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Estimating a high-frequency New-Keynesian Phillips curve

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Abstract This paper estimates a high-frequency New-Keynesian Phillips curve via the generalized method of moments. Allowing for higher-than-usual frequencies strongly mitigates the problems of small-sample bias and structural breaks. Applying a daily frequency allows us to obtain estimates for the Calvo parameter of nominal rigidity over a very short period—for instance for the recent financial and economic crisis—which can then be easily transformed into their low-frequency equivalences. With Argentine data from the end of 2007 to the beginning of 2011 we estimate the daily Calvo parameter and find that on average prices remain fixed for approximately two to three months which is in line with recent microeconomic evidence.

Keywords Calvo staggering · High-frequency New-Keynesian model · GMM

JEL Classifications C26 · E31

1 Introduction

Since the late 1990's the estimation of the New-Keynesian Phillips curve (NKPC for short) derived from Calvo (1983) staggered pricing has been prominent in the macroeconomic literature. Economists have been interested in the accuracy of the NKPC to resemble real time data and the information on the structural parameters, especially

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the Calvo parameter of nominal rigidity. A major drawback in this analysis is the low frequency. Usually, a NKPC is estimated applying Hansen's (1982) generalized method of moments (GMM) to quarterly observations. As has been shown by Fuhrer et al. (1995), however, GMM suffers from a small sample bias with the consequence that this method demands a critical amount of observations to achieve reliable estimates. Lindé (2005) argues that it takes approximately 1,000 observations for GMM to converge to the true values when estimating a NKPC. In order to obtain that many observations in a quarterly setup—with only four observations per year—a time span of 250 years is necessary. However, going back even only 50 or 100 years covers many different periods with drastically changing economic conditions such as the Great Depression, the high inflation period after the oil price shocks, the low inflation period commonly referred to as the Great Moderation and the current financial and economic crisis, i.e., the so-called Great Recession. Assuming that behavioral deep parameters remain constant across such different periods which are characterized by many structural breaks is certainly implausible.[1]

The contribution of this paper is to allow for the estimation of the Calvo parameter at a much higher frequency and thereby reducing the risk of small sample bias and structural breaks imminently. To account for such problems, we apply the standard GMM approach to estimate the NKPC for a higher frequency. Subsequently, we are able to transform the estimation results for the high-frequency Calvo parameter into pseudo lower-frequency equivalences by simply applying the rules to derive a high-frequency NKPC, which are described in Franke and Sacht (forthcoming); i.e., we account for the fact that the frequency-dependent parameters of the model should be suitably adjusted. In particular we use daily Argentine inflation data ranging from the end of 2007 to the beginning of 2011 provided by the Billion Prices Project at MIT Sloan. Doing so, our data comprises only 3 years of observations, but contains 830 data points. Our high-frequency approach gives us several advantages over the standard quarterly analysis of the NKPC. First, we are able to focus on specific events such as the financial and economic crisis and second, we can estimate the respective Calvo parameter more accurately due to a large amount of observations. However, our approach is not restricted to the use of daily data. Daily data serves mainly as an extreme example for possible applications of our approach.

In this paper we show that the corresponding estimates in daily frequency can be transformed into their lower-frequency equivalences and (in future research) can be used to calibrate business cycle models on a monthly or quarterly frequency. Therefore, we claim that this approach is helpful to get new insights into firms' pricing schemes. Furthermore, we confront our results with microeconometric evidence from the literature.

This study focuses on Argentina for the following reasons. First, Argentina is one of the very few countries for which daily observations of the consumer price index are freely available. Second, Argentina is an example par excellence for a country that suffers from structural breaks. Even in a time span as short as the last two decades D'Amato et al. (2007) identify two substantial structural breaks in inflation due to

[1] For an empirical investigation see Fernández-Villaverde and Rubio-Ramírez (2007).

shifts in the Argentine monetary regime. The first structural break was engendered by the external and financial crisis in 1982, which resulted in hyperinflation and finally in the establishment of the Convertibility Act in 1991. The second structural break occurred due to the abandonment of the Convertibility Act in 2002 as a consequence of the sharp depreciation of the Argentine Peso in the currency crisis of the early 2000's.[2] Obviously, estimating a quarterly NKPC under such extreme circumstances leads to misleading results. Finally, the so-called scraped price indices (i.e., indices calculated from online prices as done so by the Billion Prices Project) report more realistic values for Argentine inflation compared to official statistics. For instance, from October 2007 to March 2011 the scraped price data annual inflation rate was 20.14 % compare to 8.38 % reported by the National Statistics Institute (Cavallo 2012b). From our point of view the results from an estimation of a daily NKPC presented in this paper help to induce more confidence in the Argentine inflation dynamics and in general might be seen as being quite reasonable, especially in cases where the official statistics on inflation have to be challenged.

Applying our method to Argentine data we find averagely fixed prices of approximately two to three months. These values are well in line with micro evidence for Argentina from Cavallo (2012a). The results have strong implications for the modeling of monetary and fiscal policy analysis. First, they imply that for Argentina the Calvo parameter has to take much lower values going along with an increase in the frequency of price adjustment compared to the standard calibrations for the United States or the Euro Area. Second and most important, an average price stickiness of a little less than or equal to one quarter for Argentina means that at a quarterly frequency a flexible price model has to be applied to analyze the effects of policy measures. However, under standard assumptions it is well known that the analysis of monetary policy is redundant in this case. In the same vein, to analyze monetary policy in a sticky price framework, a monthly model like the equivalent (augmented) variation of the standard 3-equation New-Keynesian model (NKM) in Franke and Sacht (forthcoming) seems more appropriate.

Modeling the impact of different period lengths on the dynamics of the current workhorse-model used for monetary and fiscal policy evaluation, the NKM with sticky prices (and wages), has been done by Flaschel et al. (2008); Anagnostopoulos and Giannitsarou (2010) as well as Franke and Sacht (forthcoming) in the first place. The latter show that diverging from the standard assumption of the baseline period length to be one quarter dramatically changes the dynamic properties of the model. In particular, the authors state that while determinacy of the model remains unaffected, the impulse response functions can differ in a quantitative and qualitative significant way just by increasing the frequency of decision making (i.e., assuming a monthly, weekly or daily length of the period).

Moreover, there is an increasing interest in the literature on the high-frequency behavior of price changes. In a scanner data study for British supermarkets Ellis (2009) shows that the frequency of price changes is considerably higher in high-frequency studies compared to the traditional monthly or quarterly consumer price

[2] See the corresponding Figure in the Appendix B and D'Amato et al. (2007).

index analysis undertaken by statistical agencies. Furthermore, Ellis (2009) shows that lower-frequency data tends to overstate the true price stickiness. Abe and Tonogi (2010) strongly support this conclusion for the Japanese market. In addition, Kehoe and Midrigan (2007) find very short average price stickiness spells for suburban Chicago and Cavallo (2012a) finds this for Argentina, Brazil, Chile and Colombia.

The remainder of the paper is organized as follows. In the next section, we derive an open economy version of the high-frequency NKPC which we use for our empirical investigation. In Sect. 3 we first describe the data as well as the estimation technique and present the empirical results. Afterwards, we discuss the implications of our results for monetary and fiscal policy analysis in Argentina. Finally, Sect. 4 concludes.

2 The high-frequency New-Keynesian Phillips curve

An extensive analysis of the microfoundation of the (quarterly) NKPC under the standard assumption of the Calvo (1983) price setting scheme in closed and open economies can be found, e.g., in Galí (2008) and Walsh (2010) among others. The standard purely forward-looking NKPC reads as follows:[3]

$$\pi_t = \beta(h_d) E_t \pi_{t+1} + \kappa(h_d) \left(\mu + mc_t^r\right), \tag{1}$$

with

$$\kappa(h_d) = \frac{(1-\theta(h_d))(1-\theta(h_d)\beta(h_d))}{\theta(h_d)},$$

where $\pi_t = p_t - p_{t-1}$ denotes the domestic inflation rate. Although the structural representation of the NKPC does not differ from the one known from the literature we refer to (1) as a high-frequency NKPC since the domestic price level p_t, which is expressed in domestic goods, is not given in quarterly but in daily magnitudes instead.[4]

[3] Alternatively to the purely forward-looking NKPC, we could also apply a hybrid version of the NKPC. This can be brought about either by assuming rule-of-thumb price setters à la Galí and Gertler (1999) or by assuming that non-reset prices are indexed to inflation as in Christinao et al. (2005). However, given that price changes are—in general—costly, the assumption of indexation to daily inflation rates is simply implausible. Furthermore, the estimation of a rule-of-thumb based hybrid NKPC yields the result that the share of backward-looking agents is not statistically different from zero for each estimation considered in this paper. Thus, the hypothesis of a hybrid NKPC on a daily basis—in our setting—has to be rejected. The insignificance of lagged inflation might be the result of the low autocorrelation of daily inflation, as can be seen from the Fig. 1 (lower panel) in the Appendix B. To support this presumption we perform a serial correlation Lagrange multiplier test on inflation (Godfrey 1978 and Breusch 1979) for lag lengths between 2 days and up to 1 week. In neither case we find any statistical significant lags at the 5 % confidence level. This result implies that daily inflation is not (or only very marginally) autocorrelated. Consequently, we restrict ourselves to the purely forward-looking NKPC.

[4] Note that there is no change in the price-setting behavior of the representative firm on a higher frequency, i.e., it is still the aim of the firm to minimize the (discounted) expected deviations of all its future optimal prices (defined as the real marginal costs times a mark-up) from the future market prices. Here the future is not denoted as the next quarter, but as the next day instead. In other words, due to the underlying rational expectation hypothesis firms use all information available when applying their optimization which leads

In contrast to the standard literature we consider the underlying period length denoted by $0 < h_{i,j} \leq 1$ explicitly, where $i, j \in$ {d=daily,m=monthly, q=quarterly}. Hence, we generally allow the representative firm to make its decisions and carry out the corresponding transactions over a period length of h_i relative to the benchmark interval which is fixed as a quarter ($h_q = 1$). In particular the values of two structural parameters are dependent on the frequency of decision making: in order to extract the corresponding discount parameter and the degree of price stickiness at a frequency lower than a day these frequency-dependent parameters of the NKPC have to be suitably adjusted. Regarding Eq. (1) it follows that the daily discount parameter and the expectations operator are given by $0 < \beta(h_d) < 1$ and E_t respectively. The symbol $\theta(h_d)$ stands for the Calvo price stickiness parameter, i.e., the price of a representative firm remains unchanged with a probability $\theta(h_d)$ within a day. The price markup (due to monopolistic competition) is given by μ and $mc_t^r = mc_t - p_t$ are real marginal costs.

Generalizing Flaschel et al. (2008) and Franke and Sacht (forthcoming) we claim that for a representative firm, within a period of length h_i, the probability of resetting the price will be $\frac{h_i}{h_q}(1-\theta(h_q))$, where the symbol $\theta(h_q)$ is retained for the constituent Calvo price stickiness parameter from the quarterly setting.[5] The converse probability is then just given by

$$\theta(h_i) = 1 - \frac{h_i}{h_q}(1 - \theta(h_q)). \tag{2}$$

Thus, given the probability of not resetting the price within a quarter, $\theta(h_q)$, the corresponding probability in, e.g., daily magnitudes is $\theta(h_i = h_d)$ where $h_i = h_d$ is equal to 1/75 since a quarter consists of 75 days on average (excluding weekends).[6] In general, by rearranging the previous formula

$$\theta(h_j) = 1 - \frac{h_j}{h_i}(1 - \theta(h_i)) \tag{3}$$

we are able to extract the value of the frequency-dependent parameter from a lower frequency ($\theta(h_j)$) out of a higher frequency ($\theta(h_i)$). For instance, given the value of the Calvo parameter in daily magnitudes $\theta(h_i = h_d)$—as a result of our estimations in

Footnote 4 continued
to a quick adaptation to new information, i.e., changes in the economic environment. Obviously, this is in favor of the underlying Calvo (1983) price-setting scheme, which is specified in continuous time rather than discrete time.

[5] The following procedure can also be found in the modeling of search and matching processes; see, e.g., Mortensen (1986) and Rogerson et al. (2005). Anagnostopoulos and Giannitsarou (2010) analyze local stability under consideration of changes in the period length quite similar to Flaschel et al. (2008).

[6] Note that obviously the probability for not changing the price is higher at a higher frequency, i.e., at a period length of a day relative to a quarter. Furthermore, the stickiness remains the same in the sense that on average a firm is allowed to reset the price every $1/[1-\theta(h_i)]$ periods of length h_i (say a day) which—independently of h_i—means every $h_i/[1-\theta(h_i)] = h_i/[1-1+h_i(1-\theta(h_q))] = 1/(1-\theta(h_q))$ quarters, respectively.

Sect. 3—we are interested in the value of the Calvo parameter in monthly $\theta(h_j = h_m)$ magnitudes instead. Hence for $h_d = 1/75$ and $h_m = 1/3$ (for the latter note that a quarter consists of 3 months) the transition from monthly to daily magnitudes under consideration of the value for $\theta(h_i = h_d)$ is simply given by applying Eq. (3). Since, in New-Keynesian models, the underlying time period is one quarter by assumption, we are particularly interested in the value of the quarterly Calvo parameter, derived from daily information. Keeping Eq. (3) in mind we are going to address this issue in Sect. 3.

Finally, the discount factor β is also frequency-dependent since, e.g., a discount rate given by $\rho(h_j = h_q)$ of 1.01 % per quarter means that a certain asset is discounted by $h_i = h_d$ times 1.01 % from one day to another. Therefore, it holds that

$$\beta(h_i) = \left(1 + \frac{h_i}{h_j}\rho(h_j)\right)^{-1}. \tag{4}$$

However, while domestic prices are given in daily magnitudes by the Billion Prices Project at MIT Sloan (a detailed description of the data is provided in Sect. 3) this does not hold for the real marginal costs mc_t^r and the mark-up μ. Therefore, we consider an open economy version of the NKPC and substitute μ and mc_t^r by appropriate proxies which can be expressed in daily magnitudes as well.[7] Note that up to this point the structure of the NKPC does not differ in closed and open economies (e.g., Galí 2008 or Clarida et al. 2002). The last term in (1) can be substituted by the expressions for the domestic (y_t) and foreign (y_t^f) output gap. Hence,

$$\pi_t = \beta(h_d)E_t\pi_{t+1} + \kappa(h_d)\left[(\sigma_\alpha + \eta)y_t - (\sigma_\alpha - \sigma)y_t^f\right], \tag{5}$$

where $\sigma_\alpha = \sigma[1 - \alpha + \alpha(\sigma\gamma + (1-\alpha)(\sigma\chi - 1))]^{-1}$ is a function of the degree of openness $0 \leq \alpha \leq 1$ (calibrated to match Argentina's share of foreign goods in consumption), the substitutability between domestic and foreign goods from the viewpoint of the domestic consumer χ, the substitutability between goods produced in different foreign countries γ and the inverse intertemporal elasticity of substitution in consumption of domestic goods σ (Galí 2008). The parameter η denotes the substitution elasticity of labor. In order to get an appropriate expression for a daily NKPC, we first make use of the (log-linearized) terms of trade s_t, i.e., the terms of trade which is defined by the price of foreign goods in terms of home goods. Following Clarida et al. (2001, 2002), we assume that there exists a relationship between the terms of trade *gap* and both (domestic and foreign) output gaps

$$\frac{1}{\sigma_\alpha}(s_t - \tilde{s}_t) = y_t - y_t^f, \tag{6}$$

[7] Note that in addition neither the domestic and the foreign output gaps (as we will discuss below) nor the labor share of income stand for appropriate proxies, since both are also not available in daily magnitudes.

where $\tilde{s}_t$ stands for the terms of trade in the steady state.[8] By applying (6) on (5) we are able to derive an open economy NKPC which depends on the terms of trade and the domestic output gap:[9]

$$\pi_t = \beta(h_d)E_t\pi_{t+1} + \kappa(h_d)\left(\frac{\sigma_\alpha - \sigma}{\sigma_\alpha}(s_t - \tilde{s}_t) + (\eta + \sigma)y_t\right). \tag{7}$$

As we can see from (7) the problem concerning the frequency remains, since data on both gap specifications is also not available on a daily basis. Therefore, we consider the underlying intertemporal optimization problem of the representative household who seek to maximize its utility function under consideration of the related budget constraints. We apply optimal control theory on standard expressions for a (separable) money-in-the-utility function, a budget and a cash-in-advance constraint known from the literature (Appendix A). The latter is given by

$$y_t = m_t^r, \tag{8}$$

i.e., consumption expenditures are not allowed to exceed the real money holdings of the household, where the latter is denoted by m_t^r. The optimality condition regarding money demand depends on the nominal interest rate:

$$m_t^r = \frac{1}{\psi}(\sigma y_t - \beta(h_d)i_t), \tag{9}$$

where ψ is the inverse elasticity of money demand. Substituting (8) into (9) and re-arranging leads to

$$y_t = \left(\frac{\beta(h_d)}{\sigma - \psi}\right)i_t. \tag{10}$$

Since data on movements in the terms of trade are also not available on a daily basis, we consider two types of the high-frequency version of the NKPC. Therefore, the (re-arranged) uncovered interest parity is given by

$$s_t = E_t e_{t+1} + i_t^f - i_t - p_t + p_t^f, \tag{11}$$

where i_t (i_t^f), e_t and p_t^f denote the domestic (foreign) nominal interest rate, the nominal exchange rate and the foreign price level, respectively. The corresponding steady state expression reads

[8] Under the assumption of complete securities markets, Eq. (6) implies that relative consumption and hence (in a general equilibrium framework) the output gap relation across countries is proportional to the terms of trade. For a formal proof refer to Lubik and Schorfheide (2007) and Appendix A of Galí and Monacelli (2005). For an empirical discussion see Chari et al. (2002) among others.

[9] A related approach to estimate an open economy NKPC for a quarterly frequency has been applied by Mihailov et al. (2011a).

$$\tilde{s}_t = \tilde{e}_t + \tilde{i}_t^f - \tilde{i}_t - \tilde{p}_t + \tilde{p}_t^f. \tag{12}$$

Both lead to the following equation denoted as *Type I* NKPC:

$$\pi_t = \beta(h_d)E_t\pi_{t+1} + \kappa(h_d)\left[\phi_1\left(E_t\Delta e_{t+1} + \Delta i_t^f - \Delta i_t - \Delta p_t + \Delta p_t^f\right) + \phi_2 i_t\right] \tag{13}$$

with $\phi_1 = \frac{\sigma_\alpha - \sigma}{\sigma_\alpha}$ and $\phi_2 = \frac{(\eta+\sigma)\beta(h_d)}{\sigma-\psi}$. We define a gap by $\Delta a_t = a_t - \tilde{a}_t$ with $a_t = \{e_t, i_t, i_t^f, p_t, p_t^f\}$ and $\tilde{a}_t = \{\tilde{e}_t, \tilde{i}_t, \tilde{i}_t^f, \tilde{p}_t, \tilde{p}_t^f\}$. Within this specification the driving forces of domestic inflation are the domestic nominal interest rate, the expected bilateral nominal exchange rate gap, the domestic nominal and foreign interest rate gaps, and the domestic and foreign price level gaps. For the latter we follow Monacelli (2004, p. 201) and set $p_t^f = \tilde{p}_t^f = 0$ (i.e., $\Delta p_t^f = 0$) since in his investigation he assumes "that only a negligible small share of domestic goods is consumed in the rest of the world and therefore foreign inflation is zero."

Note that the NKPC in (13) depends on the domestic and foreign interest rate gap (and on the terms of trade gap in general) which defines the relative interest rate gap, i.e., $\delta i_t = \Delta i_t^f - \Delta i_t$. This relationship can be motivated as follows: suppose $\delta i_t > 0$ and $\tilde{i}_t = \tilde{i}_t^f = \tilde{s}_t = 0$. In this case $i_t^f > i_t$ and hence under consideration of the standard Euler equation which determines domestic and foreign consumption, c_t and c_t^f, respectively, in a DSGE context this means that $x_t^f < x_t$ via the real interest channel, where x_t (x_t^f) denotes the domestic (foreign) level of output. However, if $x_t - x_t^f > 0$, the domestic country runs a trade surplus, i.e., there exists an excess supply of domestic goods in the international goods market. It follows that domestic goods must become cheaper for market clearing and this leads to an appreciation of the terms of trade ($\Delta s_t = s_t - \tilde{s}_t$ increases). Since imported goods (used in the production process of the representative firm) become more expensive, the inflation rate must rise. This interpretation is analogous to the discussion on the exchange rate channel of monetary policy induced by Leitemo and Söderström (2005) and Ireland (2005) among others. Moreover, a direct link between real marginal costs and the terms of trade is shown by Galí and Monacelli (2005, p. 718) since "changes in the terms of trade has a direct influence on the product wage, for any given real wage."

The so-called *Type II* NKPC, which we use for robustness checks, is given by

$$\pi_t = \beta(h_d)E_t\pi_{t+1} + \kappa(h_d)\left[\phi_1\left(\Delta e_t - \Delta p_t\right) + \phi_2 i_t\right], \tag{14}$$

where for s_t and $\tilde{s}_t$ we substitute the definition of the terms of trade $s_t = e_t + p_t^f - p_t$ and the corresponding expression for the steady state $\tilde{s}_t = \tilde{e}_t + \tilde{p}_t^f - \tilde{p}_t$. Once again, $p_t^f = \tilde{p}_t^f = 0$ holds. Note that the time series for domestic prices (and, of course, for the domestic inflation rate), the interest rates and for the bilateral exchange rate are all available on daily frequencies. Hence both *Type I* and *Type II* NKPCs can be seen as high-frequency Phillips curves in daily magnitudes. Taking the Eq. (3) into account we are able to calculate the values of the monthly and quarterly Calvo parameters which correspond to the estimated degree of price stickiness in daily magnitudes.

3 Empirical analysis

In this section we analyze the empirical implications of the *Type I* and *Type II* high-frequency NKPCs—(13) and (14)—for the adjustment speed of prices on a daily basis. We define Argentina as the domestic economy. Following D'Amato and Garegnani (2009) the foreign economy comprises Argentina's three most important trading partners Brazil, the Euro Area and the United States, where Brazil gets assigned the largest weight in the basket. For robustness, we also check for bilateral arrangements with both Brazil and the United States being the foreign economy. The choice for a dominating Brazil can be justified by Brazil's relative importance in mutual trade flows arising from the geographical proximity and the joint membership in the Mercado Común del Sur, the Southern Common Market. Being Argentina's number one trading partner, Brazil accounts for approximately 22 % of exports and roughly one third of all imports. These numbers strongly outweigh the second most important trading partner, the United States, who account for only 8 % of exports and 15 % of imports (World Bank 2010). Nevertheless, to test for robustness we also report results for the United States being the foreign country.

3.1 Data

The dataset comprises daily observations for Argentina, Brazil and the United States from 12-03-2007 to 04-02-2011. For proper identification we present the mnemonic codes in parentheses after each variables.

Argentine inflation is defined as annualized daily percentage change in the consumer price index (CPI). We resort to the CPI instead of the implicit GDP-deflator since the former is the conceptually appropriate indicator in a number of open economy New-Keynesian models such as Galí and Monacelli (2005), D'Amato and Garegnani (2009), Mihailov et al. (2011a) and Mihailov et al. (2011b).[10]

One might argue that supermarket products just represent 40 percent of all CPI expenditure (i.e., if services are not considered) and therefore it is misleading to call the corresponding scraped data index considered in this paper a consumer price index. In several studies, Cavallo contradicts this statement. He states that food and household products have been the main driver of Argentine inflation during the time period considered here (Cavallo 2010). Furthermore, the author mentions that "this limitation (of not considering the prices of services) can be overcome as a growing number of firms start posting their prices online" (Cavallo 2012a, p. 6), especially those for services. Finally and most important, Cavallo argues that official Argentine inflation statistics have become rather unreliable due to the intervention of the Argentine government in the National Statistics Institute since the year 2007 (Cavallo 2012b). He shows that

[10] Holmberg (2006) even shows that CPI data results in more realistic estimates for an open economy NKPC in Sweden compared to the use of the GDP-deflator. In empirical applications of the closed economy version of the NKPC this approach is common too due to reasons of data availability. Recent examples are, among others, Ramos-Francia and Torres (2008) and Yazgan and Yilmazkuday (2005). Moreover, Nason and Smith (2008) apply both measures to test for robustness and find the differences in performance to be negligible for the United States.

Table 1 Calibration

	σ	χ	η	γ	ψ	α	$\rho(h_q)$
Escudé (2009)	1.902	0.700	1.175	0.990	1.180	0.139	0.013
Escudé (2007)	4.960	1.000	1.194	3.500	1.563	0.112	0.013

online and official estimates share a similar pattern over time, while there is a high correlation between both indices (Cavallo 2012b, Fig. 3 and Table 2 in his Appendix). Cavallo concludes that there exist a difference in the level of inflation between online and official statistics, but not in the dynamic behavior of inflation (Cavallo 2012b). Studies on Argentine inflation dynamics—like in this paper—must account for these characteristics which do not certainly hold for the Unites States or Euro Area. Therefore, our approach offers the opportunity to use high-frequency data for the empirical evaluation of low-frequency behavior in optimization.

The consumer price index (indicecanastabasica) is provided by www.inflacionverdadera.com, which is a subproject from the Billion Prices Project at MIT Sloan. The underlying price data are collected on a daily basis from large supermarkets in the metropolitan area of Buenos Aires.[11] All remaining data is taken from Datastream®. In particular, we apply the Argentine Peso to Euro and to United States dollar exchange rates, which are (TEARSSP) and (TDARSSP), respectively. In addition, we derive the Argentine Peso to Brazilian Real exchange rate from the exchange rates for the Argentine Peso to the United States dollar and the Brazilian Real to the United States dollar (TDBRLSP). As home and foreign interest rates, we apply the Argentine 1-day Buenos Aires Interbank Offer Rate (AGIBK1D) and the United States Effective Federal Funds Rate (FRFEDFD), respectively. In the Robustness exercise we also apply the Brazilian Sistema Especial de Liquidção e de Custódia (Selic) Base Interest Rate (BROVERN). For our instrument set we choose an alternative consumer price index from the Billion Prices Project at MIT Sloan, which comprises solely food and beverages (indicealimentosybebidas).

3.2 Calibration

We calibrate the real interest rate to 5.2 % according to World Development Indicators reported by the World Bank (2009), which yields a quarterly discount factor $\rho(h_q) = 0.013$.[12] All remaining parameters are calibrated according to Escudé (2009), who estimates a medium scale open economy DSGE model for Argentina. Thus, we set the inverse intertemporal elasticity of substitution for domestic goods $\sigma = 1.902$ and the inverse intertemporal elasticity of labor $\eta = 0.7$. Domestic and foreign goods

[11] In particular, the prices of 150 products are checked online every day. This methodology is sufficient since 100 % of all products in Argentine supermarkets can also be found online (Cavallo 2012a). For a thorough discussion of the methodology of the billion prices project at MIT Sloan, we refer to Cavallo (2012a,b), Cavallo and Rigobon (2011), www.inflacionverdadera.com and www.billionpricesproject.com.

[12] According to Eq. (4) the discount rate in daily magnitudes is than equal to $\beta(h_d) = 1/(1 + \frac{h_d}{h_q}\rho(h_q)) = 0.999$ where $h_d = 1/75$ and $h_q = 1$.

are assumed to be imperfect substitutes as well as are the different varieties produced in the foreign country. The elasticity of substitution between the former is set to be $\chi = 1.175$, while the latter is given by $\gamma = 0.990$. The inverse elasticity of money demand is calibrated to $\psi = 1.18$. Finally, the degree of openness is calibrated to match Argentina's share of foreign goods in consumption, i.e., $\alpha = 0.134$. To test for robustness we juxtapose the results with an earlier calibration from Escudé (2007) for Argentina. The parameter values are summarized in Table 1.

3.3 Estimation methodology

The empirical analysis rests on both types of the high-frequency NKPC given by the Eqs. (13) and (14). By substitution of the day-by-day expectations error $\varepsilon_t = \beta(h_d)(E_t[\pi_{t+1}] - \pi_{t+1})$ we obtain a regression equation of the form

$$\pi_t = \beta(h_d)\pi_{t+1} + \frac{(1-\theta(h_d))[1-\theta(h_d)\beta(h_d)]}{\theta(h_d)}\xi_{j,t} + \varepsilon_t, \tag{15}$$

with $\xi_{j,t} = \{\xi_{1,t}, \xi_{2,t}\} = \Big\{\phi_1\left(E_t\Delta e_{t+1} + \Delta i_t^f - \Delta i_t - \Delta p_t\right) + \phi_2 i_t, \phi_1(\Delta e_t - \Delta p_t) + \phi_2 i_t\Big\}$.[13] McCallum (1976) shows that under rational expectations the prediction error of future inflation ε_t is uncorrelated to the information set available to the forecaster $\mathbf{z}_t$, which comprises information dated at time t or earlier. This assumption implies that $E_t[\varepsilon_t \mathbf{z}_t] = 0$. Applying this condition to Eq. (15), we obtain

$$E_t\left[\left(\theta(h_d)\pi_t - \theta(h_d)\beta(h_d)\pi_{t+1} - (1-\theta(h_d))(1-\theta(h_d)\beta(h_d))\xi_{j,t}\right)\mathbf{z}_t\right] = 0, \tag{16}$$

with $\mathbf{z}_t$ being a vector of instruments comprising each three lags of the food- and beverage-based consumer price inflation π^{food}, food- and beverages-based consumer price index p^{food}, the United States three-month interbank interest rate i^{3M} and the respective exchange rate e. The instruments (i.e., the variables and their respective lags) are chosen to satisfy two requirements. First, they strongly correlate with the regressors in the estimation equation and second, the instrument set passes Hansen's (1982) J-test for overidentifying restrictions. Furthermore, we restrict ourselves to instruments dated time $t-1$ or earlier. The intuitive reason is straightforward, since not all contemporaneous information might be available by the time agents form their expectations (Galí et al. 2001). Finally, among the potential instrument sets we choose the particular set with the lowest average J-statistic over all applied estimations.

According to McCallum (1976) an orthogonality condition of the kind of (16) can be consistently estimated with an instrument variable technique. The latter has become standard in the literature since the prominent contribution of Galí and Gertler (1999). Therefore, we apply Hansen's (1982) GMM to estimate the structural parameter $\theta(h_d)$.

[13] Following Galí and Gertler (1999), we define deviations from steady state in terms of demeaned time series. The results are robust, however, also to the use of the Hodrick-Prescott filter with $\lambda = 6{,}812{,}100$ for daily observations.

3.4 Results

In this section we discuss the results from the empirical exercise. We focus on three different scenarios concerning the definition of the foreign country. In the first scenario the foreign country is represented by a multi-country-mix of Brazil, the United States and the Euro Area as suggested by D'Amato and Garegnani (2009). As robustness checks, we also report the results from two bilateral analysis with the Brazil and the United States instead. The point estimates for the daily Calvo parameter $\theta\ (h_d)$ and their standard errors for all three cases are summarized in the first column of Table 2.

The estimates for $\theta(h_d)$ lie remarkably close to each other in an interval between [0.9855; 0.9867], even though the admissible range for economically relevant values of $\theta(h_d)$ is from zero to unity and there are no restrictions imposed on this parameter. The standard errors indicate that the parameter estimates are statistically significant for each of the single scenarios. Since Calvo staggering follows a Poisson process, prices are fixed on average for $\mathcal{D} = \frac{1}{1-\theta(h_d)}$ days. As can be seen from column 4 in Table 2, depending on the case considered, prices are fixed between 69 and 75 days. The average duration implied by the high-frequency NKPC over all cases considered lies at approximately 73 days, or a little less than one quarter, for both the *Type I* and the *Type II* high-frequency NKPC.

This result is in stark contrast to the well-known evidence from Dhyne et al. (2006) and Taylor (1999), who report average price spells of approximately 1 year and also compared to Bils and Klenow (2004), who find prices to be fixed on average around one and a half quarters. These prominent results are, however, derived from monthly CPI data for industrialized countries such as the United States and the countries of the Euro Area. Recent micro evidence for non-industrialized economies indicates that the degree of price stickiness is much lower in Latin American economies compared to industrialized economies. Based on monthly CPI data Medina et al. (2007), Gouvea (2007) and Morandé and Tejada (2008) show that prices in Brazil, Chile, Colombia and Mexico change on average at least once a quarter.[14] Moreover, Cavallo (2012a) applies the methodology of Bils and Klenow (2004) to daily CPI data for Argentina, Brazil, Chile and Colombia and also reports average price spells below one quarter. In particular for the case of Argentina, Cavallo (2012a) finds that prices remain unchanged for 66–83 days, an interval which fully encloses the results given in Table 2. Therefore, qualitatively and quantitatively, our macroeconometric high-frequency NKPC procedure yields empirical results, which are in line with microeconometric evidence.

One of the major contributions of this paper is that we can use the daily information to derive lower-frequency information such as weekly, monthly or quarterly by simply employing the daily point estimate to Eq. (3). The results for monthly conversions are given in the second and fifth column of Table 2. For an example of the conversion look at the first line of Table 2 (*Type I*, multicountry case). According to Eq. (3), a daily probability of not being able to reset the price $\theta(h_d)$=0.9867 is equivalent to

[14] For a detailed summary of the micro evidence on the frequency of price setting from CPI data for developed and developing countries we refer to Alvarez (2008) and Klenow and Malin (2010).

Table 2 Price adjustment frequency in Argentina

Foreign country	Calvo parameter			Average duration in			p(J-statistic)	p(AR-statistic) χ^2	p(AR-statistic) F
	$\theta\,(h_d)$	$\theta\,(h_m)$	$\theta\,(h_q)$	Days	Months	Quarters			
Multicountry									
Type I	0.9867	0.6667	0.0002	75	3.00	1.00	0.9115	0.3818	0.3837
	(0.0203)								
Type II	0.9867	0.6671	0.0014	75	3.00	1.00	0.9115	0.3817	0.3836
	(0.0203)								
Brazil									
Type I	0.9867	0.6678	0.0032	75	3.01	1.00	0.8661	0.3801	0.3820
	(0.0204)								
Type II	0.9867	0.6778	0.0035	75	3.01	1.00	0.8661	0.3801	0.3820
	(0.0205)								
United States									
Type I	0.9855	0.6387	0[a]	69	2.77	0.92	0.4525	0.3910	0.3928
	(0.0181)								
Type II	0.9856	0.6390	0[a]	69	2.77	0.92	0.4524	0.3909	0.3927
	(0.0181)								

Note The parameter $\theta(h_d)$ is estimated from the orthogonality conditions given by (16). The standard errors are given in brackets. We apply a 12-lag Newey-West covariance matrix

[a] The parameters $\theta(h_m)$ and $\theta(h_q)$ are calculated according to $\theta(h_j) = 1 - \frac{h_j}{h_d}(1 - \theta(h_d))$ with $j = \{m, q\}$. Note in few cases, applying the previous equation based on our estimated value for $\theta(h_d)$ results in a negative value for $\theta(h_q)$. This implies that prices change more than once within a quarter. We account for this by setting the Calvo parameter in quarterly magnitudes equal to zero which is the natural lower bound of $\theta(h_q)$: $\theta(h_q) = \begin{cases} 0 < (\cdot) < 1 & \text{for} \quad 0 < \varpi < 1 \\ 0 & \text{for} \quad \varpi < 0 \end{cases}$ where $\varpi = 1 - \frac{h_q}{h_d}(1 - \theta(h_d))$. Nevertheless given $\theta(h_q) < 0$ it is still possible to calculate the corresponding average duration of fix prices in quarters. Hence under consideration of the Poisson process we get a duration of less than one

a monthly probability of not being able to reset the price of $\theta(h_m)$=0.6667, which is equivalent to an average duration of fixed prices of exactly 3 months. An analogous transformation can also be applied to the quarterly frequency. In the case of the example just stated above the quarterly Calvo parameter yields $\theta(h_q)$=0.0002, which implies that prices are sticky on average for exactly one quarter. Furthermore, the estimates from Table 2 reveal that the quarterly Calvo parameter might not always be determinable, as in the case for the United States being the foreign country. This is because the Calvo parameter in general reflects a probability and therefore is bounded between zero and unity. Given that prices are on average fixed for about 3 months implies that, on a quarterly basis, prices change once within every period. This, however, is equivalent to a quarterly flex-price model. The implications of this result for the choice of the simulation frequency for policy analysis are discussed in Sect. 3.6 below.

To test whether or not these results are reliable, column 7 of Table 2 present the results of Hansen's J-test for overidentifying restrictions. The p-values for the J-statistic indicate that the validity of the overidentifying restrictions cannot be rejected for any of the specifications estimated in Table 2. Thus, we consider the instruments being valid.

Even though the overidentifying restrictions are satisfied, the estimation results might be biased due to a lack of identification of the structural model or due to the presence of weak instruments. Applying the parameter estimates for $\theta(h_d)$ to Eq. (15), together with a calibrated discount factor close to unity $\beta(h_d) = 0.999$, yields a reduced form parameter for the real marginal cost term of the high-frequency NKPC $\kappa(h_d)$ very close to zero, raising the question whether the structural parameters of the high-frequency NKPC are identified. Mavroides (2007), however, analyzes the structural version of the NKPC in the absence of identification and shows that if the true value of the price stickiness parameter is close to the limiting case of unity—as it holds for the parameter estimates in this paper—the structural model is weakly identified. Furthermore, Mavroides (2007) argues that inference on the structural parameters is still possible under weak identification, if tests are applied which are robust to weak identification. In order to test for weak identification we follow Yazgan and Yilmazkuday (2005) and apply the nonlinear Anderson–Rubin (AR) statistic.[15] As has been shown by Stock et al. (2002) the nonlinear AR-statistic is fully robust even in the case of poorly identified parameters or weak instruments. The AR statistics test the null hypothesis that given our instrument set the estimated parameter values are the true parameter values. Columns 8 and 9 of Table 2 summarize the p-values for two complementary AR-statistics, given the null hypothesis stated above. For neither specification in Table 2 the null hypothesis can be rejected, which implies that the parameter estimates in Table 2 are admissible to the data, i.e., the point estimates for the Calvo parameter in daily magnitudes lie in the identification robust confidence set. Therefore, we render our estimates valid, even in the

[15] For a general discussion of the nonlinear Anderson-Rubin-statistic we refer to Stock et al. (2002). For applications of this test statistic to the NKPC we refer to Ma (2002), Khalaf and Kichian (2004) and Yazgan and Yilmazkuday (2005).

absence of full identification and independently of whether the instruments are weak or strong.

3.5 Robustness exercises

To analyze the robustness of our estimates the upper half of Table 3 presents the results for an alternative calibration from Escudé (2007). For the *Type I* specification, the estimated duration of fixed prices decreases a little, while it remains hardly unchanged for the *Type II* specification. Still, all values are well in line with microeconometric evidence. Thus, quantitatively and qualitatively the results are robust to the results from the standard calibration. Further testing for robustness includes the variation of the instrument set. Given the base calibration, the lower half of Table 3 presents the estimated Calvo parameters conditional on three alternative instrument sets. Under the alternative instrument sets the price adjustment speed increases slightly from an average duration of 3 months to an average duration of approximately two and one-half months. However, the point estimates remain in the close neighborhood of our results in Table 2. The p-values of the J-statistics and the AR-statistics given in columns 5 and 6 once again indicate that the overidentifying restrictions are satisfied and that the parameter estimates are data admissible in each of the cases presented. The robustness analysis supports the assertion that for Argentina a quarterly model would need to be calibrated as a flex-price model since in all cases $\theta(h_q)$ is set to zero, which is its natural lower bound.

3.6 Discussion

In this section we discuss some implications of our empirical results for theoretical interpretation and policy analysis. As indicated in the introduction, our approach is not restricted to the use of daily data. Daily data serves mainly as an extreme example for possible applications of our approach. It follows from the transformation Eq. (3) that every frequency, for which the appropriate time series are available, is applicable. Nevertheless, the choice of high-frequency data over lower-frequency data in the estimation process is appealing for several reasons. First, even a short period of time delivers a large data sample, reducing the small sample bias. Second, relying on shorter time spans reduces the risk of having structural breaks in the data. Third, Ellis (2009) points out that studying the frequency of price changes at lower frequencies, such as quarterly, tends to overstate the true price stickiness, an argument which is also supported by Abe and Tonogi (2010).

As has been shown in Sect. 3.4, estimating the high-frequency NKPC with daily data results in an almost flat Phillips Curve tradeoff, i.e., as $\kappa(h_d) \approx 0$. However, the flatness of the NKPC neither contradicts its economic interpretation nor its quantitative importance. Franke and Sacht (forthcoming) show that monetary policy shocks have a strong impact on economic dynamics even on a high frequency. The authors juxtapose the impulse response functions for inflation to a monetary policy shock at high and low frequencies and show that the reaction of inflation is even slightly stronger for the daily model relative to the quarterly model. Intuitively, the closer the future is, for

Table 3 Robustness exercises

Foreign country	Calvo parameter $\theta\ (h_d)$	$\theta\ (h_m)$	Average duration in Days	Months	$p(J\text{-stat.})$	$p(\text{AR-stat. } \chi^2)$ $p(\text{AR-stat. } F)$
Result for an alternative calibration from Escudé (2007)						
Multicountry						
Type I	0.9824	0.5600	57	2.27	0.9117	0.3758
	(0.0240)					0.3778
Type II	0.9828	0.5696	58	2.32	0.9112	0.3759
	(0.0246)					0.3779
Brazil						
Type I	0.9827	0.5667	58	2.31	0.8657	0.3836
	(0.0245)					0.3855
Type II	0.9828	0.5690	58	2.32	0.8655	0.3832
	(0.0245)					0.3850
United States						
Type I	0.9807	0.5167	52	2.07	0.4504	0.3922
	(0.0215)					0.3940
Type II	0.9809	0.5227	52	2.10	0.4510	0.3922
	(0.0213)					0.3940
Robustness with respect to the instrument set						
Set 1 = z_t^{1a}						
Type I	0.9798	0.4953	50	1.98	0.9018	0.3379
	(0.0435)					0.3401
Type II	0.9800	0.4989	50	2.00	0.9018	0.3379
	(0.0437)					0.3401
Set 2 = z_t^{2a}						
Type I	0.9855	0.6381	69	2.76	0.6563	0.4530
	(0.0186)					0.4545
Type II	0.9858	0.6380	69	2.76	0.6564	0.4530
	(0.0187)					0.4545
Set 3 = z_t^{3a}						
Type I	0.9841	0.6017	63	2.51	0.9386	0.4413
	(0.0156)					0.4428
Type II	0.9841	0.6020	63	2.51	0.9386	0.4413
	(0.0156)					0.4429

Note The parameter $\theta(h_d)$ is estimated from the orthogonality conditions given by (16). The standard errors are given in *brackets*. We apply a 12-lag Newey-West covariance matrix. The parameter $\theta(h_m)$ is calculated according to $\theta(h_m) = 1 - \frac{h_m}{h_d}(1 - \theta(h_d))$

We resort to report the values for $\theta(h_q)$ as well as the average duration of price setting in quarters since they are 0 and less than one respectively. This is because in all cases prices change less than once a quarter

[a] The alternative instruments are defined as follows: $(\Delta\pi_{t-i}^{\text{food}}, \Delta p_{t-i}^{\text{food}}, \Delta i_{t-i}^{3M}, \Delta e_{t-i}) \in \mathbf{z}_t^1$ for $i = 1, 2, 3$; $(\pi_{t-i}^{\text{food}}, p_{t-i}^{\text{food}}, i_{t-i}^{3M}, e_{t-i}) \in \mathbf{z}_t^2$ for $i = 1, 2, 3, 4$; $(\pi_{t-i}^{\text{food}}, p_{t-i}^{\text{food}}, i_{t-i}^{3M,US}, i_{t-i}^{3M,AR}, e_{t-i}^{US}, e_{t-i}^{EU}, e_{t-i}^{BR}) \in \mathbf{z}_t^3$ for $i = 1, 2$

instance one day ahead versus one quarter ahead, the stronger is the impact on today's choices (Anagnostopoulos and Giannitsarou 2010). Consequently, today's reaction of a firm becomes stronger as the length of the decision period decreases. Thus, from the theoretical point of view, an almost flat high-frequency NKPC is absolutely feasible and what is to be expected, when analyzing a daily frequency.

The above arguments support our choice for the use of daily data to estimate the frequency of price adjustment using a high-frequency NKPC. However, we would like to stress that this does not call in any form for the reduced-form or the structural version of the model in daily magnitudes to be the most appropriate framework to analyze fiscal and monetary policy measures. In fact, the frequency for theoretical policy analysis should be appropriately chosen given the question at hand. The model builder needs to consider that even though the impulse response functions remain qualitatively unaltered by the frequency of observation, quantitatively this need not to be true. Franke and Sacht (forthcoming) show that the lower the observation frequency, the larger the quantitative differences between the impulse response functions to a monetary policy shock. Despite these differences being not fully negligible, they are small. It can be shown analytically and numerically that a quarter is not a robust period, which is defined as an upper-bound on the length of the decision intervals with essentially similar impulse response functions. In particular their results account for the fact that for the variables the deviations of the weekly and monthly from the quarterly economy are not negligible.[16]

Given the choice for a specific frequency, our estimates guide the decision, whether to conduct policy analysis for a particular country using a flexible price framework like a Real Business Cycle model or a sticky-price model of New-Keynesian type. For instance, if economic policy analysis is to be conducted in a quarterly setting, our results simply imply that for Argentina a flexible price setting ought to be chosen rather than a sticky price setting. The reason is straightforward since, according to our estimates and the microeconometric evidence on price setting by Cavallo (2012a), prices in Argentina (on average) change at least once a quarter and therefore are flexible in a quarterly setting. Note that in this case the analysis of monetary policy measures is lapsed since in the standard flex-price model money is neutral with respect to real variables both, in the short and in the long run.[17] This does not hold true for fiscal measures, which can easily be analyzed also in a flex-price economy. Alternatively, our results can be interpreted in a way that they state the minimum frequency applicable, given the choice for the model type. For instance, in order to analyze the real effects of (monetary) policy measures in a staggered pricing framework for Argentina, the model frequency must be no lower than monthly.

4 Conclusion

This study aims at mitigating the shortcomings of the estimation of the NKPC arising from small sample bias and structural breaks. To account for such problems, we apply the standard GMM approach to estimating an open economy NKPC at daily frequency and transform the results into pseudo-quarterly equivalences. We show that this methodology is highly sufficient to estimate the Calvo parameter of price stickiness from the end of 2007 up to now just under the consideration of daily time

[16] An explicit empirical investigation concerning the period length h in sticky and flexible price models can be found in Aadland (2001) and Christiano (1985). Further investigation of this issue is necessary.

[17] This result only holds for the standard flex-price model. Once we take into account for instance non-separable utility between consumption and money-in-the-utility function, shopping time models or cash-in-advance constraints, monetary policy is—although in quantitative terms very little—effective (Walsh 2010).

series data. Applying our method to Argentine data we find the daily Calvo parameter over all cases considered to lie within the narrow interval of (0.9798;0.9867), which implies averagely fixed prices of approximately 2–3 months. These values are well in line with microeconometric evidence for Argentina from Cavallo (2012a). Our results have some implications for the modeling of monetary and fiscal policy analysis. First, they imply that the frequency of price adjustments has to be calibrated much higher for Argentina compared to the United States or the Euro Area. Second and most important, an average price stickiness of nearly one quarter means that at a quarterly frequency a flexible price model has to be applied to analyze the effects of policy measures, a situation, which—under standard assumptions—renders the analysis of monetary policy redundant. In the same vein, to analyze monetary and fiscal policy in a sticky price framework, a monthly model like the equivalent (augmented) variation of the standard 3-equation NKM in Franke and Sacht (forthcoming) seems more appropriate. Our results turn out to be robust not only to an alternative calibration of the model, but also to variations in the instrument set. In the end we offer a way to expose the dynamics on a quarterly basis given by the underlying transformation approach presented in this paper. On the second stage this pseudo-quarterly values might be useful for the analysis of monetary and fiscal policy in a low-frequency environment.

The question arises if the approach presented in this paper is sufficient to reproduce equivalent results for industrial countries like the United States or the Euro Area, where several studies (Bils and Klenow 2004 and Klenow and Malin 2010 among others) show higher degrees of price stickiness. However this depends, of course, on the availability of the data. Moreover in order to compare the impact of a high-frequency NKPC to its standard low frequency version in an optimal monetary or a fiscal policy experiment a frequency-adjusted two- or multi-country model is needed. Due to several structural breaks (especially in the case of Argentina) the estimation of the value of the quarterly Calvo parameter based on the past 50 years seems to be inappropriate as we have discussed in this paper.

Finally for a better understanding of the implications of our results for conducting optimal monetary (and fiscal) policy one might consider a model in which firms adjust their prices fully flexibly to firm-specific (idiosyncratic) shocks but sluggishly to aggregate shocks. The former captures strategic motives, e.g., the (re)filling of stations where the latter stands for policy or other exogenous shocks which effect the overall economic conditions. The question arises, whether we observe price changes on a day-to-day basis due to idiosyncratic or aggregate shocks? Furthermore which of these shocks appears more frequently? Using a model where it is assumed that aggregate shocks occur rarely while idiosyncratic shocks happen frequently, might help to understand the information content of daily and of quarterly data. Klenow and Malin (2010) report that micro price changes in the Euro Area and the United States do not keep up with overall inflation. The authors claim that this could be explained most probably by the dominance of idiosyncratic over aggregate shocks. In this respect the approach of Svensson and Woodford (2003) could be helpful. The authors discuss optimal monetary policy in a NKM under the assumption of a symmetric partial information distribution between agents. In particular, they account for uncertainty regarding the level of the potential output and the cost-push shock which are unobservable and partially observable, respectively. Following Svensson and Woodford (2003) by considering a measurement equation that accounts (in our case) for data with different frequencies could shed light on the importance of idiosyncratic or aggregate shocks. We leave these aspects to further research.

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Appendices

Appendix A: Household's optimization problem

Following Woodford (2003) and Galí (2008) we consider a standard intertemporal utility function of a representative household, which is additive separable in consumption C_t, in real money holding M_t/P_t (where P_t is the price level) and in leisure (labor supply N_t, respectively), subject to the underlying frequency in daily magnitudes:

$$U_t = E_t \sum_{k=0}^{\infty} \beta(h_d)^k \left[\frac{1}{1-\sigma} C_{t+k}^{1-\sigma} + \frac{1}{1-\psi} \left(\frac{M_{t+k}}{P_{t+k}} \right)^{1-\psi} - \frac{1}{1+\eta} N_{t+k}^{1+\eta} \right], \quad (17)$$

where $\sigma > 0$ represents the inverse intertemporal elasticity of substitution between present and future consumption of domestic goods, $\eta > 0$ is the inverse of the substitution elasticity of labor, $\psi > 0$ stands for the inverse of the elasticity of money demand, $\beta(h_d) > 0$ is the (frequency-dependent) discount factor and E_t stands for the expectations operator.

The period-by-period budget constraint is characterized by money and bond holdings of the representative household and is given in real terms by

$$\begin{aligned} C_{t+k} = & -\frac{B_{t+k}}{P_{t+k}} - \frac{\Delta M_{t+k}}{P_{t+k}} + \frac{W_{t+k}}{P_{t+k}} N_{t+k} \\ & + (1 + i_{t+k-1}) \frac{B_{t+k-1}}{P_{t+k}} + \Pi_{t+k}^r + \frac{T_t^n}{P_t}, \end{aligned} \quad (18)$$

where $M_t (B_t)$ is the household's nominal holding of money (one-period bonds). Bonds pay a nominal interest rate i_t and $1 + i_t$ represents the gross nominal interest rate. W_t denotes the nominal wage. Real profits received from firms are equal to Π_t^r and T_t^n are nominal lump-sum taxes or transfers. Note that $\Delta M_{t+k} = M_{t+k} - M_{t+k-1}$ holds.

The household seeks to maximize (17) subject to (18) and the following cash-in-advance constraint:

$$P_{t+k}C_{t+k} = M_{t+k}, \tag{19}$$

i.e., nominal consumption expenditures are not allowed to exceed the nominal money holdings of the household. This expression is given in real terms and under consideration of $y_t = c_t$ in Sect. 2. In can be easily shown that the optimality condition for the demand of real money holdings (in daily magnitudes) is given by:

$$m_t^r = \frac{M_t}{P_t} = \frac{1}{\psi}(\sigma y_t - \beta(h_d)i_t). \tag{20}$$

Appendix B: Time series for the Argentine consumer price index (CPI)

See Fig. 1.

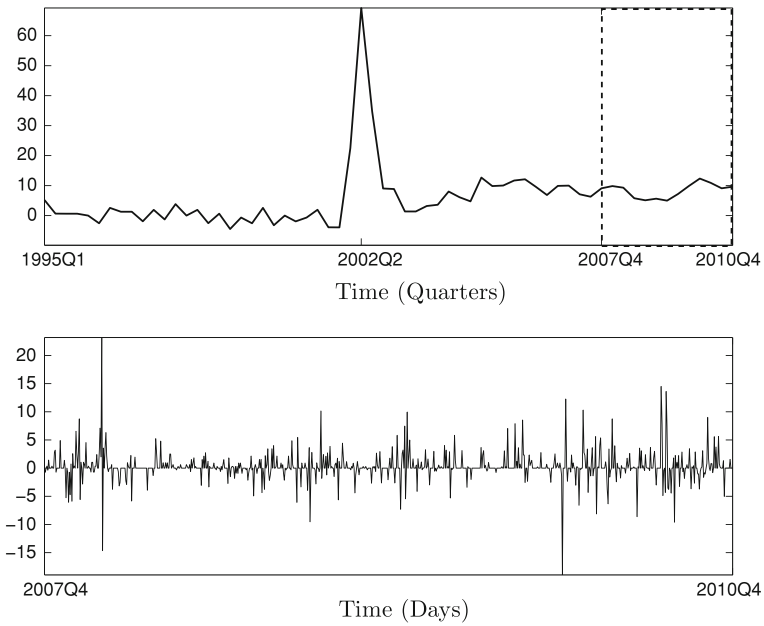

Fig. 1 The upper panel depicts the development in the Argentine CPI in *quarterly* magnitudes from 1995Q1 until 2010Q4. The lower panel depicts the corresponding Argentine CPI in *daily* magnitudes (scraped data index) from 2007Q4 until 2010Q4. Quarterly data is taken from Datastream®. Daily data is provided by www.inflacionverdadera.com

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