Openness and Inflation Volatility: Panel Data Evidence*

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Abstract

Trade openness can affect inflation volatility via the incentives faced by policy-makers or the structure of production and consumption, but the sign of this effect, as predicted from economic theory, is ambiguous. This paper provides evidence for a negative effect of openness on inflation volatility using a dynamic panel model that controls for the endogeneity of openness and the effects of both average inflation and the exchange rate regime. Our results offer one explanation for the recent decline in inflation volatility observed in many countries. The relationship is shown to be strongest amongst developing and emerging market economies, and we argue that the mechanisms linking openness and inflation volatility are likely to be strongest amongst this group of countries.

Keywords: Openness, inflation, volatility, globalization.

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1 Introduction

A striking feature of recent global macroeconomic performance has been the substantial decline in inflation volatility. In the United States inflation volatility has fallen by two thirds since the mid-1980s and similar trends have been observed in other OECD countries (Blanchard and Simon 2001). Even developing countries, which continue to experience higher and more volatile inflation than the industrial countries, have seen inflation volatility fall since the early 1990s. Understanding such trends is important because existing research suggests that volatile inflation undermines other aspects of macroeconomic performance, for example Elder (2004) and Byrne and Davis (2004) report negative effects of inflation volatility on investment and growth respectively. In this paper we consider two channels through which openness to international trade may impact upon inflation volatility: (i) openness affects the costs associated with inflation volatility, creating an incentive for governments and central banks to implement policies that imply a different level of volatility; (ii) openness affects the structure of consumption and production, either increasing or decreasing the amount of diversification and hence the scope for individual price shocks to cancel out in the aggregate price index. In section 2 we review these mechanisms, drawing on papers by Cavelaars (2009) and Badinger (2009). The main message from the theory is that the effect of openness on inflation volatility is ambiguous. Our contribution is to then undertake a detailed empirical investigation of the relationship between openness and inflation volatility. We provide robust evidence for a negative effect of openness on inflation volatility using a dynamic panel model fitted using a large cross-country dataset. The effect is strongest amongst emerging market and developing economies, and we relate this finding to the theoretical mechanisms that may link openness and volatility.

The link between openness and inflation volatility has been explored by only a small number of papers. The relationship is briefly considered by Bleaney and Fielding (2002), Gruben and McLeod (2004) and Aisen and Veiga (2008a). A more detailed analysis of the openness effect is provided by Lo, Wong and Granato (2007), who present cross-country regressions indicating a negative correlation between openness and inflation volatility. In related literature, Bowdler and Nunziata (2006) show that greater openness reduces the probability of an upturn in inflation in OECD countries, consistent with reduced volatility, while Aisen and Veiga (2008a) present panel data evidence indicating that higher degrees of political instability and social polarization, less democracy, and lower *de facto* central bank independence are associated with more volatile inflation rates.[1] In this paper we implement three important extensions of the literature on the openness-inflation volatility relationship. Firstly, we control for the possible endogeneity of openness in this context. Secondly, we focus on the temporal relationship between openness and inflation volatility, which has not been evaluated previously.[2] This aspect of the relationship is important in understanding the reasons for the recent moderation of inflation fluctuations. Thirdly, we investigate 'common cause' explanations for the relationship, for example institutional structures promoting both trade openness and stable inflation, which have not been examined in past work. In this paper we address each of these issues through instrumenting a dynamic panel model using data spanning 96 countries and more than four decades. Our results indicate a negative effect of openness on inflation volatility that is robust to a range of controls that capture many potential common cause interpretations of the openness-inflation volatility relationship.

The evidence that we present parallels the negative relationship between openness and average inflation described by, *inter alia*, Romer (1993) and Chen, Imbs and Scott (2009).[3] However, we show that openness has a negative and statistically significant effect on inflation volatility even after controlling for mean inflation. Furthermore, our results are not affected by cross-country differences in the exchange rate regime, the use of inflation targets, data quality, or participation in IMF structural adjustment programmes. The relationship is, however, much stronger amongst developing and emerging market countries than amongst OECD countries, and we argue that this is because the channels linking openness and inflation volatility are more likely to apply amongst low income countries.

The remainder of the paper expands on these points and is organised as follows. Section 2 describes the channels through which openness may affect inflation volatility. Section 3 sets out

[1] In a separate and related literature, Krause and Mendez (2008) consider determinants of the inflation aversion parameter in a policy-maker loss function, but do not explore the impact of openness to trade in this context.

[2] The paper by Gruben and McLeod offers very brief evidence on the temporal link between openness and inflation volatility, but does not analyse the relationship in any detail and instead focuses mainly on the impact of openness on the level of inflation.

[3] The Romer evidence has been challenged, see for example Terra (1998), Bleaney (1999) and Alfaro (2005).

the econometric approach and describes the data used in testing for a negative effect of openness on inflation volatility. Section 4 presents the basic empirical results and various robustness tests, and Section 5 concludes.

2 Openness and inflation volatility

A useful starting point in thinking about the determinants of inflation volatility is the conduct of macroeconomic policy. A regime in which there are regular changes in the stance of fiscal and monetary policy, for example due to political business cycle effects, will likely be characterised by relatively volatile inflation. The conduct of policy will depend on policy-makers' attitudes to deviations of inflation from target. One possibility is that openness to trade increases the perceived costs of inflation fluctuations, creating an incentive for disciplined monetary policy. In relatively open economies, if home producers of tradables are forced to vary prices frequently, their market share will quickly be eroded because domestic and foreign consumers can substitute towards foreign produced goods for which the price quoted is more stable. Policy-makers in open economies can avoid this adverse outcome through adopting more disciplined macroeconomic policies, leading to lower inflation volatility. In a recent contribution, Cavelaars (2009) shows that the effect of increased openness on policy-maker incentives is more complicated than this example suggests. When openness is driven by reductions in technologically determined iceberg trade costs, greater policy discipline can arise. In contrast, rises in openness driven by decreases in import tariffs can undermine policy discipline. The reason is that declines in revenues from import tariffs (from the expenditure switching effect of currency depreciations) serve as a deterrent to monetary policy expansions. If greater openness is associated with the elimination of tariffs, this deterrent to discretionary policy interventions is weakened, such that the conduct of monetary policy is less disciplined and inflation volatility rises.[4] A different mechanism through which openness to trade may influence policy-maker incentives and inflation volatility outcomes is the slope of the output-inflation trade-off. Romer (1993) famously demonstrated

[4] Cavelaars also shows that when increasing openness to trade is characterized by greater product market competition, the expenditure switching effect of exchange rate adjustment is amplified, and that this causes welfare outcomes to shift such that policy-making is less disciplined.

that greater openness can reduce the trade-off (steepen the short-run Phillips curve) such that the incentive for surprise policy expansions is mitigated. As noted by Badinger (2009), recent theoretical contributions demonstrate that this effect can operate in the opposite direction in micro-founded models. The key idea is that openness negates incentives for price adjustment, inducing aggregate price stickiness and a larger output-inflation trade-off, see Razin and Yuen (2002) and Daniels and VanHoose (2006). Empirical evidence on the relationship between openness and the output-inflation is mixed. Badinger (2009) and Daniels, Nourzad and VanHoose (2005) find that openness raises the trade-off (pointing to less policy discipline) while Bowdler (2009) reports evidence of a negative openness effect on the trade-off, consistent with the original Romer hypothesis.

Given that economic theory suggests that the impact of openness to trade on policy discipline is ambiguous, we turn to the empirical evidence. Aisen and Veiga (2008b) show that a *de jure* measure of openness reduces the level of seigniorage collected by governments. In our own investigations we have extended the latter results to consider the volatility of seigniorage rates in 96 countries over the period $1961-2000$.[5] A panel regression in which the dependent variable is the volatility of seigniorage and the controls are the average growth rate of seigniorage and the log share of imports and exports in GDP (plus fixed effects and time dummies) yields an openness elasticity of $-.04$ (*t-ratio* $= 2.81$).[6] This evidence is consistent with the notion that monetary conditions are more stable in open economies, even though theoretical arguments point to multiple and conflicting effects of trade openness on policy discipline. To the extent that more stable monetary conditions engender inflation stability, openness to trade will exert a negative effect on inflation volatility.

As well as affecting the conduct of macroeconomic policy, openness to trade may influence the amount of volatility to which countries are exposed. A standard result in trade theory is that openness promotes diversity in consumption, as goods and services that are unavailable domestically can be imported, and specialisation in production, as countries exploit comparative

[5] The volatility of the seigniorage rate is defined as the natural log of one plus the decimal standard deviation of the annual growth rate in reserve money.

[6] This OLS estimate of the impact of openness on the volatility of reserve money growth is robust to GMM estimation and to the inclusion of additional controls such as per capita income.

advantage. A more diverse consumption basket will reduce aggregate consumer price volatility if shocks to individual markets are not perfectly correlated, because these shocks will in part cancel when aggregated. Such an effect will be more pronounced if international trade shifts consumption towards high value added goods, whose prices tend to be more stable due to greater flexibility in their supply. Trade may play a particularly important role in shifting consumption towards high value added goods in the case of developing countries, in which some manufactured goods are not produced domestically.

An example of the consumption diversification effect occurred in South Africa following the removal of tariffs and other trade barriers at the start of the 1990s. This caused the share of exports plus imports in GDP to rise from 45% during the period $1991-95$ to 53.2% during the period $1996-2000$. Furthermore, most of this increase in trade was driven by increased imports of the high value added manufactured goods that South Africa did not produce domestically. Aron and Muellbauer (2000) develop an econometric model for the import share of manufactures and find that the underlying trend in this quantity rose by 40% between 1991 and 1998. Against this background of increasing openness and a restructuring of consumption towards manufactured goods, the standard deviation of consumer price inflation fell from 3.04% for the period $1991-95$ to 2.16% for the period $1996-2000$. These trends are consistent with the idea that openness reduces inflation volatility through promoting a more balanced pattern of consumption.

On the other hand there are several channels through which openness may exert a positive effect on inflation volatility. As noted above, international trade can induce specialisation in production, increasing vulnerability to sector specific shocks. Such volatility in export earnings could influence the volatility of domestic demand and, through that channel, inflation volatility. A similar effect could arise if trade openness is associated with increased financial openness. Aghion, Bacchetta and Banerjee (2004) develop a model in which capital flows behave pro-cyclically and therefore increase output volatility.[7] These factors may offset part of any negative effect of openness on inflation volatility that arises through greater policy discipline and diversity in consumption. In order to address this possibility in our empirical work, we control for output

[7] A counter-argument is presented by Cavallo (2009) who estimates a negative effect of openness on output volatility amongst a cross-section of countries and attributes this finding to the fact that open economies offset external shocks through capital account transactions.

volatility, export concentration and capital account openness, variables that capture some of the mechanisms through which trade openness may exert positive rather than negative effects on inflation volatility. Our main results are not sensitive to the use of these controls.

3 Data and methodology

The inflation data that are used in this paper measure the annual rate of consumer price inflation at the quarterly frequency and are taken from *International Financial Statistics* (the appendix provides comprehensive notes on data sources). We compile data for an unbalanced panel of 96 countries covering the period 1961 : 1 to 2000 : 4.[8] It is important to note that as we have defined inflation as growth in the price index over the last year rather than the last quarter, there cannot be seasonal effects in the data that may induce spurious volatility.

The choice of our sample period, especially the decision to restrict it to the year 2000, is driven by important monetary policy developments in the European Union post-1999. As is well-known, member countries of the European Monetary Union (EMU) witnessed a harmonization of monetary policy after 1999, which effectively meant these countries lacked an independent monetary policy. Since these European states form the bulk of our OECD sample and monetary policy discipline is also one of the proposed mechanisms through which trade openness might affect inflation volatility, we terminate our sample in 2000. However, to ensure the robustness of our findings, we later re-estimate our core specifications on a sample that extends to the year 2015 (see section 4.3).

In order to measure inflation volatility we divide the data for each country into a maximum of 8 windows, each of 20 quarters (1961 : 1 to 1965 : 4, 1966 : 1 to 1970 : 4 and so on).[9] For each sub-period we then compute inflation volatility ($VINF$) as

$$VINF = \ln[1 + sd(INF)] \tag{1}$$

[8]The appendix lists the 96 countries included in the sample.

[9]The full set of time observations are typically only available for OECD countries and the larger emerging market economies. The maximum number of time observations per country is 8, the minimum 3 and (in the largest sample) the average is 5.6.

where sd is a standard deviation and INF is the decimal inflation rate (3% inflation is represented as 0.03).[10] A standard practice in the literature is to take log transforms to downweight very large readings that may occur during hyperinflation episodes. One disadvantage of the log transform is that it overweights observations very close to zero. To avoid this problem we consider the log of one plus the decimal standard deviation of inflation.[11]

In order for the standard deviation to be a valid measure of volatility, the mean of the data must not change during an observation window. If the mean were to change during an observation window, calculating the standard deviation around a unique mean would be invalid. An even greater challenge would arise if inflation trended over time, since then the mean inflation rate changes each period and a standard deviation calculated for a constant mean could be dominated by spurious components. The requirement for constant mean inflation **w**ithin 5 year windows is closely related to assumptions on the order of integration of the data. If quarterly inflation is a unit root, or $I(1)$, process then the mean inflation rate again changes each quarter such that the standard deviation of inflation is undefined and there can be no meaningful investigation of the economic determinants of inflation volatility. Evidence on the time series properties of inflation provides no clear indication as to whether inflation should be treated as stationary or non-stationary, with test outcomes dependent on the testing procedure and sample period used, see for instance Culver and Papell (1998).[12] Our approach is to make the assumption that inflation is at least *locally stationary*, so that over the 5 year observation windows the mean of inflation can be treated as constant, or at least constant in most cases

[10] A GARCH measure of volatility is not used for several reasons. Firstly, the time series available for some countries are too short to justify GARCH estimation. Secondly, fitting a GARCH model for 1961 : 1 to 2000 : 4 and using the parameters of that model to infer inflation volatility in each 5 year window implies that volatility in, say, the 1960s depends on parameters estimated using data from later decades such as the 1980s. Clearly, policy-makers in the 1960s could not have targeted this measure of volatility, which implies that it is not the right statistic for assessing the relationship between openness and inflation volatility. Thirdly, if 5 years is taken as the window length, high frequency information on the conditional variance taken from a GARCH model would be lost, as quarterly observations would be averaged to form 5 year observations.

[11] Our results are robust to using two alternative definitions of inflation volatility, namely $\ln sd(100 * \text{inf})$ and $sd \ln(1 + \text{inf})$. Full details are available on request.

[12] Hendry (2001) argues strongly that the inflation rate should be treated as an I(0) process rather than an I(1) process.

so that any distortions to our volatility measures are limited. In other words, by measuring volatility at the 5 year frequency rather than over several decades, we reduce the probability of identifying spurious volatility associated with shifts in mean inflation, or the presence of trended inflation. There are some disadvantages to this approach. Firstly, it could be that inflation is stationary over 5 year windows but that these windows run from $1963-67$, $1968-72$ and so on, not for the intervals we consider starting with $1961-65$. Secondly, short window lengths risk identifying persistent shocks around a stable mean as examples of breaks in the mean, e.g. if a shock occurs at the start of a window and persists for most of the next 5 years. In such a case volatility would be under-stated by the standard deviation within each window. In order to get some insight into the sensitivity of our results to both the length and starting points of observation windows, we consider window lengths of 3 and 8 years in our robustness section. Our interpretation of the results from these exercises is that the estimated effect of trade openness on inflation volatility is not the result of spurious movements in volatility arising from the assumption of stable mean inflation within 5 year observation windows.[13]

The data for inflation volatility include some outliers, even after the transformation in (1). In order to ensure that our results are not driven by outliers we exclude observations more than three standard deviations from the mean of the unconditional distribution. This leads to 12 observations, approximately 1.5% of the sample, being dropped. These observations are mainly for Latin American countries that experienced extreme inflation during the 1980s.

Openness is defined as the natural log of the average value of imports plus exports relative to GDP in each 5 year period, and is denoted $OPEN$.[14] The log trade to GDP ratio is a frequently used proxy for openness and can arguably account for the core mechanisms linking

[13] A different approach to selecting observation windows would entail first testing for breaks in mean inflation, and then constructing observation windows of variable lengths that are consistent with the break dates. In the present context, one disadvantage of this approach would be that as the country time series are quite short in some cases, for example just 60 quarterly observations, the size and power properties of tests for structural breaks do not allow for reliable inference. Bilke (2005), for example, concludes that samples of several hundred are required in order to have confidence in the test outcomes (these sample sizes refer to monthly observations and so should be divided by 3 to ensure comparability with the quarterly data samples used in the present study). In view of such practical difficulties, we proceed using fixed windows to calculate volatility statistics.

[14] We checked $OPEN$ for outliers using the criterion applied to $VINF$, but none were found.

openness and inflation volatility emphasised in our earlier discussion, namely the extent of foreign competition for domestically produced goods and the ability to achieve diversity in consumption.[15] In the cases of both $VINF$ and $OPEN$, panel autoregressions of order one yield autoregressive coefficients that are significantly less than unity. These outcomes suggest that both series are stationary, or integrated of order zero, $I(0)$.

In Figure 1 we plot $VINF$ against $OPEN$. Even before controlling for country fixed effects, time dummies, other regressors and potential reverse causation, a negative relationship between openness and inflation volatility can be observed. Each graph reveals some extreme observations, even after the steps taken to deal with outliers. However, in the robustness section we show that our main results do not depend on these observations.

3.1 The econometric model

In order to estimate the effect of openness on inflation volatility we consider the following model:

$$VINF_{it} = \alpha + \beta VINF_{it-1} + \gamma OPEN_{it} + \eta_i + \lambda_t + \varepsilon_{it} \tag{2}$$

where i denotes a country, t a 5 year period, η_i a country fixed effect and ε_{it} the error term. The lagged dependent variable in (2) controls for persistence in inflation volatility, which may be intrinsic or simply a proxy for other determinants of volatility that are omitted at this stage.

The approach to estimating (2) follows Arellano and Bond (1991) and Arellano and Bover (1995). In order to eliminate the time invariant fixed effects we take first differences of (2) to obtain

$$\Delta VINF_{it} = \beta \Delta VINF_{it-1} + \gamma \Delta OPEN_{it} + \lambda_t + \Delta\varepsilon_{it} \tag{3}$$

Estimating (3) by least squares is problematic. Firstly, the transformed error term is correlated with the lagged dependent variable (both include ε_{it-1}) and this will lead to biased parameter estimates. Secondly, a *de facto* measure of openness based on the trade share is endogenous to inflation volatility, e.g. in the previous section we argued that volatile inflation

[15] The use of the log transform is not crucial to our results: If we use the untransformed trade share in our regressions the main implications of our empirical analysis are unchanged.

could erode the competitiveness of exporters. Alternatively, there may exist a common cause for openness and inflation volatility, for example both may be a consequence of deeper preferences over macroeconomic outcomes, or both may be subject to terms of trade shocks (a collapse in export prices may reduce the nominal value of trade and at the same time reduce aggregate demand so that inflation volatility rises). In order to test the hypothesis that increased openness leads to lower inflation volatility it is necessary to identify exogenous movements in trade that are not plausibly related to the current level of inflation volatility.

Arellano and Bond (1991) suggest a generalised method of moments (GMM) technique for estimating (3). Assuming that the errors in equation (2) are serially uncorrelated and that the explanatory variables are uncorrelated with future realisations of the errors (their endogeneity implies that they are correlated with only current values of the errors) lags of $VINF$ and $OPEN$ dated $t-2$ and earlier are valid instruments with which to identify the exogenous variation in openness.[16] A potential drawback of this *Differenced-GMM* estimator is that in the presence of high time-series persistence and short panels, lagged *levels* of the variables may be poor instruments for subsequent first differences, leading to finite sample biases (Blundell and Bond 1998). An alternative approach, suggested by Arellano and Bover (1995) and Blundell and Bond (1998), is the *System-GMM* estimator, which uses lagged differences of each variable as instruments in estimating the levels relationship in (2), and combines this information with the *Differenced-GMM* estimates of equation (3). The validity of these instruments requires a constant correlation between $VINF_{it}$ and the fixed effect, and between $OPEN_{it}$ and the fixed effect. If this is the case, $\Delta VINF_{it-1}$ and $\Delta OPEN_{it-1}$ are orthogonal to future realisations of the error terms and represent valid instruments for estimating the parameters of (2).[17]

In implementing the *System-GMM* estimator we utilise external instruments based on lagged values of log population size, POP. This term is the time-varying element of a standard gravity model of trade flows; see for example Frankel and Romer (1999). Although gravity equations typically use population size to explain cross-sectional differences in openness, we find that past

[16] To be precise, the GMM estimator for equation (3) uses the following moment conditions: $E(VINF_{i,t-s}\Delta\varepsilon_{it}) = 0$; $E(OPEN_{i,t-s}\Delta\varepsilon_{it}) = 0$ for $t = 3, 4, ...T$, and $s \geqq 2$.

[17] Specifically, the following additional moment conditions are available: $E(\Delta VINF_{it-1}(\eta_i + \varepsilon_{it})) = 0$, $E(\Delta OPEN_{it-1}(\eta_i + \varepsilon_{it})) = 0$.

population size helps to predict the evolution of openness and can therefore be used to increase the efficiency of the GMM estimator. The role of the external instrument is further examined in the discussion of the empirical results in Section 4.

The validity of the instruments can be evaluated using the Sargan test of over-identifying restrictions and Lagrange Multiplier tests for the absence of second order serial correlation (Arellano and Bond 1991). It is important to note that the first differenced transformation yielding (3) induces an MA(1) error structure, and therefore we expect that the first-differenced residuals will be negatively autocorrelated at the first lag but uncorrelated at the second lag. The estimated standard errors take account of the first-order negative autocorrelation and any heteroscedasticity in the residuals, see Arellano and Bond (1991). As recommended by Blundell and Bond (1998) the estimates that we report are based on 1-step GMM estimation in which equal weight is placed on each moment condition.[18]

4 Empirical results

In Table 1 we present our basic empirical results. Columns $1-4$ list the ordinary least squares (OLS), within groups (WG), *Differenced-GMM* and *System-GMM* estimates of a model in which inflation volatility depends on its own lag and openness, plus a full set of time dummies.[19] The *Differenced-GMM* estimates use as instruments $VINF_{t-2}$, $OPEN_{t-2}$, POP_{t-2} and POP_{t-3} and the *System-GMM* estimates use as additional instruments $\Delta OPEN_{t-1}$ and ΔPOP_{t-2}.[20] In each case the effect of openness on inflation volatility is negative and this relationship is propagated through time by the positively signed autoregressive term.

The model that we emphasise is the *System-GMM* estimate in column 4 which shows that

[18] All estimations are conducted using the DPD package in *Pc-Give*, see Doornik and Hendry (2001).

[19] These are estimates of equation (3) except in the case of the OLS results which are estimates of equation (2), i.e. the relationship between inflation volatility and openness before differencing out purely cross-sectional variation in the data.

[20] The differenced lagged dependent variable is not used as an instrument in the levels part of the system estimator because the marginal restrictions required in order for it to be a valid instrument were rejected by a Sargan test. It appears that there has been some 'inflation volatility convergence' during the sample period - countries with initially high volatility experience relatively large reductions in volatility during later periods. This implies that $\Delta VINF_{it-1}$ is not orthogonal to η_i in equation (2).

$OPEN$ impacts $VINF$ with a coefficient of $-.086$. The within groups standard deviation of openness is 0.198 while that for inflation volatility is 0.067. Based on this estimate, a one standard deviation increase in openness yields a 0.25 standard deviation reduction in inflation volatility in the first 5 years. The lagged dependent variable implies further reductions in inflation volatility depending on the persistence of the increase in openness. If a one standard deviation increase in openness is sustained the eventual reduction in inflation volatility will be approximately twice that observed in the first five years, i.e. on the order of 0.5 standard deviations, though it should be noted that additional controls have not yet been introduced.

Before investigating the robustness of our results we consider the properties of the instruments. The Sargan test for the validity of the over-identifying moment conditions used in column 4 yields a *p-value* of 35%. In order to check that this outcome is not a Type II error based on pooling valid and invalid instruments we perform separate Difference-Sargan (D-Sargan) tests for the moment conditions associated with each variable. The *p-values* are 20% ($VINF$), 76% ($OPEN$) and 34% (POP), suggesting that each type of instrument is individually valid. Furthermore, the AR(1) and AR(2) tests provide strong support for the hypothesis that the errors in (2) are serially uncorrelated, a necessary condition for instrument validity.

A related question concerns the explanatory power of the instruments. If the instruments are weak the exogenous variation in openness will be limited and this may distort inference (Stock, Wright and Yogo 2002). To address this issue we regressed $\Delta VINF_{t-1}$ and $\Delta OPEN_{it}$ on the instruments used for the differenced equation, and $VINF_{t-1}$ and $OPEN_t$ on the instruments used for the levels equation, and performed *F-tests* for the joint significance of the regressors. The test statistics were 27.96 ($\Delta VINF_{t-1}$ equation), 23.16 ($\Delta OPEN_{it}$), 20.04 ($VINF_{t-1}$) and 83.29 ($OPEN_t$), each of which is significant at the 0.1% level.[21] Hence, the instruments appear to have considerable explanatory power.[22]

In the final two columns of Table 1 we take a further look at the role of the external instru-

[21] Each regression contained a full set of period dummies, but these dummies are *not* included in the *F-tests*.

[22] On a related theme, Blundell and Bond (1998) show that a GMM estimate of the autoregressive parameter that is close to the WG estimate typically reflects a problem of weak instruments. The centrality of the *Differenced-GMM* and *System-GMM* estimates with respect to the OLS-WG range is further evidence that our results are not due to weak instruments.

ment, POP. In column 5 all terms in POP are dropped from the instrument set. A comparison of these results with those in column 4 indicates that the main role of POP is to increase the efficiency of the estimation. This is seen most clearly in the case of the openness coefficient, which is actually of greater magnitude in column 5 than in column 4 but yields a smaller *t-ratio* because its standard error increases three-fold. Hence, the external instrument does not induce the sign or magnitude of $OPEN$ but instead increases the precision of the estimation. In column 6 we address the possibility that the significance of $OPEN$ is due to the contemporaneous value of POP having been excluded from the regression (recall that lagged values of POP are used in the instrument set and therefore a regression in which POP enters contemporaneously is still identified). The results indicate that the use of population size only as an instrument, rather than as both an instrument and a regressor, is not critical to our main result.

4.1 Robustness: Adding further controls

In this sub-section we add further controls to the basic model. The largest sample for which all of the regressors are available comprises 451 observations drawn from 84 countries and this is the sample that we use in each column of Table 2 (the 12 countries that drop out of the 96 country sample used in Table 1 are listed in appendix A). In order to conserve space we focus on the *System-GMM* estimates.

The first column reproduces the simple specification for the new sample size. The magnitude of each coefficient falls slightly but the qualitative results are robust. Column 2 controls for the natural log of one plus mean inflation (INF) and uses as instruments INF_{t-3} and ΔINF_{t-2} (instruments at shorter lags are invalid according to a D-Sargan test). Mean inflation is highly significant, reflecting its strong correlation with inflation volatility. The openness coefficient falls to $-.044$, a much more modest effect than that estimated in column 4 of Table 1, but owing to greater precision in the estimation it remains significant at the 5% level, i.e. openness reduces inflation volatility even amongst countries that have the same average inflation. The autoregressive parameter is close to zero after controlling for mean inflation. Indeed, deleting the autoregressive term from column 2 leaves the results practically unchanged - the openness coefficient remains $-.044$ and the *t-ratio* is 2.26 (static estimates of each of the regressions in columns $3-10$ are available on request).[23]

We experimented with two variations on the column 2 specification. Firstly, given the important relationship between mean inflation and inflation volatility we defined the regression in terms of the log coefficient of variation for inflation (the standard deviation of inflation relative to its mean). A *System-GMM* regression of this term on its first lag plus openness yields an openness effect that is significant at the 5% level. Secondly, we experimented with a non-linear relationship between the first two moments of inflation: Adding the square of mean inflation to column 2 gives an openness coefficient of $-.041$ and a corresponding *t-ratio* of 2.11 (full details of these two experiments are available on request).

In column 3 we control for the natural log of GDP per capita ($RGDP$) and add the second lag of that variable to the instruments.[24] This is measured in 1996 US$ and corresponds to the first year from each of the 5 year windows for which inflation volatility is measured. The effect of $RGDP$ is negative but close to zero. Further (unreported) experimentation showed that this is due to INF having been included in the regression; much of the effect of $RGDP$ is mediated through mean inflation. The openness coefficient is further diminished relative to column 1 but is significant at the 6% level. Column 4 controls for the log product of population and per capita income, a measure of economic size. This is a potentially important control when analysing the effects of openness, see Lane (1997), but does not alter the main results in this case.

We next address the possibility that openness increases volatility through inducing specialisation in production, as discussed in the previous section. Such an effect is arguably most relevant in the case of countries whose trading links lead to specialisation in primary commodities, which are traded in some of the most volatile international markets. In order to address this possibility we define a dummy variable, $PRIMEXP$, that is equal to unity for countries for which more than 50% of exports during the period $1988-92$ are fuels or other primary commodities, and zero otherwise. In column 5 the negative impact of openness is weaker amongst countries with

[23] It should be noted that although originally proposed for dynamic panel models, the *System-GMM* technique is an efficient estimator for static panels and has often been used in this context, see Beck (2002).

[24] Unless otherwise stated, regressions $3-9$ use the second lag of the marginal variable as an additional instrument. Instruments based on lagged first differences of the marginal terms are not used because we found that in some specifications the Sargan *p-value* was close to unity, which is a sign that the instrument set is too large and that estimation may be imprecise.

the most concentrated export structures, but the difference is not significant at conventional levels and the marginal effect of openness is negative amongst both groups of countries.[25]

In section 2 we also noted that openness may increase inflation volatility if it is associated with larger capital flows, which are often pro-cyclical and can amplify volatility. In order to investigate this issue we define the variable $CAPFLOWS$ as the log ratio of private capital flows to GDP. This index is available for just 325 of the 451 episodes included in the Table 2 regressions and therefore it is not included in those results. However, when $CAPFLOWS$ is added to column 2 in Table 2 the coefficient on $OPEN$ is $-.061$ and the *t-ratio* is 2.71. Hence, trade openness exerts a more powerful negative effect on inflation volatility when holding constant exposure to capital flows (although the $CAPFLOWS$ term itself is insignificant in that regression).

The next idea that we explore is that the decline in inflation volatility is due to the size of supply and demand shocks having fallen exogenously, and that the link between openness and inflation volatility is coincidental. In order to proxy the size of these shocks we control for the natural logs of one plus the decimal standard deviation of annual output growth (VOL) and one plus the decimal standard deviation of the trade weighted mean of output growth in a country's five largest trading partners ($TPVOL$).[26] In column 6 the measure of domestic output volatility is positively signed but insignificant at the 5% level, while in column 7 the measure of foreign output volatility, $TPVOL$, is positively signed and significant at the 5% level. The key point, however, is that the negative and significant effect of trade openness on inflation volatility is robust in both cases.

In column 8 we control for the natural log of one plus the average rate of economic growth ($GROWTH$), the rationale being that during 'good times' inflation volatility may be more easy

[25] A caveat that should be added here is that $PRIMEXP$ is time invariant. A better measure would allow for time variation in this index.

[26] The second measure of volatility is more exogenous and arguably more likely to be positively associated with inflation volatility. For example, domestic output volatility may vary positively with inflation volatility due to the common shocks effect, but negatively with inflation volatility if openness increases output volatility, e.g. through specialisation effects, at the same time as reducing inflation volatility. Foreign output volatility is unlikely to be affected by domestic trade policies but remains subject to the international supply and demand shocks, and is therefore more obviously a positive predictor of inflation volatility.

to control. The results indicate some evidence for this and the openness effect is diminished and significant at only the 10% level. However, in the general model reported in column 10, which controls for $GROWTH$ and each of the other regressors considered, $OPEN$ is significant at the 5% level.

In column 9 we control for the measure of political constraints on the executive branch of government developed by Henisz (2000) and denoted PCI. Satyanath and Subramanian (2004) argue that concentration of power amongst a small number of actors in the executive branch of government can increase nominal volatility as temporary bouts of surprise inflation are used to effect real resource transfers within society. A large number of veto points in the policy process exerts a negative effect on inflation volatility in column 9 but the effect is far from significant, possibly reflecting the limited time variation in PCI relative to inflation volatility. In contrast, both the size and sign of the openness effect are robust.[27]

4.2 Robustness across sub-samples

In this section we present results based on various sub-samples. In column 1 of Table 3 we exclude the 5% most extreme values for inflation volatility and openness (2.5% from each tail of each unconditional distribution). This sub-sample omits Hong Kong and Singapore, an important robustness check because trade ratios may overstate the openness of these countries given that many imports are almost immediately exported. Another rationale for excluding outliers is that the boundedness of openness and inflation volatility (neither can turn negative before the log transformation) raises the possibility of skewed distributions which may affect the small sample distribution of the GMM estimator and call for alternative estimation methods such as those presented in Papke and Wooldridge (2008).[28] We leave the application of such methods for future research and instead use outlier exclusion as one simple way to mitigate skewness in the sample. The effect of openness is robust.

The next question that we address is whether or not our results are driven by exchange

[27] In addition to the controls listed in Table 2 we considered measures of financial depth, government size and the share of agriculture in GDP. The negative effect of openness was robust in each case. Full details are available on request.

[28] We are grateful to an anonymous referee for pointing out this possibility.

rate regime choices. A fixed exchange rate combined with capital mobility is thought to restrict discretionary monetary policy by forcing a country to follow foreign monetary policy, and this may reduce inflation volatility. Alfaro (2005) finds that the exchange rate regime choice is a more important determinant of temporal inflation patterns than is openness to trade. In columns 2 and 3 we present results for fixed and floating exchange rate samples. The first sub-sample comprises observations from regimes in which the exchange rate is classified by Reinhart and Rogoff (2004) as a peg or a crawling peg with bands of $\pm 2\%$, while the second sub-sample is based on regimes in which the exchange rate is judged freely floating or adjusting within bands of $\pm 5\%$.[29] The results show that openness reduces inflation volatility in both the fixed and floating exchange rate sub-samples, though in the latter case the effect is significant at only the 10% level. The fact that openness appears to reduce inflation volatility in the fixed exchange rate sub-sample suggests that the relationship is not an artefact of pass-through from exchange rates to consumer prices having fallen in recent years, and greater openness having co-incided with that trend, because exchange rate fluctuations were clearly minimal in the fixed exchange rate sample.

In column 4 we exclude the 38 countries that are not awarded at least a grade C for data quality by Summers and Heston (1988). The resulting 'good data' sample yields an openness effect that is significant at the 5% level, suggesting that our findings are not due to low data quality in closed economies leading to spurious volatility. The next sub-sample comprises the 68 low debt countries listed in Terra (1998). Terra argues that the negative relationship between the levels of openness and inflation exists only amongst heavily indebted countries; here we investigate whether the same is true of openness and inflation volatility. Openness is significant at the 5% level and its coefficient is of similar magnitude to that obtained in Table 2.

In column 6 we omit the 20 countries that maintained an inflation targeting regime during some part of the sample (dates are taken from Fatas, Mihov and Rose, 2007).[30] The openness

[29] Reinhart and Rogoff (2004) also report a fifth category of 'freely falling' exchange rates, often observed during periods of economic crisis. We do not include observations drawn exclusively from these crisis periods because they are likely endogenous to inflation volatility.

[30] Two countries, South Africa and Thailand, adopted inflation targeting during 2000, the final year of the sample. As a reform occurring in this year is unlikely to affect our results we do not exclude these countries.

effect is robust, suggesting that our main results are not due to inflation targeting schemes having caused a reduction in inflation volatility and such changes being correlated with openness by chance.

The next hypothesis that we address is that our results arise because structural adjustment programmes associated with IMF/World Bank loans secure both macroeconomic stability and trade openness. In column 7 we exclude 18 countries identified by Easterly (2005) as being amongst the top 20 recipients of IMF and World Bank loans (the other two countries, Bangladesh and Mali, are not part of our sample). The negative effect of openness remains intact, suggesting that our results do not depend on the effects of IMF/World Bank interventions.

Column 8 focuses on 71 developing and emerging market economies.[31] The effect of openness is larger and more significant in the developing country sample than in the full sample. The opposite is true in column 7, which looks only at the 23 OECD countries excluded from the column 6 sample ($\Delta RGDP_{t-1}$ is added to the instruments used in column 7 because initial results indicated very imprecise estimation). These results may arise because there exist reasons to pursue macroeconomic stability in OECD countries even in the absence of openness, e.g. Posen (1993) emphasises the importance of financial sector organisations in lobbying policy-makers in industrial countries. Similarly, consumption patterns in OECD countries may have been diversified at the start of the sample, in which case greater trade openness is less likely to reduce inflation volatility through this channel.

4.3 Robustness to varying the data frequency

In Table 4 we consider the sensitivity of our results to changing the data frequency. Column 1 uses a measure of inflation volatility based on 5 year windows but calculated from annual rather than quarterly data. Although quarterly data provide more observations for calculating volatility, they may be subject to larger measurement errors than are annual data. The effect of openness in column 1 is similar to that obtained for equivalent specifications based on quarterly

[31] These are the 96 countries in the core sample minus 23 countries that have been OECD members since 1961. Turkey has been an OECD member since 1961, but we include Turkey in the 71 country sub-sample on the grounds that it is best regarded as an emerging market economy. Hong Kong and Singapore are excluded even though they are not OECD members (including these two countries does not change the results).

data, however, suggesting that possible measurement errors in quarterly data do not drive our results.

The second column of Table 4 measures inflation volatility using quarterly data and 8 year windows ($1961-68$, $1969-76$ and so on) whilst the third column uses 3 year windows ($1961-63$, $1964-66$ and so on). As discussed at the start of section 3, a key assumption in our measurement of volatility is that mean inflation is stable within 5 year windows. If there are changes to mean inflation within these observation periods the $VINF$ statistic will over-state volatility as mean shifts are conflated with genuine inflation variance. An even more problematic case arises if inflation is trended within each 5 year window, since then mean inflation changes each period and all such changes would be incorrectly recorded as inflation volatility. Natural responses to such possibilities are to reduce the window length as much as possible, to minimise the number of cases in which windows span mean shifts, or to first detrend inflation and calculate detrended inflation volatility, but there are trade-offs associated with such choices. When mean inflation is constant but shocks to inflation are persistent, very short window lengths will under-state volatility as persistent fluctuations are treated as mean shifts. Similarly, over a 5 year window a persistent inflation shock could give the impression that inflation is trended and fitting a trend would lead to volatility being under-stated. Our response to this trade-off is to consider alternative window lengths both shorter and longer than the 5 year windows in our baseline analysis. The results show that although the magnitude of the effect of openness on inflation volatility changes, it remains negative and statistically significant in both cases. We highlight in particular the relationship estimated from 3 year windows in the final column of Table 4. If our main results from 5 year windows are driven by spurious movements in volatility linked to non-modelled trends in inflation then one would expect the results to be sensitive to the use of shorter windows, since these would limit the amount of trend variation that can be mis-classified as volatility. In the final column of Table 4 the estimated effect of openness is smaller than in the baseline estimate in column 4 of Table 1, but its statistical significance actually increases such that the effect is significant at the 1% level. Thus, whilst we have not explicitly accounted for possible inflation trends within observation windows, we argue that our results provide evidence that the impact of openness on inflation volatility is not driven by spurious volatility that might arise from such trends.

4.4 Robustness to extending the sample to 2015

To ensure that our results are not driven by the choice of the sample period, we replicate the main estimations for the extended sample period, $1961-2015$. The expanded sample necessitates the inclusion of additional temporal dynamics.[32] To do so, we include in our main specification two lags of the dependent variable, as opposed to one. The results are presented in Table 5. As before, columns $1-3$ provide results based on OLS, within-groups and GMM estimators. The results suggest strong persistence effects, as demonstrated by the statistically significant coefficients on the first two lags of inflation volatility (especially in columns 2 and 3). Reassuringly, our openness measure continues to be a significant predictor of inflation volatility. Moving to System GMM estimates in column 4, the coefficient on openness is once again negative and statistically significant at the 10% level but not at the 5% level. This result continues to hold even if we drop population from the instrument set (column 5). The slightly reduced statistical significance of trade openness in explaining inflation volatility is in part driven by a lower point estimate on openness in regressions fitted over the extended period, which is consistent with the evidence on falling inflation volatility over time.

Taken together, the results in Table 5 support our argument and show that the negative relationship between trade openness and inflation volatility is robust to both extending the sample period to 2015 and including additional dynamics (i.e., further lags of the dependent variable), though the magnitude and statistical significance of the relationship diminishes somewhat.

5 Summary and concluding remarks

In this paper we have considered two mechanisms that link inflation volatility and openness to international trade. Firstly, openness impacts policy-maker incentives in a way that can either increase or decrease monetary policy discipline. Accepting that inflation volatility outcomes are closely linked to the volatility of policy, the implied relationship between openness to trade and

[32] It should be noted that the inclusion of two lags of inflation volatility results in some sample attrition and this limits the increase in our sample size as a result of extending the time dimension of the panel to 2015.

inflation volatility is ambiguous, and the commonly held view that stronger trade links induce policy discipline and macroeconomic stability need not follow (Cavelaars, 2009). Secondly, openness to trade leads to greater diversity in consumption, which can reduce inflation volatility through decreasing the sensitivity of consumer prices to price shocks in specific markets. On the other hand, openness may induce specialization in production that increases exposure to volatility.

The main contribution of the paper was to use pre-determined variables as instruments to identify exogenous movements in national trade shares and then estimate the effects of those changes in trade openness upon inflation volatility. The principal finding was that countries that have opened up to trade more rapidly than the global average have experienced larger reductions in inflation volatility, suggesting that the channels supporting a negative effect of trade openness on inflation volatility have been dominant in recent episodes. The relationship was consistently significant at the 5% level, although in the most general regressions that we considered the quantitative importance of the relationship was modest.

A second finding to emerge from the analysis was that amongst a sub-sample of OECD countries the negative impact of openness on inflation volatility was much weaker than amongst developing and emerging market economies. One explanation for this finding is that the mechanisms that we have proposed are less effective in high income countries. Such countries may pursue disciplined macroeconomic policies even in the absence of exposure to foreign competition, e.g. if firms in the financial services sector lobby for such an outcome (Posen 1993). Similarly, consumer price inflation in developed countries may be stable even in the absence of a highly diversified consumption basket, for example if spending is concentrated on high value added products for which prices are typically less volatile.

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Appendix A: Countries included in Table 1 regressions

Algeria, Argentina, Australia, Austria, Barbados, Belgium, Bolivia, Burkina Faso, Burundi, Cameroon, Canada, Central African Republic, Chad, Chile, Colombia, Cost Rica, Cyprus, Democratic Republic of Congo, Dominican Republic, Denmark, Ecuador, Egypt, El Salvador, Ethiopia, Fiji, Finland, France, Gabon, Gambia, Germany, Ghana, Greece, Guatemala, Haiti, Honduras, Hong Kong, Iceland, India, Indonesia, Ireland, Israel, Italy, Ivory Coast, Jamaica, Japan, Jordan, Kenya, Korea, Lesotho Liberia, Luxembourg, Madagascar, Malawi, Malaysia, Malta, Mauritania, Mauritius, Mexico, Morocco, Niger, Nigeria, Nepal, Netherlands, New Zealand, Norway, Pakistan, Panama, Papa New Guinea, Paraguay, Peru, Philippines, Portugal, Rwanda, Senegal, Sierra Leone, Singapore, South Africa, Spain, Sri Lanka, Sudan, Surinam, Swaziland, Sweden, Switzerland, Tanzania, Thailand, Togo, Trinidad and Tobago, Turkey, Uganda, Uruguay, United Kingdom, United States, Venezuela, Zambia, Zimbabwe.

The following 12 countries drop out of the sample used in Table 2 due to data for some of the controls being unavailable for those countries: Belgium, Fiji, Iceland, Liberia, Lesotho, Luxembourg, Malta, Mauritania, Sudan, Singapore, Surinam, Swaziland.

Appendix B: Data sources

$VINF, INF$: *International Financial Statistics*, line 64

$OPEN$, POP and $RGDP$: Penn World Tables, version 6.1 (October 2002)

$PRIMEXP$: *World Development Indicators*

VOL: Calculated from $RGDP$, see above for source

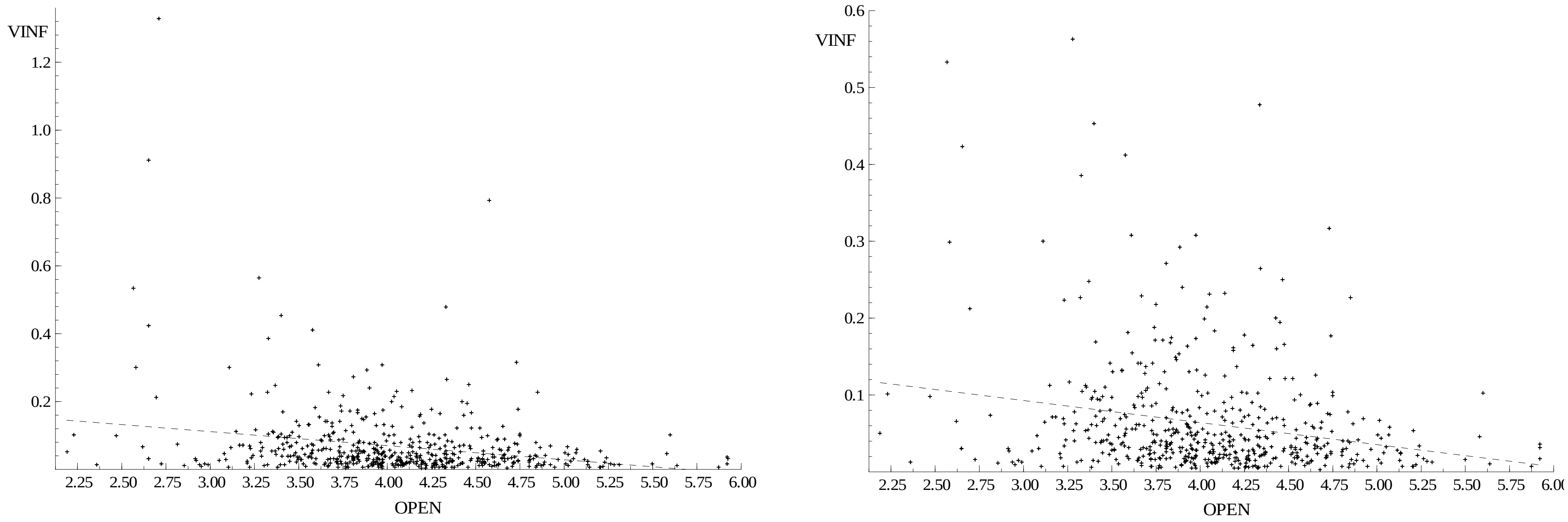
$TPVOL$: IMF's *Direction of Trade Statistics* database

$GROWTH$: Calculated from $RGDP$, see above for source

$CAPFLOWS$: *World Development Indicators*

PCI: Henisz (2000)

FIGURE 1: SCATTER PLOTS FOR OPENNESS (X-AXIS) AND INFLATION VOLATILITY (Y-AXIS)



Notes: Inflation volatility (VINF) is the log of one plus the decimal standard deviation of quarterly observations on the annual inflation rate. Openness (OPEN) is the log of exports plus imports as a percentage of GDP. The left-hand side graphs plots the 538 observations used to fit column 4 of Table 2. The right-hand side graph is the same except that the three largest readings for inflation volatility are omitted.

TABLE 1: BASIC RESULTS

	DEPENDENT VARIABLE— *Inflation volatility (VINF)*					
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
	OLS	*WG*	*GMM*	*System*	*System*	*System*
$VINF_{t-1}$	.737	.384	.530	.581	.729	.587
	(4.13)	(1.91)	(2.36)	(3.28)	(2.82)	(3.06)
OPEN	-.021	-.056	-.124	-.086	-.111	-.104
	(2.58)	(2.51)	(1.94)	(2.34)	(1.03)	(2.48)
POP						-.001
						(0.09)
	DIAGNOSTIC TESTS (P-VALUE)					
SERIAL CORR. TESTS						
First-Order	0.632	0.113	0.017	0.000	0.001	0.005
Second-Order	0.196	0.017	0.762	0.658	0.579	0.826
Sargan	-	-	*0.547*	*0.348*	*0.403*	*0.336*
	Difference-Sargan statistics for column (4): VINF(0.20) OPEN(0.76) POP(0.34)					
NT	538	538	442	538	538	512
	INSTRUMENTS					
	VINF(t-2) OPEN (t-2) POP(t-2, t-3)					
	Instruments for level equations : ΔOPEN (t-1) ΔPOP(t-2)					

Notes:
Estimates are based on a sample of 96 countries. Period dummies are included in all specifications (but are not reported) and are also part of the instrument set. Numbers in parentheses are absolute t-statistics based on robust standard errors. OLS denotes ordinary least squares, WG denotes within groups, GMM denotes 1-step generalised method of moments estimation of the first differenced equation and System denotes 1-step joint generalised method of moments estimation of the first differenced and levels equations.
The Sargan and Difference-Sargan tests of over-identifying restrictions are based on 2-step GMM estimates in order to correct for heteroscedasticity and are asymptotically distributed as $\chi2(n-p)$, where n is the number of moment conditions and p is the number of parameters. The serial correlation tests are asymptotically distributed as $N(0,1)$ under the null of no serial correlation.

TABLE 2: ADDITIONAL CONTROLS

	DEPENDENT VARIABLE— *Inflation volatility (VINF)*									
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*	*(8)*	*(9)*	*(10)*
$VINF_{t-1}$	.367	.020	.007	.002	-.002	.013	.016	-.023	.004	-.052
	(2.01)	(0.33)	(0.10)	(0.03)	(0.03)	(0.22)	(0.27)	(0.43)	(0.06)	(0.87)
OPEN	-.064	-.044	-.037	-.040	-.045	-.046	-.040	-.028	-.040	-.036
	(2.02)	(2.19)	(1.92)	(2.15)	(2.35)	(2.27)	(2.17)	(1.84)	(2.03)	(2.05)
INF		.609	.609	.610	.589	.629	.604	.579	.615	.606
		(9.51)	(9.46)	(9.50)	(8.43)	(10.60)	(9.20)	(6.89)	(10.20)	(8.95)
RGDP			-.003							.004
			(0.47)							(0.36)
*RGDP*POP*				-.004						.002
				(0.67)						(0.23)
*PRIMARY*OPEN*					.007					.008
					(0.97)					(0.79)
VOL						.304				-.091
						(1.13)				(0.42)
TPVOL							2.498			1.391
							(2.03)			(1.30)
GROWTH								-.465		-.303
								(1.86)		(1.24)
PCI									-.015	-.018
									(0.36)	(0.29)
	DIAGNOSTIC TESTS (P-VALUE)									
SERIAL CORR. TESTS										
First-Order	0.002	0.001	0.001	0.001	0.001	0.001	0.004	0.003	0.002	0.003
Second-Order	0.765	0.364	0.350	0.355	0.351	0.228	0.299	0.129	0.358	0.175
Sargan	*0.174*	*0.248*	*0.245*	*0.387*	*0.246*	*0.537*	*0.161*	*0.306*	*0.278*	*0.721*
Diff-Sargan		*0.91*	*0.52*	*0.75*	*0.66*	*0.82*	*0.20*	*0.32*	*0.16*	
NT	451	451	451	451	451	451	451	451	451	451

See notes to Table 2. The instruments in column 1 are as in Table 2 except that in the first half of the system *OPEN(t-2, t-3)* is used. In columns 2-10 the additional instrument is the marginal regressor at *t-2*, except for the variables *INF*, *GROWTH* and *PCI* for which *t-3* is used. Δ*INF(t-2)* is used to instrument the levels equations in columns that control for *INF*. Column 6 uses Δ*VOL(t-2)* as an instrument in the second half of the system. *Diff-Sargan* tests the validity of moment conditions based on the marginal instruments.

TABLE 3: SUB-SAMPLE ANALYSIS

	DEPENDENT VARIABLE — *Inflation volatility* (*VINF*)								
	(1) *Extreme obs Omitted*	*(2)* *Fixed ex rates*	*(3)* *Floating ex rates*	*(4)* *Good data*	*(5)* *Low debt*	*(6)* *No inf targeters*	*(7)* *Min IMF Intervention*	*(8)* *Non-OECD*	*(9)* *OECD*
$VINF_{t-1}$	.022	-.113	.064	.075	-.038	.072	.069	-.063	.162
	(0.31)	(0.80)	(1.04)	(1.05)	(0.62)	(0.75)	(1.10)	(0.87)	(1.17)
OPEN	-.042	-.042	-.047	-.029	-.041	-.040	-.047	-.059	-.010
	(2.17)	(2.03)	(1.72)	(2.07)	(2.10)	(2.02)	(2.08)	(2.47)	(1.33)
INF	.425	.350	.588	.629	.432	.619	.478	.634	.179
	(7.28)	(2.02)	(8.55)	(9.98)	(3.99)	(7.48)	(6.64)	(10.10)	(1.67)
	DIAGNOSTIC TESTS (P-VALUE)								
SERIAL CORR. TESTS									
First-Order	0.000	0.004	0.016	0.086	0.007	0.001	0.002	0.000	0.074
Second-Order	0.568	0.301	0.092	0.227	0.907	0.265	0.792	0.271	0.612
Sargan	*0.483*	*0.664*	*0.570*	*0.324*	*0.161*	*0.678*	*0.572*	*0.794*	*1.000*
NT	478	286	218	353	402	413	452	376	154

See notes to Table 2. The instrument set varies slightly. Col (4) omits *OPEN(t-2)* and Δ*POP(t-2)*. Col (8) omits *OPEN(t-3)* and Δ*POP(t-2)*. Col (7) omits *OPEN(t-3)*, *POP(t-3)* and Δ*OPEN(t-1)*.

TABLE 4: VARYING THE DATA FREQUENCY

DEPENDENT VARIABLE— *Inflation volatility* (*VINF*)			
	System GMM Estimates		
	(1)	(2)	(3)
	Measure volatility using annual data	*Quarterly data and eight year windows*	*Quarterly data and three year windows*
$VINF_{t-1}$	-.116	.053	.052
	(1.21)	(0.21)	(0.95)
OPEN	-.053	-.103	-.069
	(2.07)	(2.74)	(2.63)
INF	.593	.494	.551
	(7.72)	(3.07)	(4.16)
	DIAGNOSTIC TESTS *(p-value)*		
SERIAL CORR. TESTS			
First-Order	0.002	0.724	0.069
Second-Order	0.213	0.800	0.690
Sargan	*0.546*	*0.549*	*0.576*
NT	*538*	272	967
	Instruments *VINF(t-2) OPEN (t-2, t-3) POP(t-2, t-3) INF(t-3)*		
	Instruments for level equations : ΔOPEN (t-1) ΔPOP(t-2) ΔINF(t-2)		

See notes to Table 2. Column (1) excludes Δ*POP (t-2)* from the instrument set.

TABLE 5: RESULTS FOR THE EXTENDED SAMPLE PERIOD

	DEPENDENT VARIABLE— *Inflation volatility (VINF)*				
	(1)	*(2)*	*(3)*	*(4)*	*(5)*
	OLS	*WG*	*GMM*	*System*	*System*
$VINF_{t-1}$	.573	.373	.456	.553	.536
	(8.60)	(8.03)	(11.53)	(13.99)	(14.14)
$VINF_{t-2}$	-.104	-.265	-.185	-.125	-.142
	(1.26)	(3.72)	(2.27)	(1.83)	(2.01)
OPEN	-.025	-.057	-.048	-.016	-.050
	(3.28)	(2.68)	(1.93)	(1.84)	(1.89)
	DIAGNOSTIC TESTS (P-VALUE)				
SERIAL CORR. TESTS					
First-Order	-	-	0.032	0.031	0.031
Second-Order	-	-	0.422	0.290	0.316
NT	603	603	511	603	603
	INSTRUMENTS				
	VINF(t-3) OPEN (t-2) POP(t-2, t-3)				
	Instruments for level equations : ΔOPEN (t-1) ΔPOP(t-2)				

Notes:
Estimates are based on a sample of 96 countries with data extended to the year 2015. Additional temporal dynamics have been considered through the inclusion of the 2-period lag of the dependent variable. Period dummies are included in all specifications (but are not reported) and are also part of the instrument set. Numbers in parentheses are absolute t-statistics based on robust standard errors. OLS denotes ordinary least squares, WG denotes within groups, GMM denotes 1-step generalised method of moments estimation of the first differenced equation and System denotes 1-step joint generalised method of moments estimation of the first differenced and levels equations. The serial correlation tests are asymptotically distributed as *N(0,1)* under the null of no serial correlation. Additional temporal dynamics have been considered through the inclusion of the two-period lags of the dependent variable. Column 5 drops population as an external instrument.