A note on the determinants of inflation starts in the OECD

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Abstract

Boschen and Weise (*Journal of Money, Credit and Banking*, 2003) model the probability of a large upturn in inflation. We extend their work to show that openness to trade exerts a negative effect on the probability of such an event.

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1 Introduction

In a recent study Boschen and Weise (*Journal of Money, Credit and Banking,* 2003) model the probability of a large upturn in inflation during a period of either stable or declining inflation, an occurrence that they term an 'inflation start'.[1] The results indicate that three factors tend to precipitate these sustained increases in inflation. First, high rates of real GDP growth increase the probability of an inflation start, the idea being that rapid growth reflects policy-makers' attempts to exploit the short-run Phillips curve, which must eventually lead to higher inflation. Second, the gap between inflation in the United States and domestic inflation raises the probability of an inflation start, because inflation shocks in the world's largest economy tend to be propagated internationally. Third, if a general election takes place in a particular year then the probability of an inflation start in that year is higher, the interpretation being that government policies aimed at 'buying' votes are often inflationary. In contrast, oil price hikes, fixed exchange

[1] Throughout the remainder of this note, Boschen and Weise (2003) will be referred to as BW.

rate regimes, fiscal policy variables and the political orientation of the government do not exert a robust effect on the probability of an inflation start, see BW for further details.

One determinant of the probability of an inflation start that BW do not examine is openness to international trade. The relationship between openness and the *level* of inflation has received a lot of attention in the literature. Romer (1993) and Lane (1997) show that greater trade openness decreases the time consistent inflation rate through reducing the net marginal benefit of surprise inflation, and both authors present empirical evidence suggesting that openness and inflation are negatively correlated across countries (Romer (1993) finds that the correlation is weak amongst a sub-sample of OECD countries, but Lane (1997) shows that it becomes much stronger after controlling for country size). On the supply-side, greater trade openness is likely to increase competition in product markets, such that firms with monopoly power are less able to push through inflationary price increases. Such a mechanism is emphasised in Aron and Muellbauer (2000), and illustrated using South African data.

In view of the link between the levels of openness and inflation that has been established in the literature, one might expect openness to trade to exert a negative effect on the conditional probability of an inflation start. The purpose of this note is to test that hypothesis using the methodology introduced by BW. The empirical results show that greater trade openness decreases the probability of an inflation start, even after one controls for the variables emphasised by BW. A comparison of different model specifications indicates that it is changes in openness over time, rather than cross-country differences in openness, that matter for the probability of an inflation start.

2 Methodology and data

In order to determine the timing of inflation starts, BW first calculate trend inflation as a centred nine quarter moving average of actual quarterly inflation. From this series trough (peak) dates in the inflation process are defined as dates at which trend inflation is lower (higher) than in the preceding and succeeding four quarters. An inflation episode is then defined as the period over which trend inflation rises by at least 2% from trough to peak and which is preceded by four or more quarters of stable or declining inflation. An inflation start is said to occur in the year following the year in which a trough occurred.

Using data from 19 OECD countries for the period $1961 - 93$, BW identify 73 inflation starts.[2] Binary variables set equal to one in years during which inflation starts occurred and to zero during years of stable or declining inflation are then created for each country, and these time series are stacked to produce a single variable. The years during which an inflation upturn is ongoing, namely the years after an inflation start and up to and including the next inflation peak, are excluded from the index, i.e. the observations for such years are treated as missing values, see BW for details.

The inflation starts indicator can be denoted Y_{it}, where i refers to a country and t to a year. A model for the probability of an inflation start that captures the core variables in the BW study can then be obtained by estimating the following probit regression:

$$Pr\ (Y_{it} = 1 \mid \bullet) = \beta_0 + \beta_1 GDP\ growth_{i\ t-1} + \beta_2 INFDUS_{i\ t-1} + \beta_3 ELECT_{i\ t} \tag{1}$$

where $\bullet$ summarises the information set, $GDP\ growth$ is the annual percentage change in real GDP, $INFDUS_i$ is the annual rate of consumer price inflation in the United States minus the annual rate of consumer price inflation in country i and $ELECT$ is a dummy variable set to unity when general elections take place and to zero otherwise.

The new results presented in this note introduce openness to trade as a further covariate in model (1). Openness is measured as the percentage share of imports of goods and services in nominal GDP, an approach that is standard in the literature, see Romer (1993). Some variations on this measure of openness are also considered. As noted above, BW also examined the role of oil price shocks, exchange rate regimes, fiscal policy and the political orientation of the government in determining the probability of an inflation start, but robust supporting evidence could not be found for any of them. Some of these variables, especially oil price shocks, are likely to affect the magnitude of inflation surges, but they do not seem to account for the timing of the *starts* of these inflation episodes. In view of this, we do not consider these variables in testing for a relationship between openness and the probability of an inflation start.

[2] The countries in the sample are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, the UK and the US.

The data on inflation starts were extracted from BW and cross checked with data supplied by Boschen. The data on elections were extracted from Alesina and Roubini (1997), the source used by BW.[3] Data on GDP volumes, consumer prices indices, nominal GDP and nominal import spending were taken from the *International Financial Statistics* database, which is also the source used by BW.

3 Empirical results

In Table 1 we present our empirical results. Regression (1) comprises the core BW variables. The marginal effects, calculated at the means of all variables, are positive, and the corresponding t-ratios are very close to those obtained by BW for comparable model specifications, see, for instance, Model 7.1 in their paper. The estimates that we obtain are generally smaller in absolute size than those reported by BW (though the differences only appear significant in the case of GDP growth). This could be due to differences in specification (the BW models contain further variables that are not robustly significant), or slight differences in the sample size (the basic sample used to fit the models in Table 1 consists of 377 observations, while the sample used to fit Model 7.1 in BW consists of 368 observations). The key point here is that our dataset yields results qualitatively similar to those in BW.

In regression (2) we add country dummies to regression (1), allowing for some idiosyncratic element in the probability of an inflation start. The results are robust to this extension of the model. In regression (3) we condition on the first lag of openness to trade and the country dummies. The openness variable is negatively signed and highly significant, providing strong support for the proposition that the likelihood of an inflation surge is smaller when an economy is more open to international trade (the sample is larger in this case due to the exclusion of $INFDUS$ and hence the addition of U.S. observations, but this is not crucial to the results). In regression (4) we replace openness with the Hodrick-Prescott trend in openness, which is calculated separately for each country using a smoothing parameter set to 400, the recommended value for annual data. The marginal effect of openness on the probability of an inflation start is

[3] Alesina and Roubini do not provide information on general elections in Spain. We obtained such information from the website www.auswaertiges-amt.de.

slightly smaller in this case, but it is still significant and supports our main hypothesis. It will be noted that openness enters once lagged in regressions (3) and (4), to ensure consistency with the BW approach. If we instead use contemporaneous values of openness in these two regressions the absolute t-ratios are 1.94 and 2.65 respectively, indicating that the basic correlation is not especially sensitive to the dynamic structure of the model. In model (5) we add back in the BW variables. Although the marginal effect of openness is smaller than in column (3), the effect is still highly significant. If GDP growth is excluded from model (5), the marginal effect of openness increases in absolute value to .014 (the absolute t-ratio is 3.49).

What of the quantitative significance of these estimates? A one standard error increase in openness reduces the probability of an inflation start by 4.9%.[4] The increases in the probability of an inflation start following one standard error increases in GDP growth and inflation relative to the US are 9.22% and 11.47% respectively.[5] These calculations suggest that although trade openness makes a non-negligible contribution to the likelihood of an inflation start, its effects are less important than those generated by the variables identified by BW. This may be because one of the channels through which openness affects the probability of an inflation start is a reduction in the propensity for policy-makers to exploit the Phillips curve trade-off, an effect that BW control for using the growth rate of real GDP. If the latter term is excluded from column (5), the reduction in the probability of an inflation start after a one standard error increase in openness is 7.46%, which is much closer to the size of the effects estimated for the core BW variables.

In model (6) we exclude the country dummies, which means that the size of the openness effect is determined by both the time series variation in the data and the cross-sectional variation (in models (2)-(5) the country dummies control for the cross-sectional variation). This causes the openness term to lose significance, indicating that while increases in openness to trade within OECD countries are associated with a smaller probability of an inflation start, it is not true

[4] This calculation is based on the results in column (5). As this specification contains fixed effects that send the cross-sectional means of the covariates to zero, the standard error of openness is calculated after controlling for differences in average openness across countries. To be precise, it is the residual standard error obtained after regressing openness on a set of country dummies.

[5] The standard errors used for these calculations control for cross-sectional differences in the means of the variables, but as the means differ relatively little across countries almost identical results follow when using unconditional standard errors.

to say that countries with higher average levels of openness experience fewer inflation starts. This is not surprising when one considers the construction of the inflation starts indicator. As the years during which an inflation episode is ongoing are excluded from the sample, countries that experience protracted inflation episodes can only experience a handful of inflation starts over a 30 year period, e.g. Spain is the most inflation prone country in the sample in the sense that it has the highest average rate of inflation, yet it experiences just 3 inflation starts between 1960 and 1995, the joint lowest number. Further, the cross-sectional distribution of inflation starts turns out to be compact (5 countries experience 3 starts, 12 experience 4 starts and 2 experience 5 starts). This means that cross-country differences in the frequency of inflation starts are unlikely to correlate with the level of openness, which varies by a factor of 5 across countries.

In view of these considerations, it makes more sense to analyse the impact of openness on the chances of an inflation start using time series information, as in models (2)-(5) in Table 1. It is worth noting that if model (6) is augmented with a *single* country dummy for Japan, the least open country in the sample, the absolute t-ratio for openness rises to 1.85, which is significant at the 7% level. This indicates that one need only introduce very limited controls for the cross-sectional variation in the data in order to 'revive' the openness effect. An alternative approach is to replace the level of openness with the annual percentage change in openness, since this cleans out cross-sectional variation. We report such a model in column (7), and in column (8) we report the same model with 5 outlying observations excluded from the sample.[6] The absolute t-ratios for the openness effects are significant at the 6% and 2% levels respectively, and the impact of a one standard error increase in the annual percentage change in openness is a 4.82% reduction in the probability of an inflation start (calculated using the column (8) estimates). These results show that even when country dummies are omitted from the model, some measure of trade openness affects the probability of an inflation start.

Insert Table 1 about here.

[6] The five observations are the two largest absolute readings for the percentage change in openness, the two largest absolute readings for $INFDUS$ and the largest absolute reading for GDP growth (the second largest reading is already outside the sample because it occurs when an inflation episode is ongoing).

We estimated three further specifications that are not reported in Table 1. First,we augmented model (5) with 18 linear time trends, one for each of the countries included in the sample. The purpose of this step was to control for any trends in openness and the frequency of inflation starts that are common to both variables, and which may be inducing spurious correlations. The absolute t-ratio on openness falls slightly to 2.35, but is still significant at the conventional level. Second, we added time dummies to the model in order to check that the statistical significance of openness is not dependent on some short time interval during which lots of countries experienced inflation starts, e.g. 13 countries experienced an inflation start in 1979. The time dummy for 1961 was omitted to avoid multicollinearity and those for $1974-76$, $1980-83$, and $1989-93$ were omitted because no inflation starts occurred during those years, meaning that numerical optimisation problems arise when the corresponding time dummies are included in the model. The absolute t-ratio on openness in this specification was 2.26, confirming the robustness of our main finding. Third, we took a logit estimate of model (5) in Table 1. This gave an absolute t-ratio for openness of 2.53, and thus supports the result established in Table 1.

4 Summary

This note hypothesised a negative link between trade openness and the probability of a large upturn in inflation. This could arise because high levels of openness reduce the incentive for policy-makers to pursue expansionary policies, or because strong foreign competition limits the ability of price-setting firms to push through price increases. A range of probit regressions fitted using OECD data showed empirical support for this conjecture. A comparison of different specifications indicated that the negative correlation between openness and the probability of an inflation start arises because of the time series variation in the OECD panel rather than because of the cross-sectional variation.

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Table 1

The determinants of the probability of an inflation start

Regression	1	2	3	4
Constant	-.304 (9.33)	-.212 (3.03)	-.096 (1.12)	-.128 (1.45)
Openness (-1)			-.016 (3.78)	
Trend openness (-1)				-.013 (2.71)
GDP growth (-1)	.033 (5.13)	.039 (5.46)		
INFDUS(-1)	.034 (4.84)	.043 (5.46)		
Elect	.124 (3.48)	.112 (3.36)		
Country dummies	No	Yes	Yes	Yes
Log likelihood	-143.94	-133.85	-181.13	-185.38
No. of observations (0/1)	(310/67)	(310/67)	(332/73)	(332/73)

Regression	5	6	7	8
Constant	.020 (0.19)	-.272 (5.24)	-.304 (9.30)	-.310 (9.30)
Openness (-1)	-.009 (2.65)	-.001 (0.77)		
Trend openness (-1)				
Change in openness (-1)			-.005 (1.90)	-.006 (2.54)
GDP growth (-1)	.034 (4.99)	.033 (4.99)	.036 (5.22)	.037 (5.24)
INFDUS(-1)	.042 (5.51)	.034 (4.88)	.035 (5.06)	.036 (5.11)
Elect	.108 (3.36)	.123 (3.46)	.122 (3.49)	.121 (3.44)
Country dummies	Yes	No	No	No
Log likelihood	-129.86	-143.65	-133.67	-131.81
No. of observations (0/1)	(310/67)	(310/67)	(300/64)	(296/63)

Pooled probit regressions for periods of stable or declining inflation, 18 OECD countries, 1961-93. Dependent variable is a dummy set to one when there is an inflation start. Table reports the marginal effect of a variable evaluated at the mean of all variables. Absolute t-statistics for coefficient estimates in parentheses.