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estimating a high-frequency new keynesian phillips curve

by Steffen Ahrens and Stephen Sacht

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 well-known 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 monthly and quarterly equivalences and be employed for the analysis of monetary and fiscal policy. 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 NKM, GMM.
JEL classification: C26, E31.

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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 macro literature. Economists have been interested in the accuracy of the NKPC to resemble real time data and the information on the structural parameters, especially 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, p. 1140) 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, even going back only 50 or 100 years from now many periods with different economic conditions are covered such as the Great Depression, the high inflation periods after the oil price shocks, low inflation periods such as the Great Moderation, and the current financial and economic crisis, i.e. the so called Great Recession. Assuming the behavioral deep parameters to remain constant over periods, which are characterized by that 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 estimating the NKPC for a higher frequency. We are then able to take the estimation results for the daily Calvo parameters and transform them into pseudo-quarterly observations by simply applying the rules described in Franke and Sacht (2010, p. 6) in order to derive a high-frequency NKPC, i.e. we account for the fact that the frequency-dependent parameters of the model should be suitably adjusted. In particular we use daily 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 three years of observations, but contains 830 data points. This is very close to the value suggested by Lindé (2005). Furthermore, in only one year from now, we will be able to clearly supersede the mark of 1,000 observations. Our high-frequency approach gives us several advantages over the standard quarterly analysis of the NKPC. First, we are

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

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. This estimate in daily frequency can be transformed into lower frequency equivalents and (in future research) be used to calibrate business cycle models on a monthly or quarterly frequency.

This study focuses on Argentina for the following reasons. First, Argentina is the only country 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, p. 18) identify two substantial structural breaks in inflation due to 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, so called scraped price indices (i.e. indices calculated from online prices as by the Billion Prices Project) report more realistic values for Argentine inflation compared to official statistics. For instance, from October 2007 to August 2010 the scraped price data inflation rate was 17.1 % compare to 7.6 % reported by the National Statistics Institute (Cavallo (2011, p. 21)).[3]

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 (2011, p. 30). 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 adjustments compared to the standard calibrations for the United States or the Euro Area. Second and most important, an average price stickiness of two thirds of a quarter for Argentina means that at a quarterly frequency a flexible price model has to be applied to analyze the effects of policy measures. 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 (2010) seems more appropriate.

[2]See Figure 1 in the Appendix and D'Amato et al. (2007, p. 20).

[3]Cavallo (2011, p. 5) 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.

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), was done by Flaschel et al. (2008), Franke and Sacht (2010) as well as Anagnostopoulos and Giannitsarou (2010) in the first place. Franke and Sacht (2010) show that by diverging from the standard assumption of the baseline period length to be a quarter, dramatical changes in the dynamic properties of the model follow. In particular, the authors state that while determinacy of the model remains unaffected, the impulse response functions can differ in a quantitative and qualitative 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, pp. 10-11) shows that the frequency of price changes is considerably higher in high-frequency studies compared to the traditional monthly or quarterly consumer price index analysis undertaken by statistical agencies. Furthermore, Ellis (2009, p. 11) shows that lower frequency data tends to overstate the true price stickiness. Abe and Tonogi (2010, pp. 725-726)) strongly support this conclusion for the Japanese market. Additionally, Kehoe and Midrigan (2007, p. 9) find very short average price stickiness spells for suburban Chicago and Cavallo (2011, p. 30) 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 section 3 we first describe the data as well as the estimation technique and present the empirical results. We discuss the implications of our results for monetary and fiscal policy analysis in Argentina. Finally, section 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, Chapter 3) and Walsh (2010, pp. 379-381) among others. The standard purely forward-looking NKPC reads as follows:[4]

$$\pi_t = \beta(h_d)E_t\pi_{t+1} + \frac{(1-\theta(h_d))(1-\theta(h_d)\beta(h_d))}{\theta(h_d)}(\mu + mc_t^r), \qquad (1)$$

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 *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.[5]

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 equation (1) it follows that the *daily* discount parameter and the expectations operator are given by $0 < \beta(h_d) < 1$

[4]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 Christiano 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. Consequently, we restrict ourselves to the purely forward-looking NKPC.

[5]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 the next *day* instead.

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, p. 2) and Franke and Sacht (2010, p. 6) 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.[6] The converse probability is then just given by

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

Hence given the probability of not resetting the price within a quarter, $\theta(h_q)$, the corresponding probability in e.g. daily magnitudes is just $\theta(h_i = h_d)$ where $h_i = h_d$ is equal to $1/75$ since a quarter consists of 75 days on average (excluding weekends).[7] 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 chapter 3 - e.g. 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 (3). Furthermore we are interested in the question what the value of the *quarterly* Calvo parameter is since in a New Keynesian models the underlying time period is a quarter by assumption. Keeping equation (3) in mind we are going to address this issue in chapter 3.

[6] 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).

[7] 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.

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. Hence

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

However, while domestic prices are given on daily magnitudes by the Billion Prices Project at MIT Sloan (a detailed description of the data is provided in chapter 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.[8] Note again that up to this point the structure of the NKPC does not differ in closed and open economies (e.g. Galí (2008, p. 163) or Clarida et al. (2002, p. 890)). The last term in (1) can be substituted by the expressions for the domestic (y_t) and foreign (y_t^f) output gap:

$$y_t = x_t - \tilde{x}, \tag{5}$$
$$y_t^f = x_t^f - \tilde{x}^f, \tag{6}$$

where x_t (x_t^f) and $\tilde{x}$ ($\tilde{x}^f$) stand for domestic (foreign) output level and potential output (i.e. the level of output in the absence of nominal rigidities) respectively. Hence,

$$\pi_t = \beta(h_d)E_t\pi_{t+1} + \frac{(1-\theta(h_d))(1-\theta(h_d)\beta(h_d))}{\theta(h_d)}[(\sigma_\alpha + \eta)y_t - (\sigma_\alpha - \sigma)y_t^f], \tag{7}$$

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, p. 163)). The parameter η denotes the substitution elasticity of labor. As we can see from (7) the problem concerning the frequency remains since data on both output gaps is also not available on a daily basis.

Therefore, in order to get an appropriate expression for a daily NKPC, we make use of the (log-linearized) terms of trade, i.e. the terms of trade

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

which is defined by the price of foreign goods in terms of home goods:

$$s_t = e_t + p_t^f - p_t, \tag{8}$$

where e_t denotes the bilateral nominal exchange rate and p_t^f is the price for foreign goods. Furthermore, we claim that there exists a relationship between the terms of trade gap and both (domestic and foreign) output gaps

$$\frac{1}{\sigma_\alpha}(s_t - \tilde{s}(x_t^f)) = y_t - y_t^f, \tag{9}$$

where $\tilde{s}(x_t^f)$ stands for the terms of trade in the steady state which is a non-linear function of the foreign output level and the technology parameter a_t (Clarida et al. (2001, pp. 250-251) and Clarida et al. (2002, p. 890)).[9] On a daily basis, the impact of the foreign economic activity are negligibly small and therefore the expressions for y_t^f and $\tilde{s}(x_t^f)$ can be omitted.[10] By applying (8) and (9) on (7) we are able to derive an open economy NKPC which depends on the terms of trade:

$$\pi_t = \beta(h_d)E_t\pi_{t+1} + \frac{(1-\theta(h_d))(1-\theta(h_d)\beta(h_d))}{\theta(h_d)}\psi s_t, \tag{10}$$

where $\psi = 1 + \frac{\eta}{\sigma_\alpha}$.[11] 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

[9]Given the definition of the terms of trade in (8) and under the assumption of complete securities markets, equation (9) implies that relative consumption and hence (in a general equilibrium framework) the output gap relation across countries is proportional to the terms of trade (see also Lubik and Schorfheide (2007, p. 1072) and Appendix A of Galí and Monacelli (2005) for a proof). For an empirical discussion see Chari et al. (2002) among others.

[10]Figure 2 in the Appendix shows the percentaged deviations of the Brazilian, US and a multicountry output from its trend in quarterly and daily magnitudes. For simplicity we assume that the value of a daily foreign output gap is uniformly distributed over a quarter, i.e. between two data points: at the beginning of the quarters respectively ($y_t^{f,q} - y_{t-1}^{f,q}$ where the superscript q denotes a quarter). Recall that we consider 75 data points for the inflation rate (since a quarter consists on 75 days). Hence the mean of the daily output gap is just given by $y_t^f = (y_t^{f,q} - y_{t-1}^{f,q})/75$. It can be seen from Figure 2 that over the observed time horizon the deviations in the corresponding foreign output gaps are not greater than ± 0.1 %. Within our theoretical investigation of the high frequency NKPC we claim that $y_t^f \leq \nu$ where is ν is a arbitrage small number. For the degree of openness $\alpha < 0.2$ holds, based on our calibration (cf. section 3). These observations hold also for the foreign output (not shown in Figure 2).

[11]Furthermore, due to our empirical analysis we account for a white noise innovation to inflation only. Therefore no technology shock appears ($a_t = 0$).

version of the NKPC. Under consideration of the (re-arranged) uncovered interest parity

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

where i_t (i_t^f) denotes the domestic (foreign) nominal interest rate, this leads to the following expression denoted as *Type I* NKPC:

$$\begin{aligned} \pi_t = \beta(h_d) E_t \pi_{t+1} \quad &+ \quad \frac{(1-\theta(h_d))(1-\theta(h_d)\beta(h_d))}{\theta(h_d)} \\ &\cdot \quad \psi(E_t e_{t+1} + i_t^f - i_t - p_t + p_t^f). \end{aligned} \quad (12)$$

Within this specification the driving forces of domestic inflation are the expected bilateral nominal exchange rate, the domestic nominal and foreign interest rates, and the domestic and foreign price level. For the latter, we follow Monacelli (2004, p. 201) and set $p_t^f = 0$ 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."[12] Note that the NKPC in (12) depends on the domestic and foreign interest rates (and on the terms of trade in general) which defines the interest rate gap, i.e. $\delta i_t = i_t^f - i_t$. Suppose $\delta i_t > 0$. 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 this means in a DSGE context that $x_t^f < x_t$ via the real interest channel. However if $x_t - x_t^f > 0$ the domestic country is running 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 (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 e.g. 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."

A specification where domestic inflation is driven by the real bilateral exchange rate only is obtained by considering (8) and (10). The so called *Type II* NKPC, which we use for robustness checks, is then given by

$$\pi_t = \beta(h_d) E_t \pi_{t+1} + \frac{(1-\theta(h_d))(1-\theta(h_d)\beta(h_d))}{\theta(h_d)} \psi(e_t - p_t). \quad (13)$$

[12]This assumption is consistent with our empirical observation regarding the foreign output gap(s) in Figure 2.

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 formula (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. In the first place we therefore are going to estimate both types of NKPC, (12) and (13), via GMM under consideration of the data in daily magnitudes.

3 Empirical Analysis

In this section we analyze the empirical implications of the high-frequency NKPC for the adjustment speed of prices on a daily basis. We define Argentina as the domestic economy. Following D'Amato and Garegnani (2009, p. 5), 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, p. 2)). Nevertheless, to test for robustness, we also report results for the United States being the foreign country.

3.1 Data

The data set comprises daily observations for Argentina, Brazil and the United States from 12-03-2007 to 04-02-2011. For proper identification, mnemonic codes are presented 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 conceptually more appropriate implicit GDP-deflator due to reasons of data availability,[13] an approach common in the literature. Recent examples are, among others,

[13]GDP data is collected on a quarterly basis and consequently, the implicit GDP-deflator is quarterly as well.

D'Amato and Garegnani (2009, p. 4), Mihailov et al. (2011, p. 66), Ramos-Francia and Torres (2008, p. 276), and Yazgan and Yilmazkuday (2005, p. 3). Moreover, Nason and Smith (2008, p. 388) apply both measures to test for robustness and find the differences in performance to be negligible for the United States, while Holmberg (2006, p. 10) 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.

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 is collected on a daily basis from large supermarkets in the metropolitan area of Buenos Aires.[14] All remaining data is taken from Datastream®. In particular, we apply the Argentine Peso to EURO and US dollar exchange rates (TEARSSP) and (TDARSSP), respectively. Additionally, we derive the Argentine Peso to Brazilian terms of trade 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 Liquidação 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) and we make use of the Argentine 3-month Buenos Aires Interbank Offer Rate (AGIBK3M).

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 $\rho(h_q) = 0.013$.[15] All remaining parameters are calibrated according to Escudé (2009, pp. 85-86), 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

[14] In particular, the prices of 150 products are checked online every day. This methodology is sufficient since 100 percent of all products in Argentine supermarkets can also be found online (Cavallo (2011, p. 20)). For a thorough discussion of the methodology of the Billion Prices Project at MIT Sloan, refer to Cavallo (2011), Cavallo and Rigobon (2011), www.inflacionverdadera.com, and www.thebillionpricesproject.com.

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

Table 1: Calibration

	σ	χ	η	γ	α	$\rho(h_q)$
Escudé (2009)	1.902	0.700	1.175	0.990	0.139	0.013
Galí and Monacelli (2005)	1.000	3.000	1.000	1.000	0.100	0.013
Escudé (2007)	4.960	1.000	1.194	3.500	0.112	0.013

intertemporal elasticity of labor $\eta = 0.7$. Domestic and foreign goods 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 degree of openness is calibrated to match Argentina's share of foreign goods in consumption, i.e. $\alpha = 0.134$. For robustness, we juxtapose the results for an earlier calibration from Escudé (2007, pp. 73-74) for Argentina and an alternative calibration from Galí and Monacelli (2005, p. 723), which represents a standard calibration in the literature. The parameter values are summarized in Table 1.

3.3 Estimation Methodology

The empirical analysis rests on the high-frequency NKPC given by the equations (12) and (13). By substitution of the day-by-day expectations error $\epsilon_t = \beta(h_d)(E_t[\pi_{t+1}] - \pi_{t+1})$ we obtain a regression equation of the form

$$\pi_t = \beta(h_d)\pi_{t+1} + \frac{(1-\theta(h_d))[1-\theta(h_d)\beta(h_d)]}{\theta(h_d)}\psi\xi_{j,t} + \epsilon_t, \tag{14}$$

with $\xi_j = \{\xi_1, \xi_2\} = \{(e_{t+1} + i_t^f - i_t - p_t), e_t - p_t\}$. McCallum (1976, p. 44) 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 $E_t[\epsilon_t \mathbf{z}_t] = 0$. Applying this condition to equation (14), we obtain

$$E_t\left[(\theta(h_d)\pi_t - \theta(h_d)\beta(h_d)\pi_{t+1} - (1-\theta(h_d))(1-\theta(h_d)\beta(h_d))\psi\xi_{j,t})\,\mathbf{z}_t\right] = 0, \tag{15}$$

with the instrument set $\mathbf{z}_t = (\pi_{t-1}^{food}, \pi_{t-2}^{food}, p_{t-1}^{food}, p_{t-2}^{food}, i_{t-1}^{3M}, i_{t-2}^{3M}, e_{t-1}, e_{t-2})'$ comprising each two lags of the food and beverage based consumer price inflation π^{food}, food and beverages based consumer price index p^{food}, the Argentine three-month interbank interest rate i^{3M}, and the respective exchange rate e.[16] According to McCallum (1976, p. 44) such an orthogonality con-

[16] To test the validity of the instrument set, we perform a standard J-Test for overidentifying restrictions. In each of the following exercises, we cannot reject the Null hypothesis that the J-statistic is sufficiently close to zero at the 99% confidence level therefore, evaluating the instruments valid.

dition can be consistently estimated with an instrument variable technique. The latter is standard in the literature since the prominent contribution of Galí and Gertler (1999), we apply Hansen's (1982) Generalized Method of Moments to estimate the structural parameter $\theta(h_d)$.

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, p. 5). 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 $\theta\,(h_d)$ for all three cases are summarized in the first column of Table 2.

Table 2: Price Adjustment Frequency in Argentina

Foreign Country	Calvo parameter			Average duration of fix prices in...		
	$\theta\,(h_d)$	$\theta\,(h_m)$	$\theta(h_q)$	...days	...months	...quarters
Multicountry						
Type 1	0.9880	0.7000	0.0999	83	3.33	1.11
	(0.0723)					
Type 2	0.9813	0.5413	0*	53	2.13	0.71
	(0.0850)					
Brazil						
Type 1	0.9805	0.5119	0*	51	2.04	0.68
	(0.0765)					
Type 2	0.9804	0.5094	0*	51	2.04	0.68
	(0.0744)					
United States						
Type 1	0.9852	0.6310	0*	68	2.71	0.90
	(0.0611)					
Type 2	0.9776	0.3275	0*	45	1.79	0.60
	(0.0740)					

Notes: The parameter $\theta(h_d)$ is estimated from the orthogonality conditions given by (15). 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 $j = m, q$. The superscript * implies that prices change at least once within every quarter, i.e. the quarterly Calvo parameter cannot be determined (since $\theta(h_q) < 0$). In this case we set $\theta(h_q)$ equal to its lower bound (see also footnote 17). The standard errors are given in brackets. We apply a 12-lag Newey-West covariance matrix.

Note up front that the estimates for $\theta(h_d)$ lie remarkably close to each other in an interval between $[0.9776; 0.9880]$, 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. 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 Table 2, depending on the case considered, prices are fixed between 45 and 83 days. The average duration implied by the high frequency NKPC over all cases considered lies at approximately 67 days for the Type I and at 50 days for the Type II high-frequency NKPC. These results are close to microeconomic evidence on high-frequency pricing in Argentina provided by Cavallo (2011, p. 30), who reports that prices remain unchanged for about 66 to 83 days.

One of the major contributions of this paper is that we can use the daily information to derive lower frequency information, such as monthly or quarterly by simply employing the daily point estimate to equation (3).[17] The results for monthly and quarterly conversion are given in the second and third and fifth and sixth column of Table 2. For an example of the conversion, let us look at the first line of Table 2. According to equation (3), a daily probability of not being able to reset the price $\theta(h_d) = 0.9880$ is equivalent to a monthly probability of not being able to reset the price of $\theta(h_m) = 0.7000$, which is equivalent to an average duration of fixed prices of a little more than three months.

The price adjustment speed estimated from the data implies that a quarterly Calvo parameter can only be determined for very few cases. This is due to the fact that the Calvo parameter in general reflects a probability and therefore is bounded between zero and unity. Given that prices e.g. are fixed on average for about two months implies that on a quarterly basis prices change twice within every period. This, however, is equivalent to a quarterly flex-price model.[18] Therefore, to analyze monetary and fiscal policy in Argentina on a quarterly frequency, a flex-price model would be necessary to

[17]We exclude weekly because it is of less interest for the applications we aim to provide calibration for.

[18]In this case applying equation (3) based on our estimated value for $\theta(h_d)$ results in a negative value for $\theta(h_q)$. 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.

mimic the Argentine economy. In order to analyze the effects of policy measures in a staggered pricing framework instead, the model frequency should be no more than monthly. Only in the case of the Type 1 high-frequency NKPC with the multi-country approach we can also compute the quarterly equivalents. In this case the Calvo parameter $\theta(h_q) = 0.0999$ or equivalently averagely fixed prices of 1.11 quarters.

This result is in stark contrast to the standard results from Taylor (1999, p. 1020) that price changes occur on average once a year and even compared to Bils and Klenow (2004, p. 953), who find an average duration of fixed prices for a little more than one quarter. These prominent results are, however, for industrialized countries, or to be more specific, the United States. Klenow and Malin (2010, p. 10) report on micro evidence of price setting from CPI data sets as well as scanner data and show that there are more frequently price changes in Latin America and developing countries in general compared to the United States and the Euro Area. Furthermore, Cavallo (2011 p. 30) applies the methodology of Bils and Klenow (2004) to daily CPI data for Argentina, Brazil, Chile, and Colombia and generates supportive evidence of a relatively high price changing frequency in these countries compared to industrialized countries. In particular, Cavallo (2011, p. 30) reports an average duration of fixed prices of 66 days, which is close to the results given in Table 2. Therefore, qualitatively and quantitatively, our high-frequency macroeconomic NKPC procedure generates comparative results to the microeconomic methods of Bils and Klenow (2004). Additionally, Ellis (2009, p. 11) 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, pp. 725-726).

To analyze for robustness of our results, Table 3 presents the results for alternative calibrations from Escudé (2007) and Galí and Monacelli (2005). Quantitatively and qualitatively, the results are robust to the results from the standard calibration. For both alternative calibrations the estimates are marginally larger and thus even more in line with microeconomic evidence. In the case of Galí and Monacelli (2005), price duration increases by approximately one to two weeks for the consumption based CPI measure. Still, the average of fixed prices is in the close neighborhood of three months. On the other hand, the results from the standard calibration are almost not distinguishable from the case, when the calibration of Escudé (2007) is applied. Summing up, again for Argentina a quarterly model would need to be calibrated as a flex price model to analyze monetary policy or put differently, in order to study monetary and fiscal policy in a sticky price model a monthly calibration is necessary.

Table 3: Robustness with Respect to Calibration

Foreign Country	Calvo parameter $\theta\,(h_d)$	$\theta\,(h_m)$	$\theta(h_q)$	Average duration of fix prices in... ...days	...months	...quarters
Result for the Calibration of Galí and Monacelli (2005):						
Multicountry						
Type 1	0.9897	0.7419	0.2258	97	3.88	1.29
	(0.0663)					
Type 2	0.9838	0.5957	0*	62	2.47	0.82
	(0.0778)					
Brazil						
Type 1	0.9830	0.5758	0*	59	2.36	0.79
	(0.0700)					
Type 2	0.9829	0.5736	0*	59	2.35	0.78
	(0.0677)					
United States						
Type 1	0.9871	0.6787	0.0353	78	3.11	1.04
	(0.0554)					
Type 2	0.9805	0.5119	0*	51	2.05	0.68
	(0.0671)					
Result for the Calibration of Escudé (2007):						
Multicountry						
Type 1	0.9883	0.7062	0.1187	85	3.40	1.13
	(0.0714)					
Type 2	0.9816	0.5411	0*	54	2.18	0.73
	(0.0840)					
Brazil						
Type 1	0.9809	0.5213	0*	52	2.09	0.70
	(0.0757)					
Type 2	0.9808	0.5190	0*	52	2.08	0.69
	(0.0734)					
United States						
Type 1	0.9855	0.6381	0*	69	2.76	0.92
	(0.0602)					
Type 2	0.9780	0.4510	0*	46	1.82	0.61
	(0.0730)					

Notes: The parameter $\theta(h_d)$ is estimated from the orthogonality conditions given by (15). 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 $j = m, q$. The superscript * implies that prices change at least once within every quarter, i.e. the quarterly Calvo parameter cannot be determined (since $\theta(h_q) < 0$). In this case we set $\theta(h_q)$ equal to its lower bound (see also footnote 17). The standard errors are given in brackets. We apply a 12-lag Newey-West covariance matrix.

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 kind of 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 series data. Applying our method to Argentine data we find the daily Calvo parameter to lie within the narrow interval of [0.9776;0.9897], which implies averagely fixed prices of approximately two to three months. These values are well in line with micro evidence for Argentina from Cavallo (2011, p. 30). 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 one half of a quarter means that at a quarterly frequency a flexible price model has to be applied to analyze the effects of policy measures. 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 (2010) seems more appropriate. Our results turn out to be robust not only to alternative standard calibrations from the literature, but also to two different proxies for real marginal costs and alternative country data.

The question arises if the approach presented in this paper is sufficient to reproduce equivalent results for industrial countries like the US 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 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.[19] 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 appear 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, p. 29) 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 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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[19] Note that effective optimal policy in NKMs requires price stickiness regarding aggregate shocks.

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Appendix

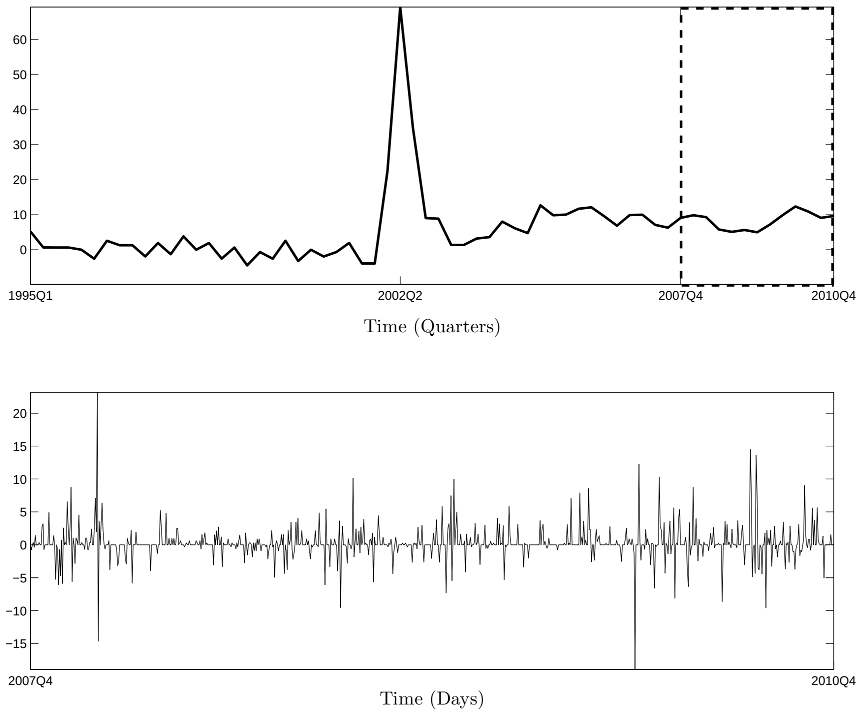

Figure 1: Time Series for the Argentine consumer price index (CPI). Note: The upper panel depicts the development in the Argentine CPI in *quarterly* magnitudes from 1995Q1 until 2010Q4. The lower panel depicts the Argentine CPI in *daily* magnitudes from 2007Q4 until 2010Q4. Quarterly data is taken from Datastream®. Daily data is provided by *www.inflacionverdadera.com.*

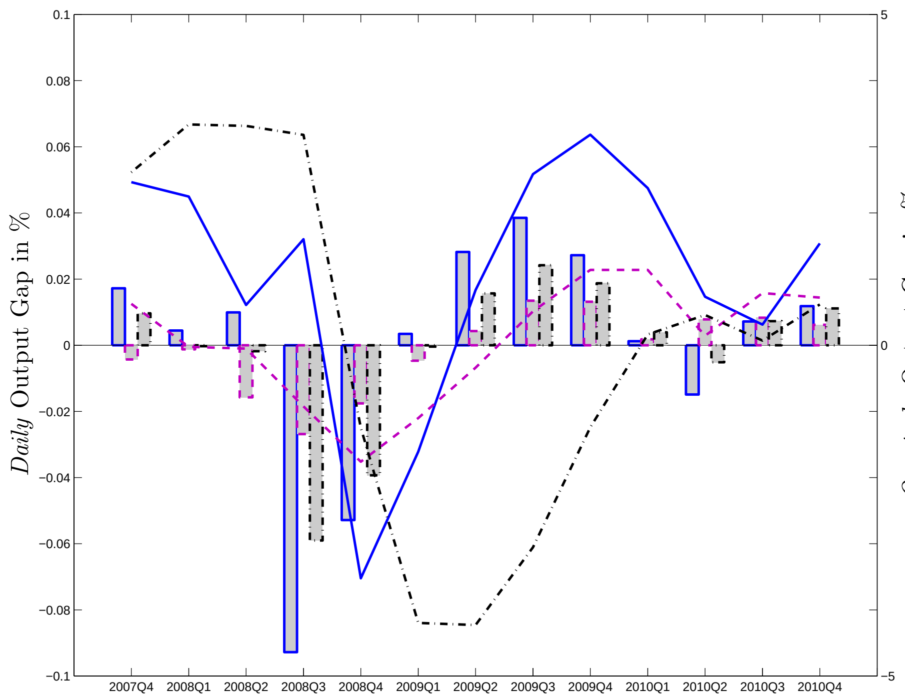

Figure 2: Development of the Brazilian, the US and multicountry output gaps in quarterly and daily magnitudes (in %) from 2007Q4 up to 2011Q1. Note: The *solid*, *dashed* and *dashed/dotted* line (bar) denote the percentaged deviations of the *Brazilian*, *US* and *multicountry* output gap from its trend in quarterly (daily) magnitudes respectively. The trend of the foreign output gap is given by applying the Hodrick-Prescott filter (with the common value for λ of 1600) based on data from 1995Q1 till 2011Q1. It goes without saying that the times series are deseasonalized. The multicountry output gap consist of one half of the value for the Brazilian and one fourth of the value of the US and EU output gaps respectively. For simplicity we assume that the value of the *daily* foreign output gap y_t^f is uniformly distributed over the whole quarter (denoted by the superscript q), i.e. $y_t^f = (y_t^{f,q} - y_{t-1}^{f,q})/75$. The data is taken from Datastream®.