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Tracking U.S. inflation expectations with domestic and global indicators

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A B S T R A C T

Are foreign variables important for tracking U.S. inflation expectations? This paper estimates a reduced-form model that takes both *domestic* and *global* indicators of economic slack and inflationary pressures into account. Our main findings point towards the instability of the estimated parameters over the last four decades. In particular, global indicators appear to have played a statistically significant role in shaping forecasters' expectations until the mid-1980s. By contrast, the U.S. monetary policy stance turns out to be relevant in the 1980s and 1990s. We relate this finding to the more aggressive monetary policy conduct implemented by the Fed since the end of the Volcker experiment.

1. Introduction

> "The challenge that lies before the Committee is to manage policy in a way that permits the economy to realize its productive potential while simultaneously maintaining firm control of inflation and *inflation expectations*." (Ben S. Bernanke, Remarks on the Economic Outlook and Monetary Policy, Annual Meeting of the Bond Market Association, New York, April 22, 2004, emphasis added)
> "[…] to make effective policy, the Federal Reserve must have a full an understanding as possible of the factors determining economic growth, employment, and inflation in the U.S. economy, *whether those influences originate at home or abroad*." (Ben S. Bernanke, Remarks on Globalization

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and Monetary Policy, Fourth Economic Summit at the Stanford Institute for Economic Policy Research, Stanford, March 2, 2007, emphasis added)

Inflation expectations are an element of key importance for monetary policy makers.[1] When thinking of a simple model for inflation expectations in a given country, it is natural to relate such expectations to *domestic* factors (i.e. factors specific to that country). However, some recent contributions (Rogoff, 2003; Ciccarelli and Mojon, in press; Mumtaz and Surico, 2008; Borio and Filardo, 2007) have stressed the role potentially played by *global* factors (i.e. factors regarding country-aggregates such as the G7 or the OECD) in affecting U.S. inflation. If one country's inflation is mainly driven by global forces, central bankers might have the incentive to coordinate at an international level in order to offer a global response to global shocks. Of course, given the role played by inflation expectations in influencing inflation (e.g. Woodford, 2003), it is crucial to understand to what extent global factors have actually influenced *expected* inflation.[2]

This paper aims at assessing the link existing between U.S. inflation expectations and international (i.e. global) forces at an empirical level. To do so, we estimate a battery of simple reduced-form models for U.S. inflation expectations. As regressors, we consider both standard domestic indicators such as U.S. inflation, the U.S. output gap, different measures of inflationary pressures, and the U.S. monetary policy conduct on the one hand, and global measures of inflation and the business cycle on the other. We conduct our empirical analysis by proceeding in two steps. First, we assume absence of breaks in the estimated relationships, and we estimate fixed-coefficient models. Then, we investigate the issue of parameter stability by running rolling-window regressions as well as sub-sample regressions.

Several results arise. We find full sample evidence for the global output gap and global inflation to be drivers of U.S. inflation expectations. Further checks show that these global variables add information with respect to a large variety of standard measures of internal and external pressures (e.g. unit labor costs, trade openness, global liquidity, financial pressures). Interestingly enough, this full sample empirical evidence turns out *not* to be robust across different subsamples. In fact, rolling-window regressions reveal that the relevance of our global indicators is not stable over time, i.e. it tends to disappear when crossing the mid-1980s. We argue that this break might be due to the aggressive monetary policy conduct implemented by the Fed at the end of the Volcker experiment. A subsample analysis confirms the significance of a measure of monetary policy stance in the last two decades, so corroborating our conjecture on the 'substitution' between global and domestic forces that might have occurred in the mid 1980s.

The structure of the paper is the following. Section 2 presents the time-series of interest and the empirical model we employ for tracking inflation expectations, and it discusses our full sample results. Section 3 analyzes the parameter instability issue, and it interprets the findings stemming from our subsample estimates. Section 4 proposes some further empirical investigations corroborating the results presented in Section 3. Section 5 concludes.

2. Tracking the U.S. inflation expectations: model and evidence

We aim at tracking the short-term U.S. inflation expectations with a simple reduced-form model. Following Erceg and Levin (2003), we concentrate on inflation expectations as reported by the Survey

[1] In this paper, we concentrate on U.S. short-term inflation expectations. For contributions dealing with long-term inflation expectations in various countries, see Castelnuovo et al. (2003) and Gürkaynak et al. (2005, 2006).

[2] In the standard new-Keynesian model à la Clarida et al. (1999), realized inflation is the sufficient statistic for expected inflation, i.e. $E_t\pi_{t+1} = \rho\pi_t$, where ρ is the autoregressive parameter of the AR(1) process for the cost-push shock. However, such a model does not capture the well-known evidence in favor of the existence of lags in the monetary policy transmission mechanism. Consider a model more suited for capturing the mentioned lags, i.e. a simplified version of the AD/AS model proposed by Rudebusch and Svensson (1999): $\pi_t = \alpha\pi_{t-1} + \beta y_{t-1} + \epsilon_t$, $y_t = \gamma y_{t-1} - \varphi(i_{t-1} - \pi_{t-1}) + \eta_t$, where π is the inflation rate, y is the output gap, i is the nominal interest rate, and ϵ, η are white noise shocks. Then, by imposing the rational expectations assumption, one obtains $E_t\pi_{t+1} = \varphi_1\pi_{t-1} + \varphi_2 y_{t-1} + \varphi_3(i_{t-1} - \pi_{t-1}) + \xi_{t+1}$, where $\varphi_1 \equiv \alpha^2$, $\varphi_2 \equiv (\alpha+\gamma)\beta$, $\varphi_3 \equiv -\beta\varphi$, and $\xi_{t+1} \equiv \alpha\epsilon_t + \beta\eta_t$. In general, the link between expected inflation and a variety of macro-variables may be interpreted as a perceived law of motion followed by the private sector under some form of learning (see e.g. Milani, 2007).

of Professional Forecasters.[3] Fig. 1 compares the 1-quarter ahead inflation expectations to the (1-period ahead) realized inflation. Evidently, the forecast error is quite persistent, a fact consistent with rational agents who must estimate the unknown time-varying inflation target (Erceg and Levin, 2003; Andolfatto et al., 2008) or must estimate some key-parameters of the model-economy of interest (Orphanides and Williams, 2005), or agents with bounded rationality.[4]

This evidence suggests the need of accounting for persistent factors – on top of realized inflation – for explaining inflation expectations.[5] One natural candidate is a measure of domestic slack, i.e. the output gap. In fact, given the role played by demand pressures in shaping an economy's price level, the evolution of the domestic business cycle may be informative for forecasting inflation. Additionally, we consider the 'G6 output gap', constructed by averaging the output gaps of Canada, Japan, Germany, France, Italy, and United Kingdom.[6] As documented by Tootell (1998), the countries belonging to the G7 have been the main trading partners of the U.S. economy after the second world war. Therefore, such a macro-aggregate is likely to provide an informative indicator of 'global slack'.[7] Borio and Filardo (2007) find evidence in favor of global measures of slack in Phillips-curve type models of inflation, so supporting the idea that the evolution of nominal prices might be driven by the world-wide business cycle.[8] Given the constantly increasing degree of openness of the U.S. economy in the last decades, the G6 output gap might have very well played a role as an indicator of future domestic inflation.

As far as globalization and global inflation are concerned, Ciccarelli and Mojon (in press) interpret a global dynamic factor computed with a panel of OECD countries as an 'inflation attractor' in an error-correction mechanism model for domestic inflation. Mumtaz and Surico (2008) find evidence in favor of a world factor significantly accounting for the decline in the level and persistence of national inflation rates. We approximate such a global indicator of inflation with the simple average of the inflation rates in the G6.[9]

[3] The Survey of Professional Forecasters, formerly conducted by the American Statistical Association and the National Bureau of Economic Research, is currently managed by the Federal Reserve Bank of Philadelphia. In this survey, forecasters are asked to provide quarterly projections up to five quarters ahead and annual projections for the current and following year on the main macroeconomic variables. For more information, see Croushore (1993).

[4] Interestingly, some of the features of the forecast error seems to be time-varying. For instance, its volatility (measured by its standard deviation) declines from 1.74 to 0.91 when moving from the sample 1970Q1–1984Q4 to 1985Q1–2006Q3. We will return on this sub-sample instability issue later.

[5] This evidence corroborates the idea of considering inflation expectations a different object with respect to realized inflation. While the relationship between these two objects as suggested by a microfounded new-Keynesian framework is very tight, we stress that such relationship is less tight when ingredients such as learning or lack of credibility by monetary policy authorities (leading to indeterminacy) are taken into account. On the latter point, see Castelnuovo and Surico (in press).

[6] We concentrate on G6 aggregates (i.e. we do not consider the U.S. output gap in building up our global indicators) to tackle multicollinearity. German data regard the sample 1991Q1–2006Q3 to remove the effect of reunification (see the Data Appendix for more details). We concentrate on the measures of the output gap provided by the OECD, which are computed on the basis of a measure of the potential output obtained by the production function approach. Our benchmark measure of the global gap is the simple average of the output gaps of the countries of interest. Crucini et al. (2008)show that there is a very high correlation between the simple average of G7 countries' real GDP growth rates and the latent world factor they compute (which is a weighted average of the G7 real GDP growth rates, with weights determined according to a statistical criterion). Our results are robust to the employment of a weighted average of the output gaps of the G6 (in the next Section).

[7] Notice that we do not take Mexico and China into account. As pointed out by Tootell (1998), data on inflation and the business cycle in Mexico and China (two among the most important exporters to the U.S. in the latter part of the sample) are not very reliable. Moreover, Bernanke (2007) estimates in about 0.1 percent per year the short-term effect reduction of the overall inflation rate due to the slowing increase in prices related to Chinese imports. Evidence against a large impact from the Chinese output gap to U.S. is also provided by Borio and Filardo (2007). Some regressions we conducted with a measure of the Chinese business cycle (the HP-filtered Chinese log-real GDP) confirmed us that such impact is not statistically significant in the model at hand. For further discussions on this issue, see Ball (2006).

[8] Notice that Ball (2006) disagrees on this point. In fact, he stresses the difference between the evolution of *nominal* prices over time (inflation) and the impact that a variation in openness of a country may have on the *relative* prices of the goods sold in that country. Bernanke (2007) underlines how increasing trade with China and other developing countries has led to a slower growth in the prices of imported manufactured goods. However, the development of such countries has also led to an increase in the prices for energy and commodities.

[9] The correlation between the global factor computed by Mumtaz and Surico (2008) - kindly provided us by Paolo Surico – and our measure of global inflation (both standardized) is 0.96. Ciccarelli and Mojon (in press) show that a simple average of the inflation rates in the OECD approximates well their global factor.

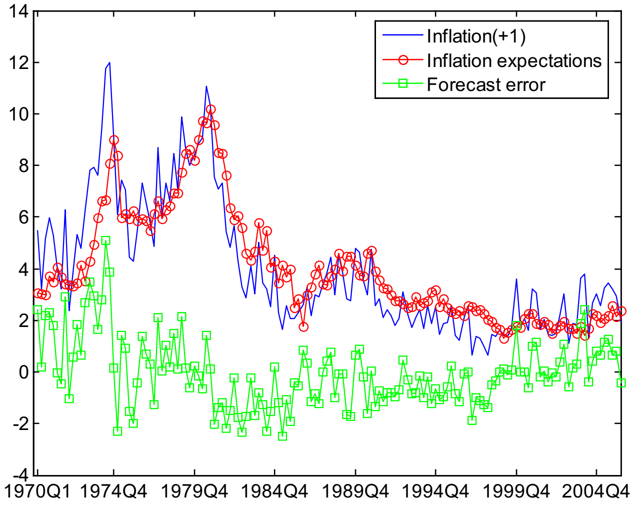

Fig. 1. Expected vs. realized inflation. Expected inflation: 1-qr. ahead expected GDP deflator inflation from the Survey of Professional Forecasters. Inflation forecast error computed as the difference between realized (1-qr. ahead) and expected inflation.

2.1. Modeling inflation expectations

We postulate the following encompassing model for inflation expectations:

$$\widetilde{E}_t\pi_{t+1} = c + \rho\widetilde{E}_{t-1}\pi_t + \alpha\widetilde{\pi}_{t-1} + \beta\widetilde{y}^D_{t-1} + \gamma\widetilde{y}^G_{t-1} + \delta\widetilde{x}_{t-1} + \epsilon_{t+1} \tag{1}$$

where $E_t\pi_{t+1}$ is the one quarter ahead annualized GDP inflation expectations as measured by the Survey of Professional Forecasters,[10] π_t is the annualized quarterly inflation as measured by the GDP deflator, y^D_t is a measure of the domestic output gap, y^G_t is the G6 log-real GDP, and x_t is a regressor taking into account other possible controls, among which G6 inflation. The wiggle above the variables indicates the *cyclical* component of the variables themselves, which is computed by applying the Hodrick–Prescott filter to the raw observations.[11] The motivation for this choice is threefold. First, all the variables considered in our study are non-stationary according to standard tests, and this could harm the reliability of our estimates.[12] Second, models of inflation expectations in the learning literature are often written in terms of deviations with respect to a known or perceived time-varying reference value (e.g. Orphanides and Williams, 2005). Third, removing a long-run trend from inflation expectations augments the ability of the model to capture the determinants of the short-run

[10] The Survey of Professional Forecasters' data refer to expected GDP deflator inflation. Consequently, the link between a global measure of the business cycle and the expected U.S. inflation we employ cannot regard imported *final* goods. However, the GDP deflator inflation is influenced by imported *inputs* employed in domestic production, which represent a considerable share of total U.S. imports. Moreover, the GDP deflator also accounts for the goods produced and exported by the U.S. economy, which are obviously influenced by the economic evolution of the foreign markets.

[11] We computed the cyclical component of output (both domestic and global) by considering its long-run trend (i.e. potential output) as computed by the OECD. In so doing, we created measures of output gaps whose economic interpretation is straightforward. As regards the remaining variables, each one was taken in deviation with respect to its Hodrick–Prescott trend (weight: 1600). A robustness check with one-sided filtered variables confirmed the robustness of our findings.

[12] We performed Augmented Dickey–Fuller and Phillips–Perron tests with a constant, a trend, and 4 lags. The null of unit root could never be rejected at conventional confidence levels. These results are available upon request.

fluctuations in inflation expectations, very much like removing trend inflation from realized inflation does (Cogley and Sbordone, 2008; Cogley et al., 2010).

Fig. 2 depicts the evolution of inflation expectations, our measure of time-varying inflation target, and the 'inflation expectations gap'. Notice that, while being a crude measure of the U.S. inflation target, the filtered series matches quite closely the estimates of the time-varying U.S. inflation target proposed by Ireland (2007). In fact, the average value of π_t^T in the subsample 1975Q1–1979Q4 is 7.50% vs. Ireland's (2007) 8%, and the average value in 2004 of 2.50% as Ireland's. Therefore, in estimating a model in gaps, we are already controlling (at least in first approximation) for the influence potentially exerted by the perceived inflation target on inflation expectations. Cogley and Sbordone (2008), Cogley et al. (2010) and Castelnuovo et al. (2008) also offer similar estimates of the low-frequency component of the U.S. inflation rate.

2.2. Filtering issues

Our filtering strategy has consequences as regards both the economic interpretation of our results and the issue of long-run relationships, i.e. cointegration. All variables entering our econometric model are expressed in *gaps*, i.e. in deviations with respect to their 'natural' values. In other words, we filter our raw variables so to obtain their cyclical component. This allows us to concentrate on short-run fluctuations of inflation expectations as driven by movements in possibly relevant drivers such as the domestic inflation gap, the domestic output gap, the global output gap, and so on. This cyclical representation leads to interpret the oscillations in the 'inflation expectations gap' in a very sensible manner from an economic standpoint. For instance, a short-run upward drift of such a gap could in principle be driven by positive realizations of the domestic 'inflation gap', and/or an expansionary phase of the domestic business cycle, and/or the international business cycle (of course, this is only one of the many possible combinations of positive and negative gaps that might occur). In contrast, economic theory suggests that, in the long run, all these variables should take values in line with their 'natural' rates, i.e. the economy should constantly remain in a situation of no macroeconomic stress of any sort, and the predicted inflation expectations *gap* should be zero.[13]

Of course, our model is empirically valid if our specification of the natural rates (trends) is 'correct'. Such rates are computed via statistical and economic filters (OECD estimates for output, Hodrick–Prescott filters for the remaining variables as for the baseline exercises; backward-looking moving averages as for our robustness check). How reliable are these filters? OECD estimates are based on a production function approach, which delivers widely accepted estimates. The Hodrick–Prescott filter is *de facto* the most employed filter to compute cyclical macroeconomic fluctuations. Interestingly, while being subject to possible critiques (e.g. end of sample issues), the Hodrick–Prescott filter delivers low-frequency representations of the objects of our interest (inflation expectations, inflation) that appear to be economically very sensible (recall the extremely high correlations between the computed low frequency representation of inflation expectations (Fig. 2) and those proposed by Ireland (2007), Cogley and Sbordone (2008), and Cogley et al, 2010).

Alternatively, one could in principle search for common trends (cointegration). Unfortunately, as recently pointed out by Müller and Watson (2009), standard inference in cointegrating models is fragile because it relies on the assumption of an $I(1)$ model for the common stochastic trends, which may not accurately describe the data's persistence. Moreover, such common trends might be difficult to estimate in a sample possibly featured by breaks in the systematic relationships linking the aggregates of interest, first and foremost breaks in monetary policy (Clarida et al., 2000; Lubik and Schorfheide, 2004; Cogley and Sargent, 2005; Boivin and Giannoni, 2006; Belaygorod and Dueker, 2007), which may cause breaks in reduced forms such as the one we deal with. Finally, the somewhat agnostic

[13] Of course, this is true up to the value assumed by the constant in our econometric models. Interestingly, the constant turns out to be insignificant in all our estimated models.

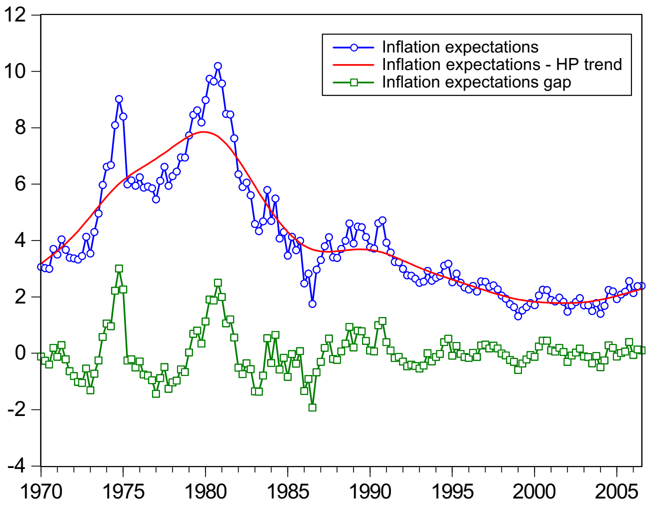

Fig. 2. Inflation expectations and trend inflation. Perceived inflation target computed as the Hodrick–Prescott filter of expected inflation (weight: 1600).

statistical approach we employ naturally allows for series-specific trends, a strategy possibly overcoming small sample biases otherwise arising when subsamples displaying high volatilities such as the 'great inflation' period of the 1970s are involved in the estimation. Therefore, while considering the cointegration issue as possibly relevant from a theoretical standpoint, these 'empirically based' considerations lead us to refrain from employing ECM-type of models in this context. We leave the analysis of this complex issue to future research.

2.3. Full-sample results

We run our regressions over the sample 1970Q1–2006Q3. Our regressors are lagged one-period in order to avoid (or at least milden) the endogeneity problems potentially arising when employing contemporaneous regressors. This allows us to estimate the model by Ordinary Least Squares.[14]

Tables 1 and 2 collect our benchmark results. Table 1 displays the estimates regarding different versions of the encompassing model (1), all considered under the constraint $\delta = 0$.[15] It is evident that the business cycle measures may be of help for tracking inflation expectations. Notably, when involving the domestic and the global output gap jointly in the estimation, the former loses its significance.[16] This finding is conceptually in line with the one proposed by Borio and Filardo (2007) in their study focusing on Phillips curves 'augmented' with the global slack, and it seems to point towards the relevance of globalization and global indicators in tracking inflation expectations.[17]

[14] This strategy is also followed by Cerisola and Gelos (2005), who estimate empirical models for tracking the Brazilian inflation expectations. Our results are robust to the employment of the GMM estimator (see the next Section).

[15] The estimated value for the constant c is not significant in all the regressions presented in the paper. This is a side-effect of considering variables in gaps.

[16] This result is robust to the employment of the U.S. and G7 output gap measures provided by the OECD (Main Economic Indicators). The significance of the global output gap measure is present also when $\tilde{y}^G_{t-2}$ (as opposed to $\tilde{y}^G_{t-1}$) is considered in the regression. When orthogonalizing the G6 output gap with respect to the domestic one, the orthogonalized G6 output is still significant in the regression.

[17] For a discussion on the robustness of Borio and Filardo's (2007) results to variations of the global output gap and inflation expectations proxies, see Ihrig et al. (2007).

Table 1
Model with output gap. Estimated model: Eq. (1) in the text. Sample: 1970Q1–2006Q3. The wiggle identifies variables in deviations with respect to their long-run stochastic trend, computed with the Hodrick–Prescott filter. The estimated constant is not reported because not significant. Significance level: ***/**/* = 99/95/90 per cent. Newey–West HAC VCV matrix (4 lags).

$\widetilde{E}_{t-1}\pi_t^D$	0.54*** (0.10)	0.55*** (0.09)	0.52*** (0.10)	0.53*** (0.10)
$\widetilde{\pi}_{t-1}^D$	0.21*** (0.05)	0.20*** (0.05)	0.19*** (0.04)	0.19*** (0.04)
$\widetilde{y}_{t-1}^D$		0.05*** (0.02)		0.02(0.02)
$\widetilde{y}_{t-1}^G$			0.09** (0.03)	0.08** (0.03)
$\sigma_{\widetilde{E_t\pi_{t+1}^D}}$	0.77	0.77	0.77	0.77
σ_ε	0.46	0.45	0.44	0.44
$\overline{R}^2$	0.64	0.66	0.68	0.67

Table 2 proposes the results of our regressions in which we allow for $\delta \neq 0$. As additional regressors, we consider the global inflation gap $\widetilde{\pi}_t^G$ as well as a more standard measure of inflationary pressures such as import price inflation $\widetilde{imppr\pi}_t$ (annualized quarterly growth rate of the import price index). To verify if the U.S. monetary policy stance has played any role in shaping inflation expectations, we follow Rudebusch and Svensson (1999) and Fuhrer and Rudebusch (2004) and consider the average real interest rate gap $\widetilde{avgrrate}$, computed via a backward-looking MA(3) transformation of the ex-post real interest rate, i.e. $avgrrate \equiv \sum_{i=0}^{3}(i_{t-i} - \pi_{t-i})/4$.[18]

A few results stand out. First, the point estimates of both expected inflation and realized inflation (both lagged) are stable across models and are highly significant. Second, our measures of external price pressures are significant. Third, the only model in which the global gap is not significant is the one in which global inflation is considered. This suggests that the global gap might actually be a proxy for the global inflation rate. In fact, when estimating a 'global Phillips curve' (sample: 1970Q1–2006Q3), we obtain

$$\widetilde{\pi}_{t+1}^G = \underset{0.11}{(0.37)}\,\widetilde{\pi}_t^G + \underset{(0.02)}{0.12}\,\widetilde{\pi}_{t-1}^G + \underset{(0.08)}{0.27}\,\widetilde{\widetilde{y}}_t^G + \widehat{\xi}_{t+1} \qquad \overline{R}^2 = 0.32, \sigma_{\widetilde{\pi}^G} = 1.62, \sigma_\xi = 1.34$$

which highlights the possible transmission channel going from external slack to external inflation, and eventually to domestic inflation expectations. Interestingly, Table 3 shows that the significance of the G6 inflation gap is robust to the inclusion of other measures of external inflationary pressures. Possibly, this is due to the presence of a global (common) component in the inflation process in OECD countries, at least when long-samples are investigated (Ciccarelli and Mojon, in press; Mumtaz and Surico, 2008). Notably, the average real interest rate gap never turns out to be significant.[19]

3. Parameter (in)stability and the role of monetary policy

Our full sample results point towards the role possibly played by global inflation in influencing inflation expectations. In contrast, the U.S. monetary policy conduct appears to have exerted a negligible impact. Taken at their face value, our full-sample estimates support the idea that G6 inflation has played a key-rule in shaping U.S. inflation, a finding supporting the globalization hypothesis put forward by several observers. However, these results rely on the assumption of stability of the estimated relationships over time. In fact, it turns out that this assumption is not warranted, and a much richer story may be told when inspecting the time-variability of the relationships at stake.

[18] We also considered as additional regressors different transformations (e.g. gaps, growth rates) of the unit labor costs, a possible proxy of marginal costs as in Sbordone (2002); a measure of trade openness proposed by Romer (1993), i.e. imports plus exports over total production; an indicator of financial stress (the S&P 500 index) as in Airaudo et al. (2006); an indicator of global liquidity (average of the money growth rate in the G7) as in D'Agostino and Surico (2009); an indicator of oil price inflation. Our results turned out to be robust to these controls.

[19] Notably, the insignificance of the monetary policy stance indicator is robust to the exclusion of the global inflation rate from the regression (as shown by the results in Table 2).

Table 2
Model with output gap, robustness checks. Estimated model: Eq. (1) in the text. Sample: 1970Q1–2006Q3. The wiggle identifies variables in deviations with respect to their long-run stochastic trend, computed with the Hodrick–Prescott filter. The estimated constant is not reported because not significant. Significance level: ***/**/* = 99/95/90 per cent. Newey–West HAC VCV matrix (4 lags).

$\widetilde{E}_{t-1}\pi_t^D$	0.53*** (0.10)	0.47*** (0.10)	0.48*** (0.09)	0.52*** (0.09)
$\widetilde{\pi}_{t-1}^D$	0.19*** (0.04)	0.16*** (0.04)	0.18*** (0.04)	0.20*** (0.04)
$\tilde{y}_{t-1}^D$	0.02(0.02)	0.04** (0.02)	0.00 (0.02)	0.02 (0.02)
$\tilde{y}_{t-1}^G$	0.08** (0.03)	0.04 (0.03)	0.06** (0.03)	0.07** (0.03)
$\widetilde{\pi}_{t-1}^G$		0.08** (0.03)		
$\widetilde{impprπ}_{t-1}$			0.02*** (0.00)	
$\widetilde{avgrrate}_{t-1}$				0.00(0.04)
$\sigma_{\widetilde{E_t\pi_{t+1}}}$	0.77	0.77	0.77	0.77
σ_ε	0.44	0.43	0.42	0.44
$\overline{R}^2$	0.67	0.69	0.69	0.67

3.1. Instability issue

To do so, we engage in the following exercise. We consider the model (1) with $\tilde{x}_{t-1} = \widetilde{\pi}_{t-1}^G$, and we estimate it over different (rolling) windows.[20] Fig. 3 depicts the evolution of the estimated parameters over the sample at hand. It is immediate to notice that the relevance of the U.S. output gap is not stable over the different subsamples we consider. Even more interestingly, the full-sample based finding regarding the relevance of global inflation is not supported by our rolling-window regressions. In fact, we observe a fall in global inflation's statistical relevance when approaching the second half of the sample. This is an interesting result, because it goes against the globalization hypothesis, which would actually suggest a *larger* relevance of global aggregates in the last two decades. We then add an extra-regressor, i.e. cyclical component of the real interest rate gap (lagged one period), and we repeat this exercise.

Fig. 4 reveals that the regressor that gains statistical relevance when during the 1980s and 1990s is the (proxy of the) Fed's monetary policy stance. The first windows confirm the insignificance of the real interest rate gap. In contrast, its significance clearly emerges when considering more recent observations, in particular those belonging to the Great Moderation.

3.2. Dating of the break: 'literature-based' approach

Our subsample regressions do not offer an exact dating of the break occurring in the relationships existing between global inflation and inflation expectations on the one hand, and monetary policy conduct and inflation expectations on the other. Nevertheless, there is an intriguing correlation between the drift in the relevance of the real interest rate for the dynamics of inflation expectations and the switch from 'passive' to 'active' monetary policy possibly occurred in the early 1980s. Clarida et al. (2000), Lubik and Schorfheide (2004), Cogley and Sargent (2005), Boivin and Giannoni (2006), and Belaygorod and Dueker (2007) find evidence in favor of a shift towards a more aggressive monetary conduct in correspondence of (or slightly later than) the appointment of Paul Volcker as Fed's chairman. Castelnuovo and Surico (in press) show that such shift is key to understand the gap between the dynamic responses estimated with a VAR in the pre-Volcker period and those produced with a microfounded new-Keynesian framework. This evidence enables us to identify a break date, i.e. 1982Q4.[21] Such a break-date is also supported by a standard Chow-breakpoint test, which rejects the

[20] Each window size is 72 observations (i.e. 18 years), a choice in line with Canova (2009). Our sample spans the years 1970–2006, so we decided to consider the first half of the sample as the very first window, and then roll over up to the end of the sample. For larger windows, our results are confirmed. When shortening the window width, our results in favor of a fading significance of global inflation are strengthened, while the results in favor of the significance of the monetary policy stance are less clear.

[21] Some researchers point towards 1984Q4 as a break date for explaining the change in the dynamics of the inflation process in the United States and other countries (e.g. McConnell and Perez-Quiros, 2000; Borio and Filardo, 2007). Our results are robust to setting to break-date to 1984Q4.

Table 3
Model with global inflation, robustness checks. Estimated model: Eq. (3) in the text. Sample: 1970Q1–2006Q3. The wiggle identifies variables in deviations with respect to their long-run stochastic trend, computed with the Hodrick–Prescott filter. The estimated constant is not reported because not significant. Significance level: ***/**/* = 99/95/90 per cent. Newey–West HAC VCV matrix (4 lags).

$\widetilde{E}_{t-1}\pi_t^D$	0.53*** (0.10)	0.47*** (0.09)	0.47*** (0.10)
$\widetilde{\pi}_{t-1}^D$	0.16*** (0.04)	0.17*** (0.04)	0.17*** (0.04)
$\widetilde{y}_{t-1}^D$	0.06*** (0.02)	0.04* (0.02)	0.06*** (0.02)
$\widetilde{\pi}_{t-1}^G$	0.09** (0.03)	0.07** (0.03)	0.10** (0.03)
$\widetilde{imppr\pi}_{t-1}$		0.01** (0.00)	
$\widetilde{avgrrate}_{t-1}$			−0.01(0.02)
$\sigma_{\widetilde{E_t\pi_{t+1}}}$	0.77	0.77	0.77
σ_ε	0.43	0.42	0.43
$\overline{R}^2$	0.68	0.70	0.69

null hypothesis of parameter instability in the quarter 1982Q4 (p-value: 0.02). We then re-estimate the model with global inflation and the average real interest rate by concentrating on two different subsamples, i.e. 1970Q1–1982Q3 and 1982Q4–2006Q3.

3.3. Sub-sample results

Table 4 collects our subsample estimates, which confirm that the significance of global inflation weakens in the second part of the sample. The opposite holds as far as the monetary policy stance is concerned.[22] This result squares with the findings recently proposed by Boivin and Giannoni (2010), who estimate a Factor-Augmented VAR on a large set of U.S. and international macroeconomic series and find no support for global forces as elements affecting the monetary policy transmission mechanism in the sample 1984–1999. This evidence does not appear to be due to an endogeneity problem. In fact, GMM regressions (instruments: constant and four lags of inflation expectations gap, U.S. inflation gap, G6 inflation gap, monetary policy stance indicator gap) implies point estimates in line with those obtained with OLS. Table 5 shows our GMM-based results, along with the p-values of the J-test confirming the goodness of the selected instruments.

3.4. Trade openness and the role of global factors

At a first glance, one may find the vanishing role of global indicators counter-intuitive. After all, the *level* of the U.S. trade openness has increased in the last two decades. However, when referring to inflation expectations – which is a measure capturing the expected *growth* rate of the price level – it is of interest to look at the *increase* of the trade openness (Ball, 2006). In fact, the mean growth rate of the U.S. trade openness in the first subsample is 4.22%, while in the second subsample is 1.92%, a markedly lower value. Fig. 5 depicts the low-frequency component of the U.S. trade openness. Evidently, the acceleration of the U.S. trade exchanges took place in the 1970s, then fell at the end of that decade, partly recovered in the 1980s and 1990s, and fell once more at the end of the last century, eventually upsurging since the beginning of the current decade. If trade openness had been one of the main drivers of the reduction in mean and variance of inflation and inflation expectations experienced in the U.S. in the last two decades, we should have probably observed a different pattern of its growth rate.[23]

[22] It is worth noticing that the correlation between global inflation and the real interest rate (both in gaps) is 0.46 in the sample 1970Q1–1984Q4, 0.44 in the next 10 years, and −0.10 in the sample 1995Q1–2006Q3. This confirms that while in the first part of the sample the information coming from global inflation renders that stemming from the monetary policy stance superfluous, things change in the second part of the sample. For similar findings, see Boivin and Giannoni (2010).

[23] Further comments along this line can be found in Ball (2006).

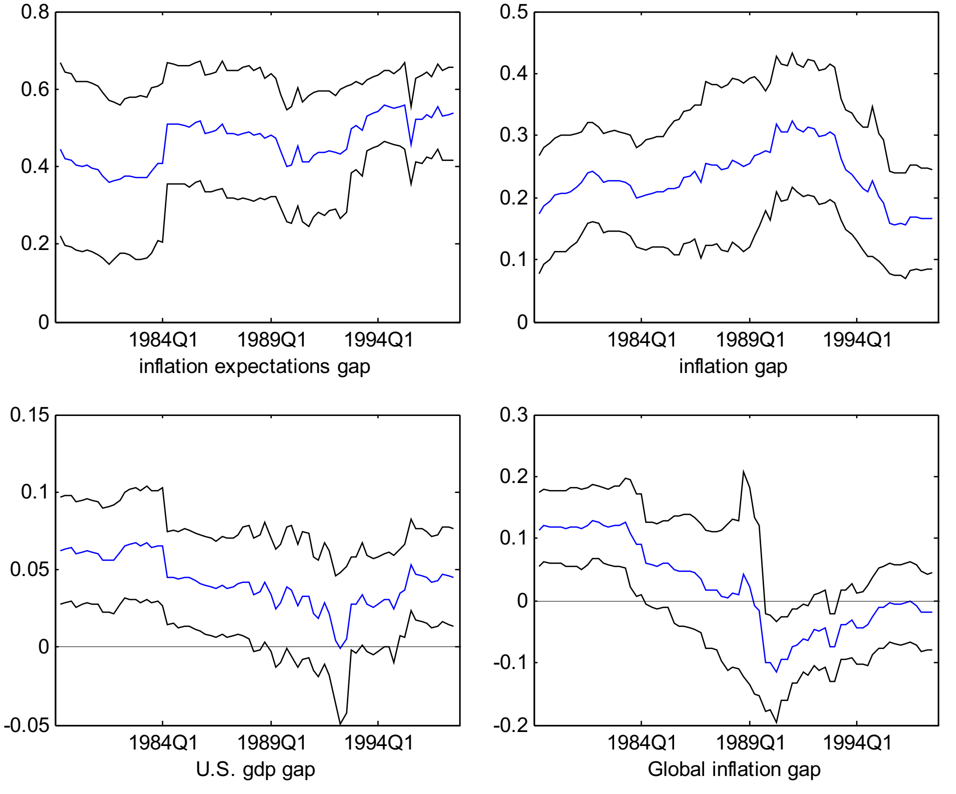

Fig. 3. Model with global inflation: rolling estimates. Label on the *x*-axis: Median observation of a given window. Rolling window size: 72 quarters. Solid line: Mean point estimate; dotted lines: 90% confidence bands. Newey–West robust standard errors (computing by allowing up to 4 lags in the autocorrelation of the estimated errors).

Our results are at odds with those proposed by Borio and Filardo (2007). They estimate Phillips curves for 16 OECD countries in the sample 1985–2005 and find that global output gap measures overwhelm domestic output gaps in affecting domestic inflation. We offer two explanations for justifying the different results we find. First, we deal with a different object, i.e. inflation expectations (as opposed to realized inflation). Second, we include the lagged dependent variable among the regressors. In fact, when we omit lagged inflation expectations from our regressions for the sample 1985Q1–2006Q3, we find a point estimate of 0.14 for the measure of global slack, significant at a 95% level. However, the model is clearly misspecified: the Breusch–Godfrey serial correlation LM test (run with 2 lags) rejects the null of independent residuals at the 99% level. When adding lagged inflation expectations, the point estimate of the global slack regressor lowers to 0.06, and it is not significant at standard confidence levels anymore.[24]

As already commented, we find support for global indicators as drivers of the inflation expectations during the 1970s. This result may appear at odds with the one proposed by Ihrig et al. (2007), who find no evidence for international demand pressures in Phillips curves estimated over the sample 1977–2005. Apart from the different objects under investigation (realized inflation in Ihrig et al., 2007, as opposed to expected inflation in this paper), subsample selection is probably the reason underlying our contrasting results. As already pointed out, when estimating Eq. (1) (with

[24] The statistical relevance of the global slack in Phillips curves seems to depend on the way a researcher builds up the measure of global pressure. See Ihrig et al. (2007) for a detailed discussion.

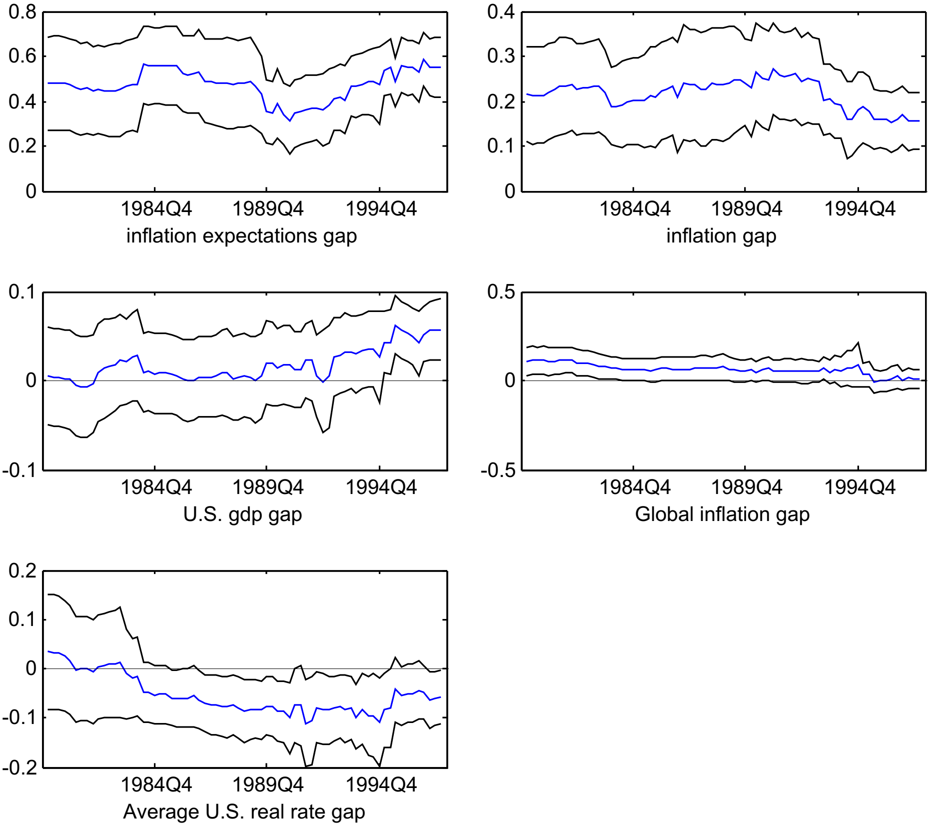

Fig. 4. Model with global inflation and real interest rate: rolling windows. Label on the *x*-axis: Median observation of a given window. Rolling window size: 72 quarters. Solid line: Mean point estimate; dotted lines: 90% confidence bands. Newey–West robust standard errors (computing by allowing up to 4 lags in the autocorrelation of the estimated errors).

$\tilde{x}_{t-1} = \widetilde{avgrrate}_{t-1}$) for the sample 1970Q1–2005Q4, we support the presence of global output in the inflation expectations equation. Interestingly, when re-estimating the very same model for the 1977Q1–2006Q3 subsample (i.e. by taking 1977 as first year of the investigated sample as in Ihrig et al., 2007), the *p*-value of the global slack regressor goes from 0.05 up to 0.12, i.e. the global gap is not significant at standard confidence levels anymore. Interestingly, the dominance of domestic factors over global indicators is also supported by Milani (2009), who estimates an open-economy DSGE model for the U.S. economy and finds just a mild evidence for the global output gap as a possible driver of U.S. inflation.

3.5. An interpretation of our results

Our evidence supports the idea of a 'substitution' between global inflation and the U.S. monetary policy stance as determinants of U.S. inflation expectations. We propose the following monetary policy-related interpretation. Forecasters exploit available information to predict the evolution of the object of interest. During the 1970s, a big oil shock severely hit the U.S. economy. The reaction of the Fed was *de facto* inflationary, with the real interest rate recording also negative realizations for several quarters. In fact, the 1970s were featured by high levels of inflation and inflation volatility. Given the weak link existing between real interest rates and inflation, forecasters might have looked at international indicators (other than those related to domestic variables) to refine their forecasts. High

Table 4
Expected inflation: model with global inflation, subsample comparison. Estimated model: eq. (3) in the text. First sub-sample: 1970Q1–1982Q3. Second sub-sample: 1982Q4–2006Q3. The wiggle identifies variables in deviations with respect to their long-run stochastic trend, computed with the Hodrick–Prescott filter. The estimated constant is not reported because not significant. Significance level: ***/**/* = 99/95/90 per cent. Newey–West HAC VCV matrix (4 lags).

	1st sub.	2nd sub.
$\widetilde{E}_{t-1}\pi_t^D$	0.57*** (0.15)	0.44*** (0.09)
$\widetilde{\pi}_{t-1}^D$	0.11* (0.06)	0.24*** (0.06)
$\widetilde{y}_{t-1}$	0.09*** (0.03)	0.04* (0.02)
$\widetilde{\pi}_{t-1}^G$	0.11** (0.04)	−0.03(0.03)
$avgrrate_{t-1}$	0.06(0.07)	−0.07**(0.04)
$\sigma_{\widetilde{E_t\pi_{t+1}}}$	1.16	0.49
σ_ε	0.54	0.36
$\overline{R}^2$	0.79	0.47

inflation rates in the G6 might have influenced the forecasters both because of the transmission of inflation via tradeables, and via the observed comovements in the inflation rates at an international level.[25] With the end of the Volcker-experiment, the Fed became more aggressive against inflation fluctuations and its credibility increased.[26] This is possibly one of the reasons why forecasters might have raised their attention on the U.S. monetary policy conduct. Interestingly, Ihrig et al. (2007) find that the link between international economic conditions and domestic inflation in 11 industrialized countries is tenuous at best, and attribute this 'missing link' to the improved monetary policy in the countries they analyze. They argue that a better monetary policy might have anchored inflation expectations and stabilized inflation, so rendering it less sensitive to resource utilization and relative prices and potentially offsetting the impact of globalization. Our empirical findings clearly corroborate Ihrig et al.'s (2007) conjecture.

4. Further investigations

We perform some further investigations to assess the robustness of our results as well as the suitability of the break-point our subsample analysis hinges upon.[27]

- *Higher lag structure.* Our baseline models for inflation expectations involve only one lag for the 'dependent' as well as 'independent' variables. While leading to a parsimonious and somewhat more easily interpretable model (at least from an economic standpoint), such an ad hoc lag selection may bias our results, especially if it takes several quarters for monetary policy to influence inflation expectations. To tackle this issue, we first estimate models with additional lags of the dependent variable (two and four lags), and find our qualitative conclusions unchanged. Then, we consider a richer lag structure as for the 'independent' regressors. To select a suited lag structure, we proceed as follows. We start off with four lags for each independent variables, with the exception of our measure of real interest rate, which already involves four lags due to its moving-average representation. For each regressor, we then drop the highest lag (e.g. the fourth lag in the initial round) if its associated p-value assumes a value higher than 0.10, re-estimate the (at that point constrained) model, and re-check the p-values of the highest lag of each variable. The algorithm stops when all the

[25] We recall the already cited contributions by Ciccarelli and Mojon (in press) and Mumtaz and Surico (2008) regarding the comovements in the inflation rates in the OECD.

[26] Schaumburg and Tambalotti (2007) build up a framework for analyzing a continuum of monetary policy rules featured by different degrees of credibility, in which commitment and discretion are special cases of what they call "quasi (i.e. imperfect) commitment" regime. In short, each period a central bank formulates optimal commitment plans, but it faces an exogenous probability of reneging its promises. This probability is interpreted as a measure of lack of credibility. Hakan Kara (2007) estimates this credibility parameter, and finds evidence in favor of an increase of the Fed's credibility when moving over the Volcker regime.

[27] We thank an anonymous referee for stimulating us to write this Section of the paper.

Table 5
Robustness check: GMM estimates. Estimated model: eq. (3) in the text. First sub-sample: 1970Q1–1982Q3. Second sub-sample: 1982Q4–2006Q3. Instruments: constant, first four lags of inflation expectations gap, domestic inflation gap, global inflation gap, average domest real interest rate gap. The wiggle identifies variables in deviations with respect to their long-run stochastic trend, computed with the Hodrick–Prescott filter. The estimated constant is not reported because not significant. Significance level: ***/**/* = 99/95/90 per cent.

	1st sub.	2nd sub.
$\widetilde{E}_{t-1}\pi_t^D$	0.68*** (0.06)	0.42***(0.07)
$\widetilde{\pi}_{t-1}^D$	0.09** (0.03)	0.09*** (0.03)
$\widetilde{y}_{t-1}^D$	0.08*** (0.02)	0.11*** (0.03)
$\widetilde{\pi}_{t-1}^G$	0.13*** (0.02)	−0.01(0.03)
$avgrrate_{t-1}$	0.01 (0.03)	−0.09*** (0.03)
$\sigma_{\widetilde{E}_t\pi_{t+1}}$	1.19	0.49
σ_{ξ}	0.59	0.38
$\overline{R}^2$	0.75	0.39
J-stat, p-value	0.83	0.55

estimated parameters of the highest lags take p-values ≤ 0.10. Consistently with our previous estimates, we impose a minimum structure of one lag per each variable. As regards the 'great inflation' period 1970Q1–1982Q3, this algorithm leads us to retain a model with one lag for the U.S. inflation, one lag for the U.S. output gap, and two lags for the global inflation rate (on top of the remaining domestic variables). The estimated coefficient associated to the U.S. real interest rate reads 0.04, with a p-value equal to 0.51; a Wald test investigating the possible exclusion of the global inflation measure clearly reject such null hypothesis (p-value: 0.002). As for the post-Volcker sample 1982Q4–2006Q3, we select a model with three lags for the U.S. inflation and output gap, and two lags for the global inflation measure. In line with our baseline estimates, the significance of global inflation vanishes, with the Wald test suggesting its exclusion (p-value: 0.15). As far as the U.S. monetary policy stance is concerned, the value of the estimated coefficient reads −0.06, which is in line with our baseline estimate. However, its p-value is equal to 0.11. Nevertheless, when scrutinizing the sample 1984Q1–2006Q3 (model with three lags for the U.S. inflation, one lag for the U.S. output gap, and two lags for the global inflation rate), the estimated value for this coefficient becomes −0.09, with a p-value equal to 0.02. Interestingly, this last model suggests the inclusion of the global inflation measure (p-value associated to the Wald test: 0.05).

- *Choice of the break-point: The Bai-Perron tests.* In selecting the break-point for our baseline regressions, we appealed to some external information related to a possible monetary policy drift occurred in the U.S. in the early 1980s. As pointed out, a standard Chow-test supports this choice. However, different reduced form models may be associated to different break-dates, both in terms of number of breaks and as far as the temporal 'location' of such breaks is concerned. Bai and Perron (1998, 2003) develop a battery of tests (to identify breaks) that do not require any a-priori 'external' information. We then employ Bai and Perron's algorithms and re-estimate our model with global inflation and the U.S. ex-post real interest rate to validate our choice of a single break in 1982Q3.[28] Interestingly, our choice turns out to be largely supported by the data as processes by Bai and Perron's tests. The Double Maximum Tests clearly reject the null hypothesis of stability of the model in the full sample in favor of the alternative of $m > 0$ unspecified breaks. In particular, the UD max statistic reads 21.63, while the WD max ones read 24.59 (10%), 25.45 (5%), 27.37 (1%), all figures remarkably larger than the relevant critical values. As for the number of breaks, the $SupF_T(2|1) =$

[28] We employed Bai and Perron's GAUSS code, which has been recently amended to overcome the drawbacks signalled by Kleiber and Zeileis (2005). The code is available at http://people.bu.edu/perron/code.html. Following the suggestions by Bai and Perron (2003), we considered a maximum number of breaks equal to 5, a percentage trimming of 15%, a minimal length of the segment equal to 21, a number of regressors whose estimated parameters are allowed to change $q = 2$ (global inflation, the U.S. policy stance), and we allowed for heterogeneity and autocorrelation in the residuals, an AR(1) prewhitening prior to estimating the long-run covariance matrix, as well as the possibility of sample-specific variances of the residuals. Our results are robust to the assumptions of homogeneity and serial uncorrelation of the econometric error term.

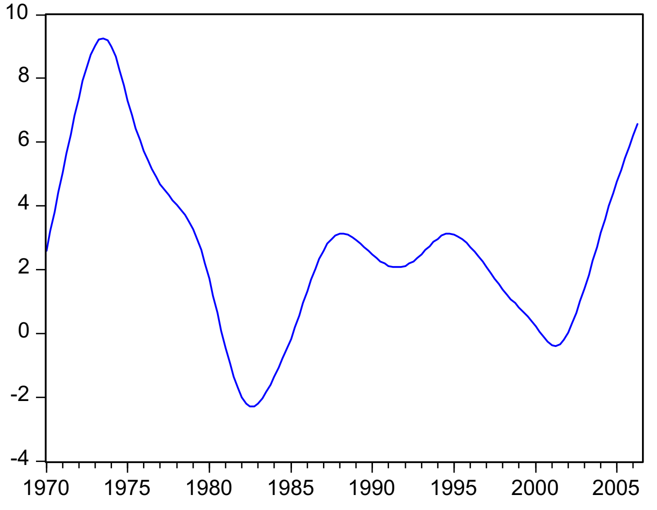

Fig. 5. Trade openness growth rate: long-run trend. Trend: Hodrick–Prescott filter (weight: 1.600) of the annualized quarterly growth rate of the U.S. trade openness. Trade openness computed as (Exports plus Imports)/GDP, as in Romer (1993).

7.02, the $SupF_T(3|2) = 0.55$, the $SupF_T(4|3) = 2.19$, and the $SupF_T(5|4) = 0.00$, all figures smaller than the critical values, i.e. the null of a single break is *not* rejected by the data. Consistently, both the BIC and the sequential method select a single break in 1981Q3, with the 90% confidence interval estimated to be [1979Q4, 1984Q4]. This interval clearly contains the break-date we picked up for our baseline regressions, i.e. 1982Q3. Moreover, when conducting subsample regressions by using 1981Q3 as a reference break-date, we verify the robustness of our conclusions, with the exception of the U.S. real interest rate in the sample 1970Q1–1981Q2, which is significant (but takes the wrong sign). This further battery of exercises, not shown for the sake of brevity, is available upon request.

- *Different data transformation.* Finally, we employ an alternative measure of G6 output gap constructed as the weighted average of the country-specific gaps (the weights being the shares of each country's real GDP on the G6 overall real GDP). We then detrend the variables of interest with an alternative filter (a one-sided backward looking MA(3)). We also consider undetrended measures of inflation. All these checks (not shown in the paper for the sake of brevity, but available upon request) confirm the robustness of our findings.

5. Conclusions

This paper estimates several different reduced form models for tracking the U.S. inflation expectations. We considered both domestic macroeconomic drivers and potentially relevant global drivers such as global inflation or the global business cycle. We engaged both in full sample analysis and in subsample-based investigations to assess the role played by global inflation and the Fed's monetary policy in influencing U.S. inflation expectations.

Our main findings point towards the instability of the estimated parameters of our empirical models. In particular, global indicators appear to have played a significant role until the mid-1980s, but they have subsequently been 'replaced' by the U.S. monetary policy stance as one of the main drivers of U.S. inflation expectations. Our interpretation of this finding points towards the enhanced credibility that the Fed began to gain after the end of the Volcker experiment.

In a recent paper, Borio and Filardo (2007) propose to move from standard 'domestic-centric' models to a novel 'global-centric' paradigm acknowledging the role potentially played by global indicators in shaping a country's inflation rate. We agree with Borio and Filardo (2007) on the importance of carefully monitoring the evolution of the impact that external pressures may exert on the domestic inflation process. However, our results suggest that U.S. inflation expectations have mainly been driven by domestic factors in a period in which the Fed has been aggressive enough to tackle the effects of macroeconomic shocks. Even if globalization is a fact, our evidence still supports the employment of 'domestic-centric' models for policy analysis, a view forcefully re-proposed by Kohn (2006) and Mishkin (2009), and theoretically supported by Woodford (2007).

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Data appendix

The sources and the treatment of the data employed in this paper are the followings:

1-quarter ahead inflation forecasts: quarter-by-quarter (annualized) GDP inflation forecasts (median values) taken from the Survey of Professional Forecasters, Federal Reserve Bank of Philadelphia. Note: During the '70s and the '80s, the Survey of Professional Forecasters was formulated in terms of GNP price deflator. However, over this period, the one-quarter annualized inflation rate computed using the GNP price deflator is extremely close to the one obtained with the GDP price deflator.

OECD U.S. output gap: Computed by the OECD as the percentualized log-deviation of the U.S. log real GDP (volume, base year: 2000) with respect to a measure of potential output. Potential output is based on a production function approach, taking into account the capital stock, changes in labor supply, factor productivities and underlying "non-accelerating inflation rates of unemployment" (Nairu) for each Member country.

G6 output gap: Simple average of the OECD output gaps of Canada, Japan, Germany, France, Italy, and United Kingdom.

U.S. inflation rate: quarter-by-quarter (annualized) changes of the U.S. GDP deflator at market prices (base year: 2000), taken from the OECD Main Economic Indicators.

G6 inflation rate: Simple average of the OECD GDP deflator inflation rates of Canada, Japan, Germany, France, Italy, and United Kingdom.

U.S. import price inflation rate: quarter-by-quarter (annualized) changes of the import price index (base year: 2000), taken from the OECD Main Economic Indicators.

U.S. oil price inflation rate: quarter-by-quarter (annualized) changes of the spot oil price (dollars per barrel) West Texas Intermediate (Dow Jones & Company), provided by the Wall Street Journal and downloaded from the Federal Reserve Bank of St. Louis's website, averages of monthly observations.

U.S. Unit labor costs: Worker compensation and benefits per unit of manufactured output, OECD Main Economic Indicators.

S&P 500: Index as reported by http://finance.yahoo.com, average of monthly observations.

U.S. Short-term nominal interest rate: Federal Funds Rate taken from the Federal Reserve Bank of St. Louis, average of monthly observations.

U.S. Long-term nominal interest rate: 10-Year Treasury Constant Maturity Rate taken from the Federal Reserve Bank of St. Louis, average of monthly observations.

Trade openness: Total Import plus Total Exports of Goods and Services, percentage of GDP, taken from the Federal Reserve Bank of St. Louis.

U.S. nominal effective exchange rate: Exchange rate of the U.S. dollar vis-à-vis other currencies weighted by their share in the U.S. international trade.

U.S. Surplus/GDP: Primary Government balance, percentage of GDP, taken from the OECD Main Economic Indicators.

G7 money growth rate: Simple average of the growth rates of the money stock in the G7 countries. For the exact definition of the money stocks employed, see D'Agostino and Surico (2009).

All the series employed in this paper are seasonally adjusted where applicable. All the gaps employed in this paper are computed as deviations (or log-deviations) of a given variable from its Hodrick–Prescott filter (weight: 1600) where not differently specified. German data regard the sample 1991Q1–2006Q3 to remove the effect of reunification.

References

Airaudo, M., Nisticò, S., Zanna, L.F., 2006. Learning, Monetary Policy, and Asset Prices. Mimeo.
Andolfatto, D., Hendry, S., Moran, K., 2008. Are inflation expectations rational? Journal of Monetary Economics 55 (2), 406–422.
Bai, J., Perron, P., 1998. Estimating and testing linear models with multiple structural changes. Econometrica 66, 47–78.
Bai, J., Perron, P., 2003. Computation and analysis of multiple structural change models. Journal of Applied Econometrics 18 (1), 1–22.
Ball, L., 2006. Has Globalization Changed Inflation? National Bureau of Economic Research Working Paper No. 12687.
Belaygorod, A., Dueker, M., 2007. The Price Puzzle and Indeterminacy. Federal Reserve Bank of St. Louis Working Paper 2006-2025.
Bernanke, B.S., March 2, 2007. Globalization and monetary policy. In: Remarks Given at the Fourth Economic Summit. Stanford Institute for Economic Policy Research, Stanford, California.
Boivin, J., Giannoni, M.P., 2006. Has monetary policy become more effective? Review of Economics and Statistics 88 (3), 445–462.
Boivin, J., Giannoni, M.P., 2010. Global forces and monetary policy effectiveness. In: Galí, J., Gertler, M. (Eds.), International Dimensions of Monetary Policy. University of Chicago Press.
Borio, C., Filardo, A., 2007. Globalisation and inflation: New cross-country evidence on the global determinants of domestic inflation. Bank of International Settlements Working Paper No. 277.
Canova, F., 2009. Monetary policy and the evolution of the U.S. economy. Journal of the European Economic Association 7 (4), 1–25.
Castelnuovo, E., Surico, P. Monetary policy, inflation expectations and the price Puzzle. Economic Journal, in press.
Castelnuovo, E., Greco, L., Raggi, D., 2008. Estimating Regime-Switching Taylor Rules with Trend Inflation. Bank of Finland Discussion Paper No. 20-2008.
Castelnuovo, E., Nicoletti Altimari, S., Rodrìguez-Palenzuela, D., November 2003. Definition of price stability, range and point inflation targets: the anchoring of long-term inflation expectations. In: Issing, O. (Ed.), Background Studies for the ECB's Evaluation of Its Monetary Policy Strategy. European Central Bank Publisher, pp. 43–90.
Cerisola, M., Gelos R.G., 2005. What drives inflation expectations in Brazil? An empirical analysis. IMF Working Paper No. 05-109.
Ciccarelli, M., Mojon, B. Global Inflation. The Review of Economics and Statistics, in press.
Clarida, R., Galí, J., Gertler, M., December 1999. The science of monetary policy: a New Keynesian Perspective. Journal of Economic Literature XXXVII, 1661–1707.
Clarida, R., Galí, J., Gertler, M., 2000. Monetary policy rules and macroeconomic stability: evidence and some theory. The Quarterly Journal of Economics 115 (1), 147–180.
Cogley, T., Sargent, T.J., April 2005. Drift and volatilities: monetary policies and outcomes in the post WWII US. Review of Economic Dynamics 8 (2), 262–302.
Cogley, T., Primiceri, G., Sargent, T.J., 2010. Inflation-gap persistence in the U.S. American Economic Journal: Macroeconomics 2 (1), 43–69.
Cogley, T., Sbordone, A., 2008. Trend inflation, Indexation, and inflation persistence in the new-Keynesian Phillips curve. The American Economic Review 98 (5), 2101–2126.
Croushore, D., November/December 1993. Introducing: the survey of professional forecasters. Business Review, Federal Reserve Bank of Philadelphia, 3–13.
Crucini, M.J., Kose, A., Otrok C., 2008. What are the driving forces of international business cycles? NBER Working Paper No. 14380, October.
D'Agostino, A., Surico, P., 2009. Does global liquidity help to forecast U.S. inflation? Journal of Money, Credit and Banking 41, 479–489.
Erceg, C.J., Levin, A.T., 2003. Imperfect credibility and inflation persistence. Journal of Monetary Economics 50, 915–944.
Fuhrer, J.C., Rudebusch, G., September 2004. Estimating the Euler equation for output. Journal of Monetary Economics 51 (6), 1133–1153.
Gürkaynak, R., Sack, B., Swanson, E., 2005. The sensitivity of long-term interest rates to economic news: evidence and implications for macroeconomic models. The American Economic Review 95 (1), 425–436.
Gürkaynak, R., Sack, B., Swanson, E., 2006. Does inflation targeting anchor long-run inflation expectations? Evidence from long-term bond yields in the U.S., U.K., and Sweden. Federal Reserve Bank of San Francisco Working Paper No. 2006-09.
Ihrig, J., Kamin, S.B., Lindner, D., Marquez, J., 2007. Some simple tests of the globalization and inflation hypothesis. Board of Governors of the Federal Reserve System, International Finance Discussion Paper No. 891, April.
Ireland, P., 2007. Changes in the Federal Reserve's inflation target: causes and consequences. Journal of Money, Credit, and Banking 39 (8), 1851–1882.
Kara, H., 2007. Monetary policy under imperfect commitment: reconciling theory and evidence. International Journal of Central Banking 3 (1), 149–177.
Kleiber, C., Zeileis, A., 2005. Validating multiple structural change models: a case study. Journal of Applied Econometrics 20 (5), 685–690.
Kohn, D.L., 2006. The Effects of Globalization on Inflation and Their Implications for Monetary Policy. Remarks at the Federal Reserve Bank of Boston's 51th Economic Conference, Chatham, Massachusetts, June 16.
Lubik, T.A., Schorfheide, F., 2004. Testing for indeterminacy: an application to US monetary policy. The American Economic Review 94 (1), 190–217.

McConnell, M., Perez-Quiros, G., 2000. Output fluctuations in the United States: what has changed since the early 1980's? The American Economic Review 90 (5), 1464–1476.
Milani, F., 2007. Expectations, learning and macroeconomic persistence. Journal of Monetary Economics 54 (7), 2065–2082.
Milani, F., 2009. Has Globalization Transformed U.S. Macroeconomic Dynamics? University of California at Irvine, mimeo.
Mishkin, F.S., 2009. Globalization, Macroeconomic Performance, and Monetary Policy. Journal of Money, Credit, and Banking 41 (s1), 187–196.
Müller, U.K., Watson, M., 2009. Low-frequency Robust Cointegration Testing. Princeton University, Mimeo.
Mumtaz, H., Surico, P., 2008. Evolving International Inflation Dynamics: World and Country Specific Factors. CEPR Working Paper No 6767.
Orphanides, A., Williams, J.C., 2005. Imperfect Knowledge, inflation expectations and monetary policy. In: Bernanke, B., Woodford, M. (Eds.), The Inflation Targeting Debate. University of Chicago Press.
Romer, D., 1993. Openness and inflation: theory and evidence. The Quarterly Journal of Economics 108 (4), 869–903.
Rogoff, K., 2003. Globalization and Global Disinflation. Paper presented at the symposium "Monetary policy and uncertainty. Adapting to a Changing Economy", Federal Reserve Bank of Kansas City, August 28–30.
Rudebusch, G.D., Svensson, L.E.O., 1999. Policy rules for inflation targeting. In: Taylor, John B. (Ed.), Monetary Policy Rules. University of Chicago Press.
Schaumburg, E., Tambalotti, A., 2007. An investigation of the gains from commitment in monetary policy. Journal of Monetary Economics 54 (2), 302–324.
Sbordone, A., March 2002. Prices and unit labor costs: a new test of price stickiness. Journal of Monetary Economics 49 (2), 265–292.
Tootell, G., July/August 1998. Globalization and U.S. inflation. Federal Reserve Bank of Boston New England Economic Review.
Woodford, M., 2003. Interest and prices: foundations of a theory of monetary policy. Princeton University Press.
Woodford, M., 2007. Globalization and Monetary Control. NBER Working Paper No. 13329, August.