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Trade Openness and Inflation Episodes in the OECD

Boschen and Weise (*Journal of Money, Credit, and Banking*, 2003) model the probability of a large upturn in inflation in the OECD (an inflation start). We extend their work to consider the impact of trade openness on the probability of such an event. The main finding is that increased openness reduces the probability of an inflation start, both directly, and indirectly through restricting the role of general elections in triggering inflation starts.

JEL codes: E31, F41
Keywords: inflation starts, trade openness.

In a recent study Boschen and Weise (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. Firstly, 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. Secondly, the gap between inflation in the U.S. and domestic inflation raises the probability of an inflation start, because inflation shocks in the world's largest economy tend to be propagated internationally. Thirdly, if a general election takes place in a particular year, the

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

We are grateful to John Boschen for supplying data on inflation starts, and to Chris Martin and John Muellbauer for helpful discussions. We also thank an anonymous referee for several suggestions that helped to improve the paper. The first author acknowledges financial support from a British Academy post-doctoral fellowship.

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Received May 10, 2004; and accepted in revised form February 14, 2005.

Journal of Money, Credit, and Banking, Vol. 38, No. 2 (March 2006)

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 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 finds that the correlation is weak amongst a subsample of OECD countries, but Lane 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 short paper 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 controlling 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. We then analyse the effects of interaction terms based on openness and the three variables emphasised by BW. The positive effect of general elections on the probability of an inflation start is found to be smaller under conditions of greater trade openness, possibly because the costs of inflation are larger in more open economies such that the political cycle in inflation is less pronounced. In contrast, the impact of openness on the marginal effects of lagged GDP growth and inflation spreads vis-a-vis the U.S. are positive but statistically insignificant.

In what follows Section 1 sets out our methodology and data, Section 2 presents the basic empirical results, some robustness tests and an analysis of interaction terms and Section 3 concludes.

1. 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. Using 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 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 index. 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:

$$\Pr(Y_{it} = 1 \mid \bullet) = \beta_0 + \beta_1 \text{GDP growth}_{it-1} + \beta_2 \text{INFDUS}_{it-1} + \beta_3 \text{ELECT}_{it}\,. \tag{1}$$

where $\bullet$ summarises the information set, GDP growth is the annual percentage change in real GDP, INFDUS is the annual rate of consumer price inflation in the U.S. minus the annual rate of consumer price inflation in country i and ELECT is a dummy variable set to unity in general election years and to zero otherwise.

The new results that we present are obtained by adding trade openness and its interaction with each of GDP growth_{it-1}, INFDUS_{it-1}, and ELECT_{it} to 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 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 did not obtain robust evidence in support of 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 the empirical analysis that follows.

The data on inflation starts were extracted from BW and cross checked with data supplied by John 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* (IFS) database, which is also the source used by BW.

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 U.S.

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

2. 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 of 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 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 the contemporaneous values of openness in these two regressions the absolute *t*-ratios are 1.94 and 2.65, respectively (results not reported), indicating that the basic correlation is not especially sensitive to the dynamic structure of the model. In Regression (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. It is interesting to note that if GDP growth is excluded from Model (5), the marginal effect of openness increases in absolute value to .014 (*t*-ratio $= -3.49$).

What of the quantitative significance of these estimates? A one standard deviation 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 deviation increases in GDP growth and US inflation relative to domestic inflation are 9.22%

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 deviation of openness is calculated after controlling for differences in average openness across countries. Specifically, it is the residual standard deviation obtained after regressing openness on a set of country dummies.

TABLE 1

The Determinants of the Probability of an Inflation Start

Regression	1	2	3	4	5	6	7	8
Constant	–.304 (9.33)	–.212 (3.03)	–.096 (1.12)	–.128 (1.45)	.020 (0.19)	–.272 (5.24)	–.304 (9.30)	–.310 (9.24)
Openness (–1)			–.016 (3.78)		–.009 (2.65)	–.001 (0.77)		
Trend openness (–1)				–.013 (2.71)			–.005 (1.90)	–.005 (2.01)
GDP growth (–1)	.033 (5.13)	.039 (5.46)			.034 (4.99)	.033 (4.99)	.036 (5.22)	.037 (5.21)
INFDUS (–1)	.034 (4.84)	.043 (5.46)			.042 (5.51)	.034 (4.88)	.035 (5.06)	.036 (5.09)
Elect	.124 (3.48)	.112 (3.36)			.108 (3.36)	.123 (3.46)	.122 (3.49)	.122 (3.42)
Country dummies	No	Yes	Yes	Yes	Yes	No	No	No
Log likelihood	–143.94	–133.85	–181.13	–185.38	–129.86	–143.65	–133.67	–133.08
No. of observations (0/1)	(310/67)	(310/67)	(332/73)	(332/73)	(310/67)	(310/67)	(300/64)	(295/64)

Notes: 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 are in parentheses.

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 once lagged 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 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 inflation rate, yet it experiences just three inflation starts between 1960 and 1995, the joint lowest number. Further, the cross-sectional distribution of inflation starts turns out to be compact (five countries experience three starts, twelve experience four starts and two experience five 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 five 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. Nevertheless, 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 needs to introduce only 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 the first difference transformation accounts for cross-sectional variation in the data. We report such a model in column (7), and in column (8) we report the same model with five outlying observations excluded from

5. The standard deviations 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 deviations.

the sample.[6] The absolute t-ratios for the openness effects are significant at the 6% and 5% levels, and the impact of a one standard deviation 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.

2.1 Robustness Tests

In Table 2 we report four regressions intended to help evaluate the robustness of our main finding to changes in the empirical procedure that has been employed. Firstly, we augment Model (5) in Table 1 with 18 linear time trends, one for each of the countries included in the sample. This version of the model controls for 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 5% level and the estimated marginal effect of openness actually increases in magnitude. Secondly, we add time dummies to the model in order to check that the statistical significance of openness is not dependent on a 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 (including these dummies induces numerical optimisation problems in the maximum likelihood estimation). The results are presented in column (2) of Table 2. The magnitude of the openness effect decreases slightly reflecting the fact

TABLE 2
EXTENSIONS OF THE CORE REGRESSION MODEL

Regression	1	2	3	4
Extension of column 5, Table 1	Include country time trends	Include time dummies	Estimate by logit	Openness is (X + M/GDP)
Constant	.167 (0.83)	–.050 (0.97)	.021 (0.22)	.074 (0.68)
Openness (–1)	–.015 (2.35)	–.004 (2.26)	–.008 (2.53)	–.005 (3.07)
GDP growth (–1)	.028 (3.91)	.014 (3.99)	.031 (4.88)	.031 (4.73)
INFDUS (–1)	.044 (5.57)	.022 (5.40)	.038 (5.31)	.043 (5.65)
Elect	.096 (3.41)	.036 (2.45)	.098 (3.40)	.109 (3.43)
Country dummies	Yes	Yes	Yes	Yes
Log likelihood	–118.84	–74.44	–130.47	–128.44
No. of observations (0/1)	(310/67)	(310/67)	(310/67)	(310/67)

NOTES: See notes to Table 1. Each regression entails one extension of the specification in Table 1, column 5.

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).

that the time dummies now control for global shifts in the data, but the absolute t-ratio for the openness coefficient is 2.26, confirming the robustness of our main finding.

Thirdly, in column (3) of Table 2 we report a logit estimate of the core model. The absolute t-ratio for openness is 2.53, which is in line with the corresponding probit estimate. Fourthly, we obtained an alternative measure of openness that is frequently used in the literature, namely the sum of nominal export revenues and nominal import expenditures, relative to nominal GDP (recall that to this point openness has been measured using only information on import penetration).[7] The sample correlation between the two measures of openness exceeds 98% and this is reflected in the results in column (4) of Table 2, which show that the absolute t-ratio for openness rises to more than three (the marginal effect of openness is approximately half that in column (5) of Table 1 because the range of the new openness measure is approximately double that of the previous one).

2.2 Interacting Openness and the BW Covariates

In this sub-section we investigate a further dimension of the relationship between trade openness and the probability of an inflation start, namely the impact of openness on the marginal effect of the three covariates emphasised by Boschen and Weise (lagged GDP growth, general elections and the lagged inflation spread vis-a-vis the U.S.). Specifically, we interact the openness measure (based on import penetration) with each of the other explanatory terms. In the interactions the measure of openness has a zero mean across the panel, i.e. a constant has been subtracted from the original openness variable. This ensures that the reported marginal effect for a particular covariate is the marginal effect for a country characterised by the sample average value of openness. The marginal effect estimated for an interaction term has the usual interpretation, namely the change in the effect of the covariate for a one percentage point increase in openness.[8]

The regression in column (1) of Table 3 suggests that the marginal effect of lagged GDP growth on the probability of an inflation start is *higher* in more open economies. Thus, while greater openness causes governments and central banks to pursue policies that make an inflation start less likely, it is also the case that once an acceleration of GDP has occurred, a relatively open economy is more likely to experience an inflation start. This finding is consistent with the evidence presented by Karras (1999), which shows that in the aftermath of a monetary shock the increase in inflation for each 1% increase in GDP is larger the more open the economy. One explanation for this phenomenon is that expansionary monetary policy both raises GDP and depreciates the exchange rate, and the latter adjustment pushes up import prices and inflation in proportion to the openness of the economy (Romer,

7. The data were downloaded from the online Penn World Tables (version 6.1). The data for Germany commence in 1991. Pre-1991 observations for Germany were obtained by splicing the Penn series to a series that we constructed using data on current price measures of exports, imports, and GDP obtained from the IFS database cited in Section 1.

8. It should be emphasised that demeaning openness does not affect the t-ratios calculated for any of the covariates.

TABLE 3

Regressions Incorporating Interaction Terms

Regression	1	2	3	4	5
Constant	.087 (0.75)	–.030 (0.77)	.020 (0.19)	.026 (0.24)	.041 (0.35)
Openness (–1)	–.011 (2.90)	–.007 (2.10)	–.009 (2.65)	–.009 (2.69)	–.010 (2.37)
GDP growth (–1)	.034 (4.96)	.033 (4.94)	.034 (4.99)	.033 (5.02)	.033 (4.96)
INFDUS (–1)	.041 (5.49)	.042 (5.65)	.042 (5.51)	.042 (5.54)	.041 (5.68)
Elect	.107 (3.35)	.108 (3.41)	.109 (3.37)	.107 (3.33)	.102 (3.27)
GDP growth (–1) • Open	.0007 (1.34)				.0006 (1.17)
Elect • Open		–.007 (2.59)			–.007 (2.38)
INFDUS (–1) • Open			–.00008 (0.14)		
INFDUS (–1) • Open • BW				.0005 (0.64)	.0005 (0.77)
Country dummies	Yes	Yes	Yes	Yes	Yes
Log likelihood	–128.9	–126.22	–129.85	–129.64	–125.31
No. of observations (0/1)	(310/67)	(310/67)	(310/67)	(310/67)	(310/67)

Notes: See notes to Table 1. All terms in openness have been demeaned.

1993 provides a formal exposition of this idea). However, we do not emphasise the results in column (1) given that the marginal effect of the interaction term is insignificant at the 5% level.[9]

In column (2) we find that greater openness reduces the marginal effect of a general election on the probability of an inflation start. This effect is significant at the 1% level and is robust to the inclusion of country specific time trends (results not reported). One interpretation of the negative interaction term is that governments in relatively open economies are less inclined to launch inflationary economic expansions during election years because the costs of inflation can be larger in more open economies, e.g. inflation may cause real exchange rate fluctuations, which are likely to be more costly in open economies, see Temple (2002) for further discussion along these lines.

An interesting feature of Regression (2) is that in the most open country in the sample, Belgium, the impact of an election on the probability of an inflation start is effectively zero, e.g. average demeaned openness for Belgium is 25%, which implies that the full marginal effect of an election in Belgium is actually slightly negative, though very close to zero. On the other hand, in the most closed country in the sample, Japan, the effect of a general election is to increase the probability of an inflation start by more than 20%. In their paper, BW present a decomposition of the factors driving the predicted probability of an inflation start in the U.S., Germany, and Japan. The results indicate that elections were particularly important in starting inflation episodes in the U.S. and Japan in the early 1970s, and BW note that this conclusion is consistent with the evidence from analytical narratives on monetary policy. The regression that we present in column (2) suggests that general elections may be especially pertinent in understanding inflation starts in the U.S.

9. It is possible that fixed exchange rate regimes undermine the mechanism linking the marginal effect of GDP growth to the extent of trade openness, but our attempts to control for fixed exchange rate regimes did not change the results.

and Japan because a lack of openness in these two countries eliminates some of the disincentives to launching inflation starts in election years. This is one channel through which cross-country differences in openness matter for the likelihood of an inflation start (recall that the direct effect of openness on the probability of an inflation start arises because of the time series rather than the cross-sectional variation in openness).

In column (3) we test the hypothesis that more open economies are more likely to import inflation from the U.S. and therefore face a larger probability of an inflation start following a surge in U.S. inflation. This may happen because open economies tend to fix their exchange rate against the U.S.$ (see Edwards 1996), and a fixed exchange rate increases the marginal effect of the inflation spread on the probability of an inflation start (a result for which BW find some support). Our results do not support this hypothesis, however, since the sign of the interaction term is negative rather than positive. The reason for this may be that the positive impact of openness on the marginal effect of the inflation spread emerged only after the end of the Bretton Woods system in 1973. Prior to this date all of the countries in our sample adopted a fixed exchange rate against the U.S.$ as part of the post-war macroeconomic policy consensus, irrespective of their openness. Only during the post-1973 period have countries started to reveal a preference over the exchange rate regime that might be related to openness. In order to test this idea we define a variable that is equal to zero for each country for 1961–72, but then takes the value of the interaction between the inflation differential and openness for 1973–93 (openness is demeaned for the period 1973–93). In column (4) this term is positively signed, as we expect, but far from significant, and the picture is essentially the same in column (5), which controls for each of the interaction terms plus the covariates from the core model.

3. SUMMARY

This paper 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 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. In contrast, both the time series and the cross-sectional variation in openness serve to restrict the marginal effect of elections on the probability of an inflation start, an effect that may be due to higher costs of inflation in open economies deterring governments from launching inflation starts during election years.

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