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JEL classification: C26, E31.

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Steffen Ahrens* and Stephen Sacht§

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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 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 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, p. 143), 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, 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 daily 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 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. 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. This estimate in daily frequency can be transformed into their low-frequency equivalences and (in future research) be used to calibrate

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

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 done so for instance 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 (2012, p. 3)). 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.

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 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, 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

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

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, 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:[3]

[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 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. The insignificance of lagged inflation might be due to the result of the low autocorrelation of daily inflation, as can be seen from the Figure (lower panel) in the Appendix. To support this presumption we perform a serial correlation Lagrange multiplier test on inflation (Godfrey (1978) and Breusch (1979)) for lag lengths between two days and up to one 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)

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

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]

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 $0 < \beta(h_d) < 1$ 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 (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

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

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.

[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).

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).[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 section 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 equation (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 section 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. Hence

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

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 section 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 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. Hence,

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

[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.

[7] 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.

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. 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. Furthermore, we claim that there exists a relationship between the terms of trade *gap* and both (domestic and foreign) output gaps

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

where $\tilde{s}$ stands for the terms of trade in the steady state (Clarida et al. (2001, pp. 250-251) and Clarida et al. (2002, p. 890)).[8] By applying (6) on (5) we are able to derive an open economy NKPC which depends on the terms of trade and 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 money-in-the-utility function, a budget, and a cash-in-advance constraint known from the literature (see Appendix). 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:

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

[8]Under the assumption of complete securities markets, equation (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 (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.

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

where ψ is the inverse elasticity of money demand. Substituting (8) into (9) and rearranging 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. Under consideration of the (re-arranged) uncovered interest parity

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

where i_t (i_t^f) denotes the domestic (foreign) nominal interest rate. The corresponding steady state expression is given by:

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

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

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

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 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 $p_t^f = \tilde{p}_t^f = 0$ 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. $\delta i_t = \Delta i_t^f - \Delta i_t$. Suppose $\delta i_t > 0$. In this case $\Delta i_t^f > \Delta i_t$ 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, where x_t (x_t^f) denotes the domestic (foreign) level of output. 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 ($\Delta s_t = s_t - \tilde{s}_t$ 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 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".

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

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

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 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.

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, 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 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, p. 711), D'Amato and Garegnani (2009, p. 4), Mihailov et al. (2011a, p. 319), and Mihailov et al. (2011b, p. 66).[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, p. 11, fn 19)). Furthermore, the author mentions that "this limitation [of not considering the price of services] can be overcome as a growing number of firms start posting their prices online" (Cavallo (2011, 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 (2012, p. 12)). He shows that online and official estimates share a similar pattern over time, while there is a high correlation between both indices (Cavallo (2012, p. 13, Figure 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 rates (Cavallo (2012, p. 18)). To sum up, studies on Argentine inflation dynamics - like in this paper - must account for these characteristics which do not certainly hold for the US or Euro Area.

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.[11] All remaining data is taken from Datastream®. In particular, we apply the Argentine Peso to EURO and

[10]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. 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, 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.

[11]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, we refer to Cavallo (2011 and 2012), Cavallo and Rigobon (2011), www.inflacionverdadera.com, and www.thebillionpricesproject.com.

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

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).

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, 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 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 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, pp. 73-74) for Argentina. The parameter values are summarized in Table 1.

[12] 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 $h_q = 1$.

3.3 Estimation Methodology

The empirical analysis rests on the high-frequency NKPC given by the equations (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

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

with $\xi_j = \{\xi_1, \xi_2\} = \{\phi_1\left(E_t\Delta e_{t+1} + \Delta i_t^f - \Delta i_t - \Delta p_t\right) + \phi_2 i_t, \phi_1(\Delta e_t - \Delta p_t) + \phi_2 i_t\}$.[13] 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 $E_t[\varepsilon_t \mathbf{z}_t] = 0$. Applying this condition to equation (15), we obtain

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

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 US 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), p. 1250). Finally, among the potential instrument sets we choose the one set with the lowest average J-statistic over all applied estimations.

According to McCallum (1976, p. 44) 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) 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

[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.

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

Note up front that 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. Since Calvo staggering follows a Poisson process, prices are fixed on average for $\mathscr{D} = \frac{1}{1-\theta(h_d)}$ 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 for both the Type I and the Type II high-frequency NKPC. These results are fully in line with microeconometric evidence on high-frequency pricing in Argentina. In a microeconometric analysis of price changes in Argentina, Cavallo (2011, p. 30) reports that prices in Argentina remain unchanged for about 66 to 83 days.

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 weak identification of the model. In order to test for weak identification we follow Yazgan and Yilmazkuday (2005) and apply the nonlinear Anderson-Rubin (*AR*) statistic.[14] 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. This result holds true independently of whether the instruments are weak or strong.

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 quar-

[14]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).

Table 2: Price Adjustment Frequency in Argentina

Foreign Country	Calvo parameter			Average duration in...			p(J-statistic)	p(AR-statistic)	p(AR-statistic)
	$\theta\left(h_d\right)$	$\theta\left(h_m\right)$	$\theta\left(h_q\right)$	...days	...months	...quarters		χ^2	F
Multicountry									
Type 1	0.9867	0.6667	0.0002	75	3.00	1.00	0.9115	0.3818	0.3837
	(0.0203)								
Type 2	0.9867	0.6671	0.0014	75	3.00	1.00	0.9115	0.3817	0.3836
	(0.0203)								
Brazil									
Type 1	0.9867	0.6678	0.0032	75	3.01	1.00	0.8661	0.3801	0.3820
	(0.0204)								
Type 2	0.9867	0.6778	0.0035	75	3.01	1.00	0.8661	0.3801	0.3820
	(0.0205)								
United States									
Type 1	0.9855	0.6387	0^*	69	2.77	0.92	0.4525	0.3910	0.3928
	(0.0181)								
Type 2	0.9856	0.6390	0^*	69	2.77	0.92	0.4524	0.3909	0.3927
	(0.0181)								

Notes: 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 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.

terly by simply employing the daily point estimate to equation (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 equation (3), a daily probability of not being able to reset the price $\theta(h_d) = 0.9867$ is equivalent to 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 three 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.

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 due to the fact that the Calvo parameter in general reflects a probability and therefore is bounded between zero and unity. Given that prices are fixed on average for about three months implies that, on a quarterly basis, prices change only once within every period. This, however, is equivalent to a quarterly flex-price model. Therefore, to analyze economic policy in Argentina on a quarterly frequency, a flex-price model would be necessary to mimic the Argentine economy. 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.[15] This does not hold true for fiscal measures, which can easily be analyzed also in flex-price economy. Therefore, in order to analyze the effects of policy measures in a staggered pricing framework instead, the model frequency should be no more than monthly.

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 to 83 days, an inter-

[15]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), pp. 126-127).

val which encompasses the results given in Table 2. Therefore, qualitatively and quantitatively, our high-frequency macroeconometric NKPC procedure generates comparative results to the microeconometric 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 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 three 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 to zero, which is its natural lower bound.

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 over all cases considered to lie within the narrow interval of [0.9798;0.9867], 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

Table 3: Robustness Exercises

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

Notes: 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 due to the fact that in all cases prices change less than once a quarter.

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

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.

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 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 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 (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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Appendix

Time Series for the Argentine consumer price index (CPI)

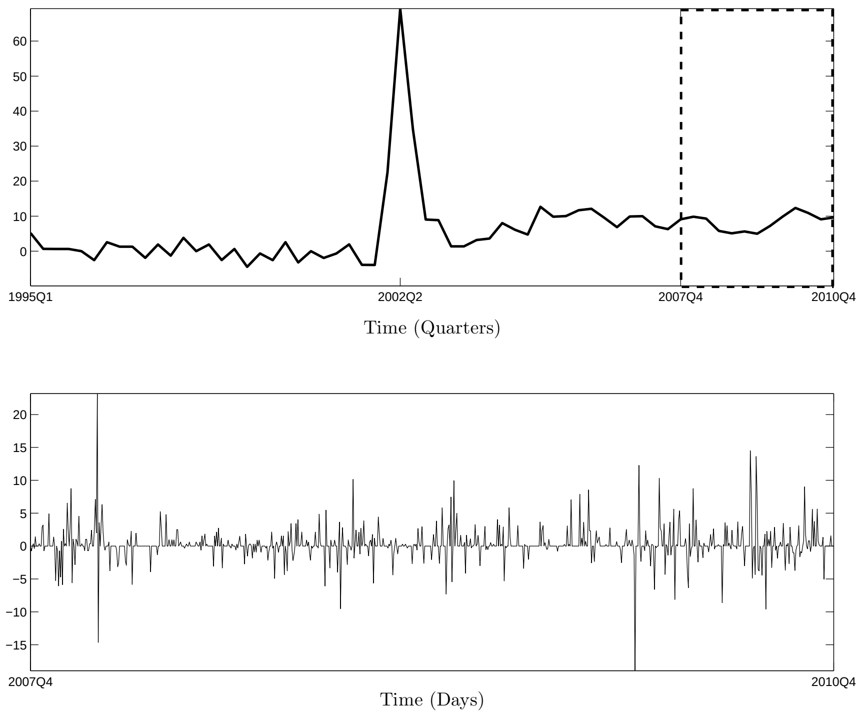

Figure 1: The upper panel depicts the development in the Argentine CPI in *quarterly* magnitudes from 1995Q1 until 2010Q4. The lower panel depicts the Argentine CPI in *daily* magnitudes (scraped data index) from 2007Q4 until 2010Q4. Quarterly data is taken from Datastream®. Daily data is provided by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

Household's Optimization Problem

Following Woodford (2003, Chapter 6) and Galí (2008, Chapter 4) 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:

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

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 discount factor. E_t stands for the expectations operator.

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

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

where M_t (B_t) is the household's nominal holding of money (one-period bonds). Bonds pay a nominal interest rate i_t. $1+i_t$ represents the gross nominal interest rate. Real profits received from firms are equal to Π_t^r. 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 section 2. In can be easily shown that the optimality condition for the demand of real money holdings (in daily magnitudes) is given by:

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