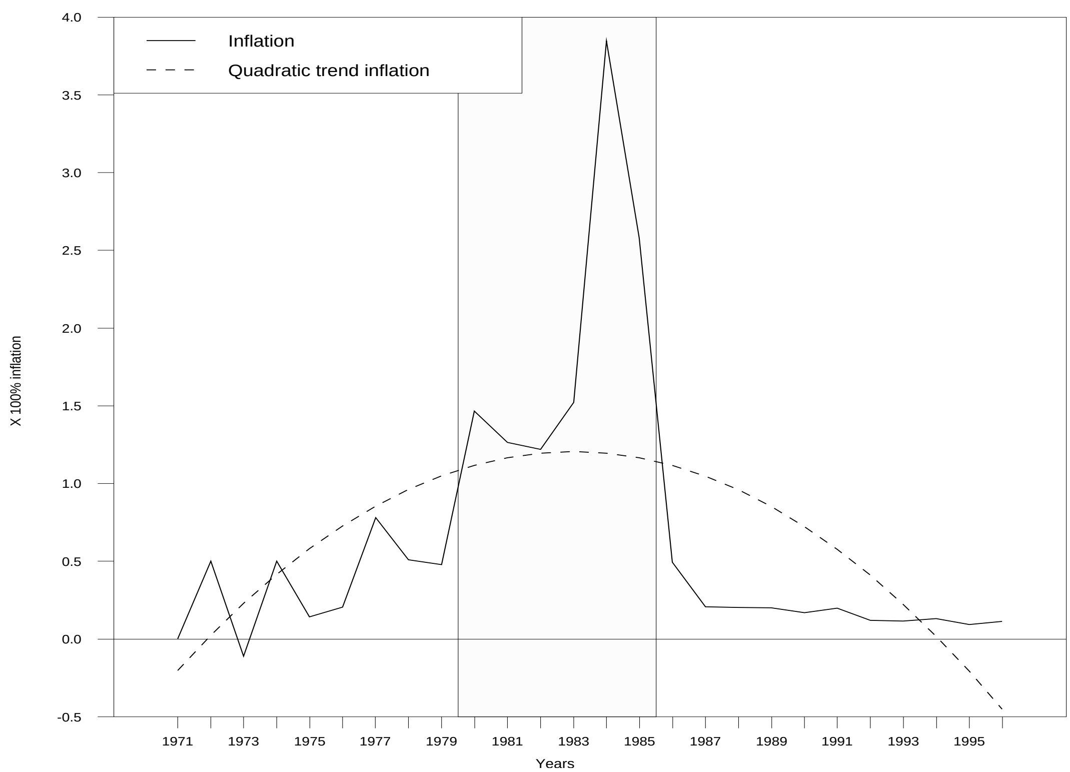

Figure 3: Israel

Table 6: Episodes of hyperinflation

Country	Period	Years	$\bar{\pi}$	σ_π	Relative volatility
Argentina	Full sample	1966−1981	0.977	1.06	4.61
	Unstable	1974−1981	1.653	1.15	5.0
	Stable	1966−1973	0.30	0.23	1
Brazil	Full sample	1966−1984	0.583	0.51	1.38
	Unstable	1984	2.105	−	−
	Stable	1966−1983	0.499	0.37	1
Israel	Full sample	1971−1996	0.652	0.90	4.09
	Unstable	1980−1985	1.983	1.04	4.73
	Stable	1971−79, 1986−96	0.252	0.22	1

amining the volatility of inflation in subsample periods (see Table 6). In Argentina, the relative volatility of inflation in the unstable period is five times that of the stable period. In Israel, the relative volatility of the unstable period is 4.7 times that of the stable period. In the unstable periods, the inflation rate was in excess of three standard deviations of the average inflation of the stable period.[19] In Israel, the liberalization process that lifted capital controls and introduced foreign exchange indexed accounts triggered the hyperinflationary phase between 1980—85. These unstable periods constitute episodes of hyperinflation. These episodes are: Argentina (1975—81, 7 years), Brazil (1984, 1 year), and Israel (1980—85, 6 years). *These observations constitute less than 1 per cent of the entire panel data.*

The results are given in Tables 7 and 8. The results are markedly different compared with the results in Tables 3 and 5. The coefficient Φ_π^{11} is statistically insignificant for the QTD case. For the H-P filtered data, the coefficient $\Phi_\pi^{\mathbf{11}}$ not only remains insignificant but has an opposite sign from the prediction of the theoretical model.

The results for the case where I exclude the episodes of hyperinflation strongly support my earlier conclusions. That is, in the international data the inverse relationship between

[19]This rule is similar to the one proposed by Tsay (1988) to identify outliers in time series.

Table 7: Persistence and average inflation: Excluding episodes of hyperinflation

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.824*	0.023	0.000	0.697*	0.027	0.000
Φ_π^{11}	-0.085	0.187	0.646	0.175	0.214	0.415
α	-0.0002	0.0005	0.716	0.0003	0.0005	0.951

persistence in output fluctuations and trend inflation is not statistically significant. This finding is difficult to reconcile with the sticky-price model in which the frequency of price changes depends on the inflationary environment. Suppose that average inflation is, in fact, an appropriate measure of the inflationary environment: the results then indicate that the inverse relationship between persistence and average inflation found for the QTD case exists *only if* one takes into account the periods of hyperinflation. However, the theoretical model of price stickiness assumes a stable and stationary environment. That is, the model is not about hyperinflationary phenomena.

Where there is support for the hypothesis, it is more likely that the inverse relationship arises because of other factors. For instance, the autocorrelation in QTD output for Argentina is -0.42 and in H-P filtered output it is -0.20. If price stickiness indeed disappears at very high levels of trend inflation, then the sticky-price model predicts that the autocorrelation should be zero and not negative.

Table 8: Persistence and average inflation: with x_t, excluding episodes of hyperinflation

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.792*	0.022	0.000	0.676*	0.026	0.000
Φ_π^{11}	0.133	0.19	0.485	0.346	0.218	0.112
Ψ_0	0.134*	0.010	0.000	0.127*	0.009	0.000
Ψ_π	-0.390*	0.035	0.000	-0.393*	0.034	0.000
α	-0.011*	0.001	0.000	-0.010*	0.0009	0.000

4.2 AR(2) output dynamics

I estimate a more general specification involving AR(2) output dynamics:

$$y_{it}^d = \alpha + \Phi_0^1 y_{it-1}^d + \Phi_\pi^{11} f(\Pi_{it}) y_{it-1}^d + \Phi_0^2 y_{it-2}^d + \Phi_\pi^{21} f(\Pi_{it}) y_{it-2}^d + \Psi_0 \mu_{it} + \Psi_\pi f(\Pi_{it}) \mu_{it} + u_{it} \tag{20}$$

Although this approach captures richer output dynamics, it complicates the definition of persistence. However, the coefficients of the interaction terms on lagged output, Φ_π^{11} and Φ_π^{21}, capture the effect of the trend inflation rate on output dynamics.[20] I directly test for the joint significance of these coefficients. That is, $H_0 : \Phi_\pi^{11} = 0$ and $\Phi_\pi^{21} = 0$.

Table 9 shows the results for the trend inflation rate. For the H-P case, the χ^2 statistic for the joint test is 1.96 with a p-value of 0.375. This implies that the null hypothesis of both coefficients being zero cannot be rejected for any significance level below 37.5 per cent. For QTD, the χ^2 statistic is 5.26 with a p-value of 0.09. That is, the null of both coefficients being zero cannot be rejected for any significance level below 9 per cent.

Table 10 shows the results for average inflation. For QDT, the χ^2 statistic is 19.12 with

[20]Because of space limitations, the results for $\mu_{it} = x_{it}$. The results for unanticipated demand shocks, $\mu_{it} = x_{it}^{UN}$, are qualitatively similar.

Table 9: Persistence and trend inflation: AR(2)

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	1.136*	0.026	0.000	0.964*	0.026	0.000
Φ_π^{11}	-0.276*	0.124	0.026	-0.170	0.124	0.172
Φ_0^2	-0.356*	0.026	0.000	-0.323*	0.027	0.000
Φ_π^{21}	0.104	0.129	0.421	0.070	0.141	0.619
Ψ_0	0.056*	0.006	0.000	0.069*	0.006	0.000
Ψ_π	-0.039*	0.005	0.000	-0.062*	0.005	0.000
α	-0.006*	0.0009	0.000	-0.007*	0.0008	0.000

a p-value of 0.0001. Since the null hypothesis is strongly rejected, there is a statistically significant relationship between second-order output dynamics and average inflation. For H-P, the χ^2 statistic value is 5.08 with a p-value of 0.08. That is, the null cannot be rejected at the 5 per cent level of significance but it is rejected at the 10 per cent level. Similar to my findings for the first-order output dynamics, this support for a relationship between second-order output dynamics and average inflation is very fragile. When I exclude the episodes of hyperinflation identified earlier, the results change drastically (see Table 11). For the QTD data, the χ^2 statistic value for the joint test is 4.12 with a p-value of 0.127. That is, the null cannot be rejected at any significance level below 12.7 per cent. For the H-P filtered data, the exclusion of episodes of hyperinflation gives a χ^2 statistic value of 0.13 with a p-value of 0.936.

The results for the AR(2) dynamics support my earlier results and conclusions. The results strongly suggest that, contrary to the prediction of the sticky-price model, trend inflation does not have any significant effect on the persistence properties of output. Further, average inflation has a very fragile negative influence on persistence, particularly for the QTD

Table 10: Persistence and average inflation: AR(2)

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	1.206*	0.030	0.000	1.013*	0.031	0.000
Φ_π^{11}	-0.735*	0.168	0.000	-0.397*	0.186	0.033
Φ_0^2	-0.399*	0.029	0.000	-0.371*	0.030	0.000
Φ_π^{21}	0.353*	0.169	0.037	0.309	0.1183	0.092
α	-0.00005	0.0005	0.914	0.00002	0.0004	0.960

data. This fragile relationship is driven entirely by a few episodes of hyperinflation.

5 OECD and Non-OECD Countries

By dividing the countries in the sample into OECD and non-OECD categories one can determine whether the inverse relationship between trend inflation and persistence in output fluctuations exists within countries that belong to a particular economic group.

The results for the OECD countries are given in Tables 12 and 13.[21] The OECD countries have experienced less within-country time variation in inflation. This characteristic makes the set of OECD countries a favourable group for testing the hypothesis. Neither trend inflation nor average inflation affects the persistence in output fluctuations within the OECD countries. This finding holds for QTD and H-P data.[22]

[21] There are 22 OECD countries in the sample (see Table 1). Cross-sectional variability exists in the average inflation rates within the OECD countries. For example, Canada and the United States. have low average inflation rates of 4.4 per cent and 4.3 per cent, respectively, whereas Portugal and Iceland have high average inflation rates of 14.5 per cent and 24 per cent, respectively, over the sample period. The results where unanticipated nominal output growth is used to capture the impact effects are qualitatively similar.

[22] For the AR(2) case, the null hypothesis $H_0 : \mathbf{\Phi_\pi^{11}} = 0$, $\mathbf{\Phi_\pi^{21}} = 0$ is not rejected. The χ^2 test statistic value is 2.46 with a p-value of 0.300, and 1.79 with a p-value of 0.407 for the H-P and QDT cases, respectively. The results when average inflation is considered are similar. The χ^2 statistic value is 4.48 with a p-value of 0.11, and 2.97 with a p-value of 0.226 for the H-P and QDT cases respectively. These results are not reported here but are available.

Table 11: Persistence and average inflation: AR(2), excluding episodes of hyperinflation

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	1.130*	0.036	0.000	0.946*	0.037	0.000
Φ_π^{11}	-0.100	0.279	0.719	0.165	0.288	0.566
Φ_0^2	-0.362	0.036	0.000	-0.343	0.037	0.000
Φ_π^{21}	0.082	0.272	0.762	0.112	0.289	0.697
α	-0.00008	0.0005	0.875	0.00003	0.0004	0.948

Table 12: Persistence and trend inflation: with x_t, OECD countries, N=22

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.740*	0.030	0.000	0.629*	0.037	0.000
Φ_π^{11}	0.104	0.361	0.774	0.282	0.248	0.528
Ψ_0	0.245*	0.018	0.000	0.264*	0.016	0.000
Ψ_π	-0.625*	0.067	0.000	-0.700*	0.065	0.000
α	-0.019*	0.001	0.000	-0.020*	0.001	0.000

Table 13: Persistence and average inflation: with x_t, OECD countries, N=22

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.737*	0.037	0.000	0.633*	0.043	0.000
Φ_π^{11}	0.179	0.481	0.373	0.018	0.268	0.674
Ψ_0	0.233*	0.019	0.000	-0.234*	0.017	0.000
Ψ_π	-0.851*	0.110	0.000	-0.852*	0.103	0.000
α	-0.017*	0.001	0.000	-0.012*	0.001	0.000

Table 14: Persistence and average inflation: with x_t, non-OECD countries, N=29

	QTD data			H-P data		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.917*	0.029	0.000	0.798*	0.033	0.000
Φ_π^{11}	-0.603*	0.170	0.000	-0.343	0.183	0.061
Ψ_0	0.027*	0.008	0.001	0.012*	0.007	0.104
Ψ_π	-0.032*	0.010	0.002	-0.019*	0.010	0.056
α	-0.004*	0.001	0.011	-0.001	0.001	0.320

For the non-OECD countries, the coefficient of interest, $\Phi_\pi^{\mathbf{11}}$, is negative for QTD and H-P cases (see Table 14). For the QTD, it is statistically significant at the 5 per cent level and for the H-P case it is almost significant at the 5 per cent level. The three countries for which the episodes of hyperinflation have been identified are all non-OECD countries. The exclusion of these episodes renders the coefficients statistically insignificant at any reasonable level of significance (see Table 15). Table 16 reports the results when Argentina, Brazil, and Israel are excluded from the estimation. The coefficients are not only statistically insignificant but of the opposite sign from what is predicted by the theoretical model.[23]

6 The Two-Stage Estimation

I use a two-stage estimation procedure similar to the one employed extensively in previous research (for example, Ball et al. (1988)) that has focused on the influence of average inflation on the impact effects of aggregate demand shocks. In the first stage, I estimate country-

[23]For AR(2) dynamics for the non-OECD countries. The χ^2 test statistic value is 12.1 and the p-value is 0.002, indicating that the null hypothesis is rejected. However, when I exclude the episodes of hyperinflation, the χ^2 test statistic value is 1.88 and the p-value is 0.391. Therefore, the null hypothesis cannot be rejected. That is, within the non-OECD countries, average inflation does not influence second-order output dynamics. Finally, when I exclude Argentina, Brazil, and Israel, the χ^2 test statistic value is 0.29 and the p-value is 0.866. The results for the QTD are qualitatively similar to the H-P data and are not reported here.

Table 15: Persistence and average inflation: with x_t, non-OECD countries, excluding episodes of hyperinflation

	QTD data			**H-P data**		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.863*	0.033	0.000	0.749*	0.033	0.000
Φ_π^{11}	-0.192	0.402	0.528	0.054	0.258	0.834
Ψ_0	0.096*	0.013	0.000	0.075*	0.012	0.000
Ψ_π	-0.270*	0.004	0.002	-0.243*	0.039	0.000
α	-0.009*	0.001	0.000	-0.006*	0.001	0.000

Table 16: Persistence and average inflation: with x_t, non-OECD countries, excluding Argentina, Brazil, and Israel

	QTD data			**H-P data**		
Coefficient	Estimate	Std. error	p-value	Estimate	Std. error	p-value
Φ_0^1	0.842*	0.034	0.000	0.751*	0.040	0.000
Φ_π^{11}	0.019	0.936	0.528	0.109	0.286	0.701
Ψ_0	0.146*	0.015	0.000	0.112*	0.014	0.000
Ψ_π	-0.491*	0.055	0.002	-0.406*	0.049	0.000
α	-0.012*	0.001	0.000	-0.008*	0.001	0.000

specific measures of persistence in output fluctuations. In the second stage, I examine how average inflation affects the estimated measures of persistence in the cross-section of countries.

This approach is intuitively appealing for testing the cross-sectional implication of the theory. However, it assumes that countries are at different levels of steady inflation. Under this assumption, average inflation is the appropriate measure of the inflationary environment. As stated in Section 5, this assumption does not hold up in the data, particularly for countries that have experienced episodes of hyperinflation.

6.1 AR(1) specification

Stage 1 estimates an AR(1) model of the detrended output and identifies the coefficient on the lagged output deviation, β_i, as the measure of persistence in country i. Specifically, the estimated regression is

$$y_{it}{}^D = \beta_i y_{it-1}{}^D + \alpha x_{it} + u_{it}, \tag{21}$$

where $y_{it}{}^D$ is the detrended output and x_{it} is the nominal output growth of country i.

In Stage 2, the estimated regression is

$$\beta_i = \delta_0 + \delta_1 \bar{\pi}_i + \delta_2 \bar{\pi}_i{}^2 + \epsilon_i, \tag{22}$$

where $\bar{\pi}_i$ is the average inflation in country i. The results are reported in Table 17.[24] The

[24]Potentially, one can allow for the β_i to be time-varying. I estimated a simple state-space model for each country where $\beta_{it} = \beta_{it-1} + u_{it}$. I obtained a time series for the persistence parameters and used panel

Table 17: Persistence and average inflation: N=51

	QTD data			H-P data		
Variable	Estimate	Std. error	p-value	Estimate	Std. error	p-value
$\bar{\pi}_i$	-0.920*	0.252	0.001	-0.566*	0.272	0.043
δ_0	0.870*	0.035	0.000	0.750*	0.036	0.000

standard errors reported in the tables are heteroscedasticity consistent.

I find that the relationship between average inflation and persistence in the full sample is negative and significant for both QTD and H-P filtered data. This finding resembles the evidence reported in Kiley (2000). He finds that the degree of price stickiness varies inversely across countries. That is, high-inflation countries exhibit less-persistent output fluctuations. Kiley interprets these results as being strongly supportive of the sticky-price models. Nevertheless, I argue that there are compelling reasons to question whether this evidence is supportive of the sticky-price model of persistence in output fluctuations.

6.2 The inverted-U relationship

I use squared average inflation as a regressor to capture any non-linearities in the empirical relationship between persistence and average inflation. There is a statistically significant inverted-U relationship between persistence and average inflation in the data. This relationship is at variance with a clear theoretical prediction of the sticky-price models, in which the frequency of price adjustment depends on the average inflation rate. Table 18 shows that the critical average inflation rate is 27 per cent for H-P filtered data and 18 per cent for the QTD. Below these critical average inflation rates, the relationship between persistence and average inflation is *positive*. In the entire sample only 5 countries are above the 27 per cent

estimation. The results were qualitatively similar to the ones reported here.

Table 18: Persistence and average inflation: the inverted-U relationship, N=51

	QTD data			H-P data		
Variable	Estimate	Std. error	p-value	Estimate	Std. error	p-value
$\bar{\pi}_i$	0.723*	0.230	0.003	1.005*	0.251	0.000
π_i^2	-2.000*	0.225	0.000	-1.921*	0.205	0.000
δ_0	0.727*	0.031	0.000	0.617*	0.031	0.000

Table 19: Persistence and average inflation: excluding episodes of hyperinflation, N=51

	QTD data			H-P data		
Variable	Estimate	Std. error	p-value	Estimate	Std. error	p-value
$\bar{\pi}_i$	-0.765	0.412	0.070	-0.189	0.385	0.624
δ_0	0.831*	0.037	0.000	0.690*	0.036	0.000

average inflation and only 10 countries have average inflation rates above 18 per cent.[25]

6.3 Episodes of hyperinflation

Section 2 made a case for excluding periods of extreme time variation in inflation for certain countries. Table 19 provides the results for when episodes of hyperinflation are excluded. These results again are striking. None of the measures of persistence has a statistically significant relationship with average inflation at the 5 per cent level of significance. This implies that only episodes of hyperinflation are informative about the prediction of the theory.

Table 20 shows the results when Argentina, Brazil, and Israel are excluded from the estimation. The relationship between average inflation and persistence is *positive* and statistically significant in H-P filtered data (see Figures 4—6), and the inverted-U relationship

[25]Similar results hold when I restrict attention to the countries in the Ball et al. (1988) data set. The results are qualitatively similar when I use unanticipated nominal output growth to capture the impact effects of aggregate demand. I also controlled for supply shocks using the producer price index for crude oil and obtained similar results.

Table 20: Persistence and average inflation: excluding Argentina, Brazil, and Israel, N=51

	QTD data			**H-P data**		
Variable	Estimate	Std. error	p-value	Estimate	Std. error	p-value
$\bar{\pi}_i$	0.032	0.255	0.901	0.340*	0.160	0.039
δ_0	0.771*	0.032	0.000	0654*	0.027	0.000

Table 21: Persistence and average inflation: the inverted-U relationship, excluding episodes of hyperinflation, N=51

	QTD data			**H-P data**		
Variable	Estimate	Std. error	p-value	Estimate	Std. error	p-value
$\bar{\pi}_i$	0.079	0.940	0.933	-0.420	0.822	0.612
π_i^2	-2.066	1.719	0.235	0.564	1.145	0.701
δ_0	0.778*	0.063	0.000	0.705*	0.057	0.000

is no longer statistically significant (see Table 21). The coefficients δ_1 and δ_2 are jointly insignificant.

6.4 Alternative measures of persistence

Stage 1 estimates an AR(1) specification,

$$y_{it}{}^D = \beta_i y_{it-1}{}^D + u_{it}, \tag{23}$$

to obtain the cumulative response of output deviations to aggregate demand shocks. This measure of persistence is given as

$$\frac{1}{1-\beta_i}.$$

As above, the prediction of the theoretical models is that the cumulative response of output fluctuations in high-trend inflation countries should be smaller. An AR(2) model of the

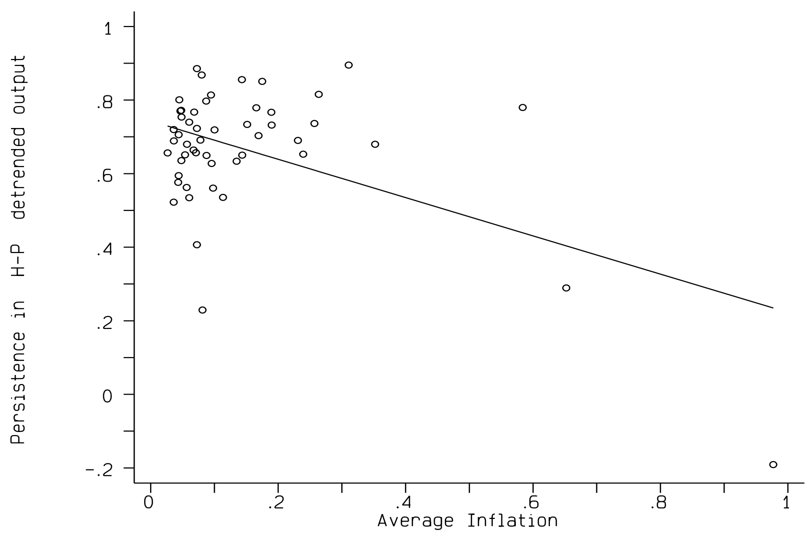

Figure 4: Full sample

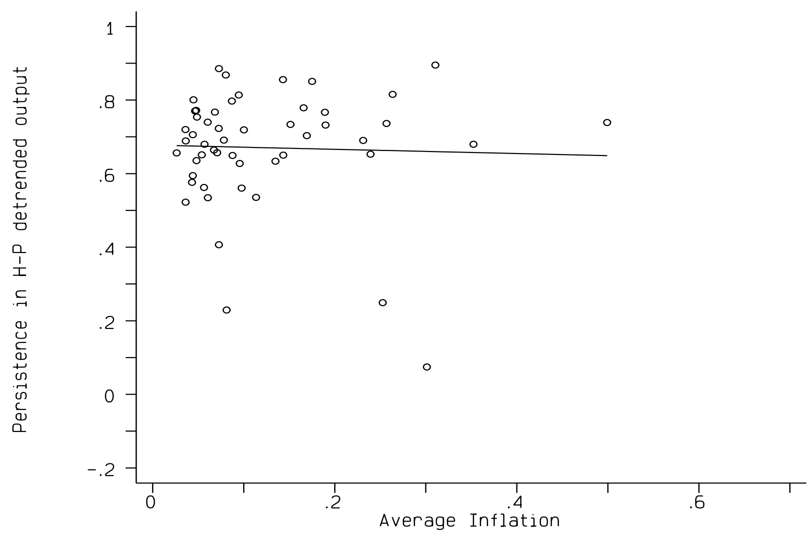

Figure 5: Excluding episodes of hyperinflation

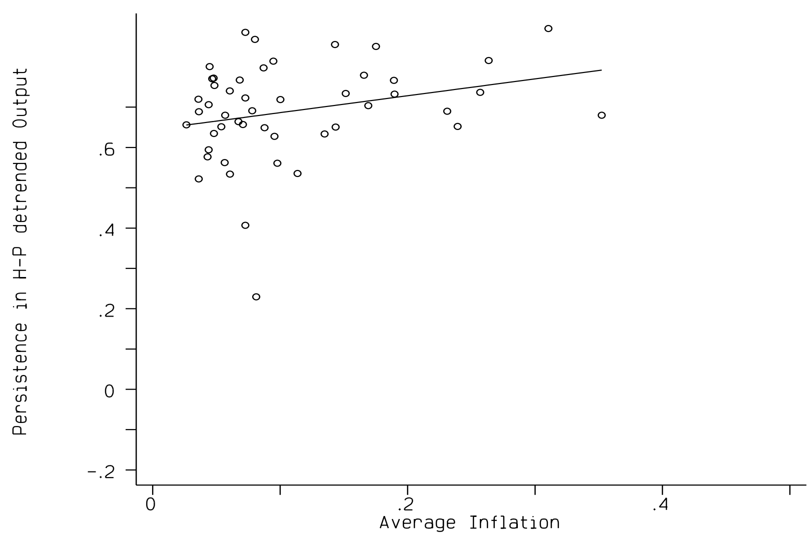

Figure 6: Excluding Argentina, Brazil, and Israel

detrended output is estimated to capture higher-order dynamics in the detrended output:

$$y_t^D = \beta_{i1} y_{it-1}^D + \beta_{i2} y_{it-2}^D + \epsilon_{it}.$$

The cumulative response of output to aggregate demand shocks is given as

$$\frac{1}{1 - \beta_{i1} - \beta_{i2}}.$$

Stage 2 uses these measures as dependent variables in (21) above. Table 22 gives the results for these alternative measures of persistence. The results indicate that no statistically significant relationship exists between average inflation and the cumulative response of output to aggregate demand shocks.[26]

6.5 Other country-specific characteristics

As in Kiley (2000), I allow for per-capita income differences PCY_{70} across countries, trade-openness of countries (OPEN), dummy variable for OECD status (OECDUM), dummy vari-

[26]I checked the relationship between the unconditional variance of the AR(1) and AR(2) specifications and average inflation rates:

$$V_{AR(1)}(y_t^D) = \frac{\sigma_\epsilon^2}{1 - \beta^2} \equiv B(\beta)\sigma_\epsilon^2, \qquad |\beta| < 1.$$

Here, $B(\beta)$ captures the extent to which the variability of output fluctuations is magnified by the persistence parameter. $B(\beta_1, \beta_2)$ is obtained from the unconditional variance of the AR(2) model, which is given as

$$V_{AR(2)}(y_t^D) = \frac{(1 - \beta_2)\sigma_u^2}{(1 + \beta_2)((1 - \beta_2)^2 - \beta_1^2)} \equiv B(\beta_1, \beta_2)\sigma_u^2.$$

The conditions for covariance stationarity,

$$|\beta_2| < 1, \qquad -2 < \beta_1 < 2, \qquad \beta_1 + \beta_2 < 1, \qquad \beta_2 - \beta_1 < 1.$$

are satisfied for all countries. None of these results are statistically significant. Similar results hold for the QTD.

Table 22: Alternative measures of persistence and average inflation: H-P data, N=51

Persistence measure	δ_1	δ_0
$\frac{1}{1-\beta}$	1.033 (8.69)	6.561* (1.39)
$\frac{1}{1-\beta_1-\beta_2}$	-1.758 (1.52)	3.768* (0.32)
S.E. in brackets. *= Sig. at 5% level.		

Table 23: With country-specific characteristics: dependent variable is β, QTD

Variable	Full sample	Excl. hyp. inf. epi.	Excl. A, B, I
$\bar{\pi}_i$	-0.671* (0.24)	-0.362 (0.44)	-0.207 (0.35)
$OPEN_{70}$	0.0006 (0.0003)	0.0008* (0.0004)	0.0006 (0.0003)
OECDUM	0.02 (0.07)	0.147 (0.11)	-0.003 (0.07)
LATINDUM	0.170* (0.06)	0.155* (0.06)	0.129* (0.05)
PCY_{70}	-0.56E-5 (0.1E-4)	-0.17E-4 (0.15E-4)	0.1E-5 (0.9E-5)
δ_0	0.775* (0.08)	0.722* (0.10)	0.702* (0.09)
S.E. in brackets. *= Sig. at 5% level.	N=51	N=51	N=48

able if a country is a Latin American country (LATINDUM), and persistence in the exogenous driving process, ρ_x, as measured by the autocorrelation in nominal output growth.

Consideration of these country-specific characteristics does not alter the conclusions (see Tables 23 and 24).[27]

[27]The variables PCY_{70} and OPEN $\equiv \frac{Imports+Exports}{GDP}$ were obtained from the Penn World Tables. Similar results hold for H-P filtered data.

Table 24: With country-specific characteristics: dependent variable is β, QTD

Variable	Full sample	Excl. A, B, I
$\bar{\pi}_i$	-0.717* (0.23)	-0.309 (0.37)
ρ_χ	0.192* (0.09)	0.154 (0.09)
$OPEN_{70}$	0.0005 (0.0003)	0.0006 (0.0003)
OECDUM	0.0001 (0.07)	-0.024 (0.06)
LATINDUM	0.162* (0.05)	0.123* (0.05)
PCY_{70}	-0.7E-5 (0.9E-5)	0.3E-6 (0.8E-5)
δ_0	0.704* (0.09)	0.655* (0.10)
S.E. in brackets. *= Sig. at 5% level.	N=51	N=48

7 Concluding Remarks

This paper investigates a key implication of the sticky-price model of business cycle that is designed to explain the observed persistence in output fluctuations around its long-run growth path. A prediction of this model is that a higher trend inflation rate will increase the frequency of price adjustment and thereby inversely affect the persistence in deviations of output from its trend. Therefore, countries with a high trend inflation rate should have less-persistent output fluctuations. The characteristics of the data set available to examine this hypothesis raise several implementation issues. This paper focuses on those issues in a manner that is consistent with the theoretical framework of the model. My estimation results, in general, do not support the hypothesis that higher inflation countries have less-persistent output fluctuations. When there is support for the hypothesis, it is extremely fragile. This support is found to be driven by a few episodes of hyperinflation. The results are in contrast with those in Kiley (2000).

The lack of support for the hypothesis raises a problem for the propagation mechanism based on sticky prices. If price stickiness is indeed the central mechanism for the transmission of monetary shocks to the real side of the economy in a persistent manner, then one must assume that the costs of nominal contracting are inelastic with respect to the inflationary environment over a very broad range of inflation, from 2—3 per cent to 60—70 per cent. However, such a position is intuitively unappealing. Therefore, the results of this paper suggest that further investigation of the sticky-price channel is required before it can provide a solid microfoundation for explaining persistence of output fluctuations.

Ball et al. (1988) focus on the prediction of the sticky-price model for the impact effect of aggregate demand shocks. They find that in high inflationary environments the impact effect of demand shocks is smaller. My results confirm their findings. The results reported here suggest that there is robust support for the prediction of the sticky-price model for one aspect of the business cycle: the impact effect of aggregate demand shocks. However, the same data does not support the prediction of the model for another key aspect of business cycles: the persistence of output fluctuations.

One direction is to examine the prediction of the sticky-price model that is used to explain persistent deviations of real exchange rates from purchasing-power parity (PPP). The sticky-price framework has been extensively studied in efforts to understand real exchange rate dynamics. According to the PPP theory, the real exchange rate (nominal exchange rate adjusted for price levels) should be constant. However, a stylized fact in international macroeconomics is that real exchange rates exhibit peristent deviations from PPP.[28] A leading explanation is based on sticky prices. In the presence of sticky prices, the national price

[28] See Froot and Rogoff (1995).

levels are either fixed or very slow to adjust. Therefore, aggregate demand shocks that cause a change in the nominal exchange rate make the real exchange rate deviate from PPP. A prediction of this model of real exchange rate fluctuations is that: *The deviations of real exchange rates from PPP should be less persistent in countries with high inflation.* The argument underlying this prediction is the same as the one described in the introduction. I test this prediction and find that the conclusions are similar to the ones reported in this paper (Khan (1999)). The support for the hypothesis is extremely fragile. When I exclude episodes of hyperinflation, the inverse relationship between persistence in real exchange rate fluctuations and inflation is statistically insignificant.

A second direction that appears to be promising is to further examine the interaction of price stickiness with elements of market incompleteness, distortionary taxation, strategic competition, and the role of non-monetary shocks within the framework of the dynamic general-equilibrium model.

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Appendix

I estimate the following specification for the 51 countries allowing for cross-sectional heteroscedasticity:

$$y_{it}^{d} = \alpha + \Phi_{0}^{1} y_{it-1}^{d} + \Phi_{\pi}^{11} f(\Pi_{it}) y_{it-1}^{d} + \Psi_{0}\mu_{it} + \Psi_{\pi} f(\Pi_{it})\mu_{it} + u_{it}. \tag{24}$$

Defining $[f(\Pi_{it}) * y_{it-1}^{d} \quad \mu_{it} \quad \pi_{it}\mu_{it}] \equiv X_{2,it}$, writing equation (24) as

$$\begin{bmatrix} \mathbf{y_1} \\ \vdots \\ \mathbf{y_{51}} \end{bmatrix} = \begin{bmatrix} 1 & \mathbf{y_1^l} & \mathbf{X_{2,1}} \\ \vdots & \vdots & \vdots \\ 1 & \mathbf{y_{51}^l} & \mathbf{X_{2,51}} \end{bmatrix} * \begin{bmatrix} \alpha \\ \Phi_0^1 \\ \Phi_\pi^{11} \\ \Psi_0 \\ \Psi_\pi \end{bmatrix} + \begin{bmatrix} \mathbf{u_1} \\ \vdots \\ \mathbf{u_{51}} \end{bmatrix}, \tag{25}$$

where $\mathbf{y_i} = [y_{i1}...y_{iT_i}]'$, $\mathbf{y_i^l} = [y_{i0}...y_{iT_{i-1}}]'$, and $\mathbf{u_i} = [u_{i1}...u_{iT_i}]'$. Equation (24) can be compactly written as

$$\mathbf{y} = \mathbf{X}\Phi + \mathbf{u}.$$

The covariance matrix is given as $E[\mathbf{uu}'] = \Omega$ where $\Omega = \Sigma_{\mathbf{NxN}} \otimes \mathbf{I_{T_i x T_i}}$. The off-diagonal elements of the Ω matrix are all zeros because of the unbalanced nature of the data. I allow for cross-sectional heteroscedasticity thus $Diag(\Sigma) = [\Sigma_1 ... \Sigma_{51}]'$.

I conduct a feasible generalized least-square estimation to obtain the parameter estimates:

$$\hat{\Phi} = (\mathbf{X}'\hat{\Omega}^{-\mathbf{1}}\mathbf{X})^{-\mathbf{1}}\mathbf{X}'\hat{\Omega}^{-\mathbf{1}}\mathbf{y}, \tag{26}$$

where $\hat{\Sigma_{ii}} = \frac{\hat{u_i}'\hat{u_i}}{T_i}$.

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