Gender Quotas, Female Representation and Public Expenditures: New Evidence from Flemish Municipalities*

John Ashworth*, Benny Geys#,‡ and Bruno Heyndels‡

** Department of Economics, University of Durham, 23-6 Old Elvet, Durham, DH1 3HY, England; email: John.Ashworth@durham.ac.uk*

 # Department of Economics, Norwegian Business School BI, Nydalsveien 37, N-0442 Oslo, Norway; email: Benny.Geys@bi.no

‡ Department of Applied Economics, Vrije Universiteit Brussel, Pleinlaan 2, B-1050 Brussel, Belgium; email: Benny.Geys@vub.ac.be; Bruno.Heyndels@vub.ac.be

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Abstract –
Over the past 15 years, various legislative steps have been taken in Belgium – including quota rules for election lists and a mandated minimal representation requirement in municipal legislative bodies – towards the equalisation of male and female representation in the political sphere. At the local level, such laws are not only exogenously imposed by higher-level governments, but also generate varying constraints across space and time. This paper exploits these legal changes to identify whether, and how, local government expenditures are affected by the gender of local politicians. The dataset includes 299 (out of 308) Flemish municipalities over the period 1998-2008. We find, in line with previous studies, that a larger share of women in the municipal population is linked to higher local government expenditures. Interestingly, however, higher female representation in the local parliament and/or government is *not* associated with higher spending (and, if anything, tends to decrease it in the short term), and does not appear to create substantive shifts in spending patterns.

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* Corresponding author: John Ashworth, Department of Economics, University of Durham, 23-6 Old Elvet, Durham, DH1 3HY, England ;tel:0044191334635; email: John.Ashworth@durham.ac.uk

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1. Introduction

Gender is often seen as an important indicator of policy preferences, with women tending to be more egalitarian, more socially aware and more favourable towards higher welfare spending (e.g., Lott and Kenny, 1999; Edlund and Pande, 2002; Ågren *et al*., 2006; Funk and Gathmann, 2008; Svaleryd, 2009; Campa, 2011; for an overview, see Croson and Gneezy, 2009). Obviously, this difference-of-opinion between the sexes can have important policy implications. First, it leads to different voting patterns between men and women (e.g., Edlund and Pande, 2002; Funk and Gathmann, 2008), which can cause substantial long-term shifts in public spending patterns. One illustration of such effect is provided by studies uncovering the economically and statistically substantive effect of the introduction of female suffrage on public policies (e.g., Lott and Kenny, 1999; Aidt and Dallal, 2008; Miller, 2008). Second, to the extent that female politicians have the same policy preferences as the general female population (for evidence that this need not necessarily be the case, see Ågren *et al*., 2006) and politicians do not merely (attempt to) represent the median voter (see Osborne and Slivinski, 1996; Besley and Coate, 1997; as compared to Downs, 1957), a similar shift in public policies might be expected when women increase their representation in government/parliament (e.g., Chattopadhyay and Duflo, 2004; Rehavi, 2007; Chen, 2010; Campa, 2011; Clots-Figueras, 2011a, b).

This article contributes to the latter strand of literature by analysing whether an increase in the political representation of women at the local government level in Flanders affects municipal expenditure levels (in general and across policy fields). Clearly, however, simply comparing

* Corresponding author: John Ashworth, Department of Economics, University of Durham, 23-6 Old Elvet, Durham, DH1 3HY, England ;tel:0044334635; email: John.Ashworth@durham.ac.uk

jurisdictions with varying shares of female representatives is inappropriate. As unobserved variations across jurisdictions' electorates may be responsible for differences in *both* female representation *and* public policies, such approach will fail to identify the causal effect of female representation (e.g., Funk and Gathmann, 2008; Svaleryd, 2009; Geys and Revelli, 2011). In a path-breaking contribution, Chattopadhyay and Duflo (2004) address this critical endogeneity concern by exploiting a natural experiment in India, in which Village Council Head positions in a random selection of villages where reserved for a woman. Rehavi (2007) – using US state-level data – and Clots-Figueras (2011a, b) – using Indian state- and district-level data – exploit the inherent randomness of very close elections for identification purposes, whereas Chen (2010) – using panel data from 103 countries over the 1970-2006 period – exploits the temporal variation in the presence/absence of gender quotas in an IV-estimation approach. Assuming gender quotas have no direct effect on policy outcomes, the latter is a feasible strategy given that gender quotas have been shown to cause higher female representation, even controlling for temporal trends and changing cultural attitudes (e.g., De Paola *et al*., 2010; Campa, 2011). All five studies find a significant impact of female representation on policy outcomes.

Interestingly, studies in settings where the electoral system is based on proportional representation have thus far provided, at best, mixed evidence. On the one hand, Svaleryd (2009) finds some effect of female representation on the relative share of public spending allocated to childcare and education relative to elderly care in Swedish municipalities. On the other hand, Campa (2011) finds no effect of gender quota (as an indirect measure of female representation) on the size and composition of public spending using a regression discontinuity design in Spanish municipalities. In this article, we analyse local governments in the Flemish Region of Belgium and exploit the introduction of a series of legislative initiatives intended to enhance the share of female politicians active at the local government level for identification purposes. First, quota rules were established in 1994 specifying that at most three quarters of the positions on a party list were allowed to be taken up by candidates of a specific sex (reduced to two thirds of the list in 2000 and half of the list in 2006). Second, a mandated minimal representation requirement was imposed as of the October 2006 municipal elections, specifying that at least one female politician should be present in the main municipal legislative body (called the College of Mayor and Aldermen, or 'College') or be appointed head of the local public welfare agency. Hence, relative to earlier work, we analyse a different institutional setting and we combine information from gender quota effects (as in Chen, 2010; Campa, 2011) with that from mandated representation requirements (as in Chattopadhyay and Duflo, 2004).

Three characteristics of the legal changes introduced make them particularly interesting for our purpose. First, quota rules aim to increase female representation on party lists, but this imposes varying constraints across municipalities depending on the historical presence of female politicians. The ensuing variation across both space and time (given the changes in the quote rule over time) in how women's representation adjusts to these quota rules can be usefully exploited in our analysis. Second, more than 80 (out of 308) municipalities failed to have *any* women in the College prior to the 2006 elections. Hence, the imposition of a mandated minimal representation requirement in 2007 constituted a substantive shock and a binding constraint for many municipalities. Finally, and crucially, all changes to the local election laws were imposed by higher-level governments upon the Flemish municipalities, and can thus reasonably be treated as exogenous to local politics (for a similar approach in different contexts, see Geys, 2006; Fiva and Folke, 2011).

Our results – employing annual data on local public expenditure in Flemish municipalities over the period 1998-2008 – indicate that a larger share of women in the municipal population is linked to higher local government expenditures in the long run. This is in line with a substantial literature illustrating that the different voting patterns of men and women significantly affect public spending patterns (e.g., Aidt and Dallal, 2008; Funk and Gathmann, 2008). Interestingly, however, a similar effect does *not* appear when concentrating on female representation in municipal political bodies – i.e., parliament, government and the mayor position. Most estimates indeed suggest a negative (and, at best, short-run) effect of female representation. Similarly, we uncover only weak evidence that increased female representation is related to shifts in local public expenditures across policy domains. One interpretation – consistent with Campa (2011) – is that female politicians at the local level in Flanders are unable to affect public policies in their desired direction (unless they are sufficiently established and dominant; see Svaleryd, 2009) – possibly because they are taken insufficiently seriously due to the positive discrimination element inherent in the quota legislation (e.g., Reynaert *et al*., 2005). An alternative explanation is that female politicians may not share the same preferences as their non-political counterparts (for empirical evidence in this direction, see Ågren *et al*., 2006), but simply behave as … politicians.

The remainder of the paper is structured as follows. Section 2 presents the institutional background of Flemish elections, the legislative changes imposed since 1994 and descriptive statistics on female representation. Section 3 considers our empirical approach and gives the main findings. Finally, section 4 offers a concluding discussion.

2. Institutional Setting and Data

2.1. Local government and electoral system

Flemish municipalities are politically organized according to a parliamentary system consisting of the local council (the legislative body or 'parliament') and the College of Mayor and Alderman (the executive body or 'government'). The size of both political bodies depends directly on the size of the municipality according to a rule set by federal legislation and which has remained essentially unchanged since 1836. Municipalities have substantial autonomy on the revenue side of their budget (e.g., Ashworth *et al.*, 2006; Goeminne *et al.*, 2008), as well as significant spending responsibilities in, for instance, public administration, education, local infrastructure, public safety, social services, and cultural and (local) environmental policies. All decisions with respect to these policy areas are taken by majority vote in the council. The party (or parties) that holds a majority position in the council after a municipal election (minority governments are very rare) thus possesses all political power in the municipality.

Elections are held during the first week of October every six years. Each party thereby presents a list containing at least one candidate and at most a number of candidates equal to the number of council seats available in a given municipality (ranging from 7 to 55 seats). Faced with these party lists, voters can either cast a 'party vote' for one party (which implies agreement with the order of candidates as presented by the party, see also below), or one or more preferential votes for candidates within one particular party. Casting votes for (candidates of) multiple parties is not allowed within the same election. Subsequently, seats are allocated to parties according to a system of proportional representation (i.e., highest averages 'Imperiali') and each party distributes its seats according to the number of preferential votes each candidate obtains, albeit with a twist. Candidates are selected into the local council on behalf of their party if they receive at least as many preferential votes as the party's vote total divided by the number of seats it obtains. This is usually, at best, only achieved by the first candidate on the list. For the remaining seats, the party adds its 'party votes' to the candidates in order of the party list, until they reach the required number of votes. Hence, party lists and 'party votes' play a crucial role in the intra-party seat allocation, and female representation in the council (and/or College) will be critically determined by female candidates' positions on their parties' lists.[1] We return to this below, as it will play a central role in the effect of gender quotas.

2.2. Female Representation

We collected data on the composition of municipal parliaments and governments for all 308 Flemish municipalities from 1998 onwards. This comprises of three electoral periods (1995-2000, 2001-2006 and 2007-2012).[2] These data were obtained from the Flemish government and cross-validated with information provided in the annual "Gemeentelijk Zakboekje" (which provides information on municipal governments). Politicians' gender was inferred from their name. In case of uncertainty (e.g., Kim can be a male and female first name in Belgium), we looked up the politicians' picture on the municipal website. Based on these data, we calculated the percentage of female politicians in the council and College, and generated an indicator variable equal to 1 for the presence of a female Mayor.

As can be seen in Table 1, just over 35% of local council members in Flemish municipalities were female following the 2006 municipal elections. Their representation in executive positions stood at just under 30%, and just under 10% of all mayors were female. These numbers represent a substantial increase over the beginning of the post-war period. After

[1] If the party runs out of 'party votes' before all its seats are allocated, allocation of the remaining seats occurs based solely on candidate's preferential votes. This, however, was a fairly infrequent event until – in 2006 – the value of the 'party vote' was reduced. Currently, each 'party vote' allows the party to add only half a vote to its preferred candidates. Clearly, this makes the party list somewhat less decisive.

[2] We start in 1998 as information on our central dependent variables (level and composition of public expenditures; see below) was unavailable prior to this year.

WWII, less than 4% of local councillors were female and less than 1% of local government executive positions were taken up by women (Reynaert *et al*., 2009). By the 1976 municipal elections, the share of female councillors had risen to roughly 13% and Colleges of Mayor and Alderman consisted, on average, of 10% women. This then remained roughly stable until the early 1990s, when a number of legislative initiatives – discussed in more detail below – generated a renewed upward trend (Reynaert *et al*., 2009). Even so, however, the share of female representatives in Flemish local politics still falls substantially short of their population share (just over 50% on average, and ranging from 46.34% to 54.51% of the municipal population). Moreover, as can be seen from Table 1, the variation in female representation across Flemish municipalities remains very high. In some municipalities, women are hard to find in either the legislative body (i.e., the council) or the executive body (i.e., the College), while in others they make up a substantial majority.

Table 1: Mean share of women in Council and College, and share of female Mayors (1994-2012)

	1994-1999	***2000-2006***	***2006-2012***
Council	20.89 [0 – 42.86]	27.14 [7.70 – 57.14]	35.19 [8.33 – 64.29]
College	16.24 [0 – 75.00]	19.51 [0 – 80.00]	29.56 [0 – 100.00]
Mayor	5.52	7.47	8.77

Note: Mean values, with minimum and maximum values between square brackets. For ‘Mayor’, we present the share of municipalities with a female mayor.

As mentioned, the increase in female representation since the early 1990s is at least in part driven by a series of legislative initiatives. The first of these – introduced with the 1994 municipal elections – intended to increase female representation indirectly by establishing quota rules on the share of single-sex candidates on party lists. Particularly, at most three quarters of the party list could be made up of male or female candidates (reduced to two thirds of the list in 2000). Failure to comply implied ineligibility of the entire party list, while a failure to find sufficient candidates of the minority sex meant the party was forced to compete with an incomplete list in the election. The idea was that a higher share of women on the party lists is a prerequisite for their higher representation in the local council and College. Nevertheless, in practice the effect remained relatively modest – especially as regards the College (see Table 1) – as no restriction was made on the placement of candidates of different sexes across the list. That is, simply filling up the bottom places with female candidates (or leaving them empty) perfectly complied with the legal framework imposed. As party lists play a central role in Belgian election outcomes (see above) due to the high share of party votes in all votes cast, this meant that a significant hurdle against increased female representation persisted.[3] The second legislative initiative therefore took a more direct approach, and imposed a minimal representation requirement starting with the October 2006 municipal elections. Particularly, it was mandated that – independent of the size of the municipality – both genders should be represented in the municipalities' main executive positions (i.e., the College *or* chairing the local office of the public welfare agency; OCMW). While this new legislation had only a limited effect on the share of women in the council, it had a strong effect on the share of women in the College (see Table 1). Note that some municipalities still had no women in the College, but these all had a female head of the public welfare agency.[4]

It is important to highlight here that local-level election laws in Belgium are determined by higher-level governments. The Belgian federal government held this legislative power until 2004, when it was transferred to regional governments (i.e., in Flanders and Wallonia). Although several politicians hold office simultaneously at the federal/regional and local level, this is unlikely to substantially affect local-level election laws in any particular direction. The

[3] Reynaert *et al*. (2009) quote a study by Johan Ackaert finding that the probability to be elected for male candidates remained twice that of female candidates in 1994 (29% versus 15.3%). They also note that the average number of preferential votes is 71% higher for male candidates compared to female ones.

[4] The share of women in the council increased as well, but this was predominantly due to a second tightening of the quota requirement. Party lists now had to have a 50-50 division according to sex, and the first three positions were not allowed to have candidates of the same sex (effectively mandating that at least one had to be female).

reason is that their individual influence is likely to be small, and their preferences concerning the 'optimal' arrangement for legislation "very likely to be conflicting (not only between municipalities, but also between different parties within the same municipality)" (Geys, 2006, 295). Consequently, and crucially, the legislative changes discussed above can reasonably be treated as exogenous shocks to the local polities (see also Geys, 2006; Fiva and Folke, 2011). Moreover, the constraints they imposed varied significantly across municipalities. For instance, roughly 27% of the Flemish municipalities (i.e., 83 out of 308) did not have *any* women in the College prior to the 2006 elections. Similarly, in some municipalities all parties already met the quota restrictions prior to their introduction in 1994, while this was not the case in (many) others. This variation in the stringency of the gender-legislation constraints across both space and time (given the change in the quote rule over time) is exploited in our analysis.

3 Empirical Analysis

3.1. Empirical model and methodology

To examine the effects of female representation on expenditure patterns, we examine annual data for the period 1998 to 2008 for 299 Flemish municipalities. As mentioned, this time period includes two municipal elections (in 2000 and 2006) and two major legislative changes (i.e., strengthened quota regulations and mandated minimal representation requirements). For the main thrust of the study, we rely on an error correction model (ECM), which allows consideration of both the long-term effects and short-run deviations (i.e., possible shocks from the long run equilibrium). The general formulation of our specification closely follows earlier work on Flemish and Norwegian municipalities by, for instance, Heyndels (2001), RattsØ and Tovmo (2002), Ashworth and Heyndels (2005) and Ashworth *et al*. (2005), but augments this with a set of gender-related variables:

$$\Delta EXP_{i,t} = \beta_0 + \beta_1 EXP_{i,t-1} + \beta_2 X_{i,t-1} + \beta_3 POL_{i,t-1} + \beta_4 GEN_{i,t-1} + \beta_5 \Delta X_{i,t} + \beta_5 \Delta POL_{i,t} + \beta_7 \Delta GEN_{i,t} + e_{i,t} \quad (1)$$

Specifically, $EXP_{i,t}$ gives total local public expenditures (per capita, in euro) for municipality i (=1, …, 299) at time t (= 1998 to 2008) and $X_{i,t}$ is a matrix of socio-economic and political control variables that creates the ceteris paribus condition within the analysis. More specifically, we have $X_{i,t}$ = (INC, GRANT, UN, POP, DEN, YOUNG, OLD, INEQ, NSM, ENPG, ICG)$_{i,t}$, where $INC_{i,t}$ represents taxable income per capita (in 1000 euro), $INEQ_{i,t}$ is an indicator of income dispersion within the municipality and $GRANT_{i,t}$ indicates higher-level government transfers to the municipalities (also per capita and in 1000 euro). UN is the unemployment rate, $POP_{i,t}$ the number of inhabitants, $DEN_{i,t}$ the population density (per km^2), $YOUNG_{i,t}$ and $OLD_{i,t}$ the proportion of young (below 16) and old (over 65) in the municipal population.[5] We also include three political variables: i.e., the number of spending ministers ($NSM_{i,t}$; Wehner, 2010), the effective number of governing parties ($ENPG_{i,t}$) and the ideological position of the ruling party/coalition on a 10-point left-right scale ($ICG_{i,t}$, with larger number reflecting a more right-wing position).[6] Also, given the nature of the data, municipality- and time fixed effects are included (Greene, 2011). Yet, the use of ECM means that the model is first-differenced to remove individual fixed effects.

Finally, as our central explanatory variables, we add indicators for both the gender composition of the population and the representation of women in elected bodies of politicians. The former – i.e., the proportion of women in the population, $FEMPOP_{i,t}$ – is

[5] A number of these variables will themselves be part of the formula used to determine grants. It is treated as an empirical issue as to whether the level of the multicollinearity becomes an issue.

[6] ENPG equals $1/\sum_k p_k^2$ with p_k the vote share of party k. Larger values indicate a higher number of effective parties, and thus a more fragmented party system. We also examined the actual number of parties (without the weighting). The inference of the results was the same, and the effective number of parties dominated. ICG was calculated based on a self-placement survey asking presidents and spokesmen of the parties in the municipalities to locate their party on an ideological scale between 0 (Left) and 10 (Right). Specifically, we weigh each party's left-right score with the share of seats it holds in the local government.

included to control for the possible effect of female preferences within the electorate. The latter – i.e. female representation – is addressed in a number of ways. Firstly, we look at the proportions of women in the local parliament (the council; $FEMCOUN_{i,t}$) and governments (the College; $FEMGOV_{i,t}$). Secondly, we add two variables reflecting the stringency of the regulatory changes: i.e., a dummy variable for those authorities that had no women in their parliament in, at least, the previous two legislative periods ($NOFEMEVER_{i,t}$) and another indicator variable for those authorities that had no female representatives in the 2000-2006 legislative term ($NOFEM2000_{i,t}$).[7]

Given the length of the time series available in the panel, and the fact that budgetary decisions are taken over a one-year time period, we implement a one-year lag-structure in the ECM. The relatively short time-dimension in the data also precludes careful, formal testing of the properties of the time-series variables. As a result, the socio-economic variables are treated as having an order of integration no higher than 1 so that they become I(0) on differencing, and the political and gender variables are deemed to be I(0) in levels.[8] We should also not that the presence of a lagged dependent variable generates biased and inconsistent estimates even if the residuals are white noise.[9] Hence, estimation is undertaken using the Generalised Method of Moments (GMM) developed by Arellano and Bond (1991). This method is ideal for short time-series as the number of valid instruments grows with t.[10] The GMM estimators are calculated using one-step estimates robust to cross-section and time-series

7 Clearly, the former is a subset of the latter as 57 of the 83 authorities with no women in the 2000-2006 parliament has never had a woman in their parliament at all.

8 Preliminary tests of the level of panel integration of the variables substantiates that the economic variables are I(1) and the political variables are I(0). However, given the length of the series and the possibility of discontinuities in some of the variables, the adoption of the ECM-approach should best be seen as a precautionary measure. Still, it is worth highlighting that estimation results in levels (given in Table A1) are consistent with those obtained from the first-difference approach taken in the main text.

9 Crucially, the bias is still present even when the fixed effects model removes the cross-section component of the error term (see Baltagi, 2005).

10 The instruments for the lagged dependent variable are lagged up to four time periods back. Using the full possible range of instruments leaves the average effects the same, but accentuates the extremes.

heteroscedasticity.[11] To test the dynamic properties of the specification estimated, we provide tests for first and second order serial correlation[12], as well as the Sargan specification test of over-identifying restriction.

Clearly, the above discussion only considers whether female representation affects the total level of spending. Yet, even when total spending remains unchanged – i.e., $\beta_5=0$ in equation (1) – it may still be that female representation affects expenditure patterns (e.g., Funk and Gathmann, 2008; Svaleryd, 2009; Chen, 2010). We address this by examining changes in expenditure patterns through the use of a *turbulence index* (proposed by Ashworth and Heyndels, 2001). This builds on Hymer and Pashigian's (1962) mobility index, which in industrial economics is employed to measure market share deviations of individual firms over time. Here, *expenditure turbulence* measures the extent to which municipality i's expenditure structure in year t differs from its spending structure in the previous year. Considering n different expenditures, the expenditure mix of municipality i in year t (E^i_t) is given by:

$$E^i_t = (E^i_{1,t} , E^i_{2,t} , ..., E^i_{n,t}) \qquad 1 \geq E^i_{1,t} \geq 0; \quad \Sigma E^i_{j,t} = 1$$

where $E^i_{j,t}$ is the share of expenditure j in i's total expenditure in year t. The expenditure turbulence index - ΔE^i_t - is then:

$$\Delta E^i_t = \sum_{j=1}^{n} | E^i_{j,t} - E^i_{j,t-1} | \qquad 2 \geq \Delta E^i_t \geq 0$$

The turbulence index takes a value 0 if expenditure structures in t-1 and t are identical; it takes its maximum value of 2 if the expenditure structure has changed completely: any expenditure in year t-1 is non-existent in year t and vice versa. We calculated expenditure turbulence for

[11] The choice of the one-step robust version of GMM was dictated by the fact that standard errors are biased for two-step estimates (see Arellano and Bond, 1991). The results from the two-step coefficient estimates are broadly comparable with those reported and, as is normal, reinforce the significance of the significant variables.

[12] If the error terms of the basic model are uncorrelated, then the first differenced error terms are an MA(1) process implying negative first order serial correlation and an absence of second order correlation.

all municipalities using 7 main spending categories: i.e., Finance & Administration; Police; Infrastructure; Education; Arts, Recreation and Libraries; Environment; Social Welfare.

Using this turbulence index, the following basic regression model is considered:

$$\Delta E^{i}_{t} = \beta_0 + \beta_1 \Delta POL_{i,t} + \beta_2 \Delta GEN_{i,t} + e_{i,t} \qquad (2)$$

Where we assume that expenditure turbulence is affected only by political variables, particularly changes in women's representational position.[13]

3.2. Empirical results

Our baseline results from estimating equation (1) are presented in Column (2) of Table 2.[14] As can be seen, the equation is well-specified and satisfies all the relevant tests of over-identification and serial correlation.

TABLE 2 HERE

Briefly looking at the control variables first, we find that, broadly in line with earlier work, income, income dispersion and unemployment increase expenditures. Population size has a negative effect, which may reflect scale economies in the provision of public services. The share of elderly has a stimulating effect on expenditures (see also Bastiaens *et al.*, 1997), whilst the share of young has no similar effect (indicating that grants are covering these areas

13 A larger regression $\Delta E^{i}_{t} = \beta_0 + \beta_5 \Delta X_{i,t} + \beta_5 \Delta POL_{i,t} + \beta_7 \Delta GEN_{i,t} + e_{i,t}$ was also examined on the premise that some expenditure turbulence comes from economic and socio-demographic factors, over which the politicians may have less direct control. The results of this exercise indicate that some expenditure turbulence is indeed related to such non-political factors. Crucially, however, there is no change in the inferences about the significance or size of the effects of the political variables and so the simpler model is reported.

14 Some initial work was undertaken with a simpler formulation of the model (see table A1 in Appendix). Specifically, we estimated the basic long-run component (i.e., what is effectively contained in the unconstrained ECM), for the final two years of the sample (2007 and 2008) and the initial years (1999 and 2000). This gives a snapshot of the effect of the imposition of the 2007 minimal representation regulation. It can be seen that there is minor evidence of an effect of women in government in the simple OLS estimation but this disappears if one considers an estimation by instrumental variables to accommodate endogeneity. Moreover, even if we accept these snapshot OLS results, the effect of female politicians is tiny relative to the effect of the female population (i.e., one twentieth).

satisfactorily).[15] With respect to the political variables, we find that right-wing governments implement lower expenditure than left-wing governments, and that a higher number of parties and a higher number of spending ministers in the governing coalition increases expenditures.[16] While the former reflect the standard assumption that left-wing parties are more inclined to support state intervention, the latter can be explained by the fact that there are more groups to satisfy (Ashworth *et* al., 2005; Wehner, 2010). Note that the effects discussed so far are long-run effects (from the lagged level equation). However, it can easily be see that the speed of adjustment over the period is slow and that there are only relatively weak immediate effects (from the difference equation) for income, unemployment, political fragmentation and the proportion of old people.

Turning now to the central gender effects, we find significant evidence that a larger female proportion in the population leads to greater spending, entirely in line with previous work (e.g., Lott and Kenny, 1999; Aidt and Dallal, 2008; Miller, 2008). One reason for this may be their greater proportion in the elderly population, but it might also reflect the difference in policy preferences of women compared to men (see section 1; for an overview, see Croson and Gneezy, 2009). The picture is quite different, however, when we look at the representation of women in the local government. In effect, the evidence here is uniform in that there is no indication (from either the proportion of females in parliament and government) suggesting that a higher representation of female politicians affects spending

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[15] As regards density there is little overall effect and omission of this variable has no effect on the other results presented. Its continued inclusion is to enable direct comparisons with previous work. The positive coefficient may indicate that small municipalities spend more than large one, though the negative in the short run may indicate that they take longer to make this adjustment. Of course, the large standard errors encourage caution in such interpretation.

[16] As previous work on Flanders (e.g., Ashworth *et al.*, 2005, 2006; Geys, 2007; Goeminne *et al.*, 2008) has found a non-linear relation, we experimented with the square of the effective number of parties. This substantiates earlier work by showing a negative sign on the number of parties and a positive sign on the square with the turning point at just above 2 parties (though both variables are significant only at just below 90% confidence). As this does not affect our other results, we chose to retain the linear specification (all other results available on request).

levels in the long run. Furthermore, the indicator variables for municipalities without women in current and/or historical governments show insignificant and positive signs – indicating that authorities where women were new in the 2006-2012 legislature were spending more, *ceteris paribus*, during the period without women on the council.[17] This effect also carries over for female mayors. It is crucial to note here that there are time dummies present in the estimation. This implies that any drift associated with a change in representation mix is picked up by these dummies, and hence the test here is for the clear associated shock of 'new' female representation.[18] Finally, we should mention the short-run effect of female representation, which is borderline significant and suggests that the immediate effect of women entering government is, if anything, to reduce spending. One possible reason may be that they use such strategy to establish their "credentials". Also, those councils that have women for the first time are reducing expenditure (remember the reverse coding of this variable, see note 7). Yet, the interpretation of this result should be treated with caution since it is unclear whether women are the cause of this, or whether this is a natural adjustment of higher spending municipalities.

Overall, our results thus far indicate that legislation on the gender mix in politics will have, at most, a minimal effect of spending in proportional representation systems. However, the results presented above are for all (save 9) municipalities in Flanders. This begs the question whether greater underlying differences exist for those municipalities that did not have female representation before the mandatory minimal representation requirement – and thus, possibly, having female representation forced upon them – and those that had already "embraced the

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[17] The results presented are for the period where there were no women; the effects of women being introduced are therefore the negative of this result.

[18] These time dummies show no indication of a positive time trend indicating higher spending when the gender-legislation was more stringent. Crucially, there is also no significant positive time shock in the years following elections, notably 2007, that might be an effect from the change in female representation (details available upon request).

change" willingly. To that end, a number of statistical tests were made to examine whether both groups of municipalities could be treated similarly, or not. First, separate estimations were made of the 83 municipalities with no women in parliament in 2000-2006 and the other municipalities to test if there a significant difference between the two sets of municipalities (see columns (3) and (4) in table 2). The test procedure follows Andrews (1993) and consists of a test that confines the instability to the parameters of the model alone (a Wald test) and a test that allows instability also on other aspects of the model (based on the over-identifying restrictions; Hall and Sen, 1999).[19]

The results are provided near the bottom of table 2. Clearly, neither of the tests for stability is significant. That is, there is no structural difference between "men only" municipalities and municipalities that contained women prior to it being a requirement. Even so, it could still be that whilst there is no difference between municipalities by group, there was a systematic shift after 2006 (i.e., in 2007 and 2008 when allowing for policy to be enacted) that is not identified in the simple dummy approach, but that is particularly relevant for those municipalities without women. To that end, both tests were re-run examining a break point at 2006. These tests were undertaken in the full model and in the model using only the sub-set of 83 "men only" municipalities. Again, there is no evidence of a break. The conclusion is that the introduction of the minimal representation requirement for female politicians has had no (immediate) effect on spending in Flemish municipalities. As a final evaluation, we re-ran all tests examining a break point at 2000 – that is, after the first election in our sample. The principle here is that the quota-legislation has been brought into force in stages and therefore may have induced some pre-emptive action by authorities. Hence, the "women's voice" in

[19] While these tests only apply to a known break point, which is clearly the case here, Hall and Sen (1999) show that the conclusion holds even with the wrong break point chosen. Note also that the first Wald test can be made using information from the original full model and these are the results given. Still, the tests were also undertaken using the sub-divisions provided columns (3) and (4) and there is no change in the inference (see Andrews, 1993, and Hall, 2005, for the justification of this).

politics could have been introduced in anticipation of formal legislation.[20] Again, there is no evidence of a significant effect of women in politics.

As mentioned above, the fact that increased female representation does not appear to affect the overall level of expenditures does not necessarily imply that it has no effect on expenditure patterns. Geys and Revelli (2011) indeed note that female representation in Flanders is associated with a different tax mix, and a similar effect may be hidden under the results in Table 2. To assess this in more detail, we report results from estimation equation (2) – which evaluate revenue structure turbulence – in Table 3. A number of different formulations were examined. First, using the entire time period available (i.e., 1998-2008), we estimate the model without year fixed effects (Column (1)), using time effects capturing both post-election periods (2001/02 and 2007/08) independently (Column (2)) and jointly (Column (3)). Second, we ran separate regressions for both post-election periods separately, in order to focus on the years in which expenditure readjustments from the new government are most likely.

What is immediately clear from the first three columns in Table 3 is that there are time-specific effects (thus rejecting the model without time effects in Column (1)) and that the treatment of these is critical (e.g., treating both elections as being the same is rejected).[21] Overall, there is a clear election effect creating expenditure turbulence for one or two years after the election. There are additional effects from changing the number of spending

[20] This is contrary to the idea that the introduction of quotas makes women politicians' credibility *decline* (Reynaert *et al*., 2005). To examine the life of the one parliament where our premise would make most sense, we also repeated the test for data to 2006 only. This did not affect our inference. Note also that, in the event that there was drift in behaviour to accommodate the forthcoming legislation, a case could be made for testing for an unknown break (Hall, 2005); the premise here is that there are shifts between parliaments. One final test was examined to see if the "men only" parliaments had a different long-run equilibrium that they were "aiming" for by. This test followed Enders and Siklos (2001) and allowed different long-run equilibria for the two groups. Tests of the coefficients indicated that there was no significant difference.

[21] It is obviously still possible that the election effects are the same, except for additional exogenous effects in these years (e.g., the move to the euro) that affect expenditure turbulence in a given year.

ministers and from a change in the fragmentation of the governments, but the gender of politicians appears irrelevant.

However, it may be argued that finding an election (year) effect indicates a composite effect of elections on tax expenditures, and that this effect should be decomposed into its constituent parts. In order to examine this possibility, two further regressions for the periods directly after the elections are presented in the final columns of Table 3.[22] Here there is some evidence that increased female representation in the local parliament (and government – though this comes only close to 10% significance) affected expenditure turbulence following the election of 2000.[23] The effect for women in parliament thereby reflects a sizeable proportion of the expenditure turbulence observed. However, a similar effect does not re-appear following the 2006 election. Overall, the conclusion must therefore be that the effect of quotas will at best have a very limited effect on expenditure and expenditure patterns in Flanders. Moreover, what (minimal) effect there has been in the past, appears to have worked its way through the system.

5. Conclusion

Geys and Revelli (2011) show that female representation in Flanders appears to affect the tax structure of municipal governments. This article demonstrates that it does not appear to affect the level, nor the distribution, of local public expenditures. Clearly, this may be due to a lack of influence. Still, this should show up when women have become mayors or held spending, as opposed to representative, posts, which is not the case. It may also reflect the fact that female politicians may simply not have the same policy preferences as the general female

[22] As in the previous, the treatment of the two elections as the same is firmly rejected with an F statistic of equality of parameters of 19.482.

[23] Results from combining the two elections indicate similar results. Moreover, some evidence then also suggests that having no women in the last parliament has a small effect on turbulence.

population (see Ågren *et al*., 2006). Independent of the underlying mechanism, ‘concerns’ that women’s desire for larger government or their different spending priorities will lead to (further) increases in the size of government as their representation in politics rises appear, based on our evidence of Flanders, ill-founded. Politicians are politicians and gender does not appear to be a source of difference between them.

TABLE 2: GMM ESTIMATION OF THE ADJUSTMENT OF CURRENT EXPENDITURE

Dependent Variable	ΔEXP	ΔEXP	ΔEXP
		Only Men in Parliament, 2000-2006 (83)	WOMEN in Parliament 2000-2006 (216)
Lagged Levels			
Expenditure	**-0.143** **(0.010)**	**-0.200** **(0.056)**	**-0.132** **(0.019)**
Income	**1.498** **(0.801)**	***1.272*** ***(0.731)***	**1.105** **(0.775)**
Population	**-0.049** **(0.010)**	**-0.037** **(0.009)**	**-0.004** **(0.001)**
Density	0.395 (0.562)	1.556 (1.702)	0.001 (0.005)
Old	**4.653** **(1.213)**	***4.468*** ***(2.396)***	**4.122** **(1.481)**
Young	-0.760 (1.467)	1.529 (1.627)	-0.607 (1.350)
Unemployed	***10.089*** ***(3.046)***	**6.811** **(2.753)**	**8.706** **(3.274)**
Income Dispersion	***0.061*** ***(0.032)***	***0.075*** ***(0.048)***	0.133 (0.333)
Grants	0.001 (0.006)	0.004 (0.003)	0.003 (0.007)
Female Proportion	**6.254** **(2.701)**	**18.139** **(9.320)**	**5.445** **(2.293)**
Ideology	**-4.875** **(2.994)**	**-4.440** **(2.057)**	***-3.982*** ***(2.104)***
Females in Government (Proportion)	-0.075 (12.701)	-6.057 (29.541)	13.286 (15.750)
Females in Parliament (Proportion)	-0.248 (0.240)	-0.284 (0.362)	-0.274 (0.278)
Female Mayor	-7.560 (7.069)	-15.116 (14.771)	-4.795 (7.232)
No Women in Present parliament	17.143 (18.900)	14.653 (8.130)	
No Women Ever in Parliament	22.413 (22.520)	2.745 (9.638)	
Number of Spending Minister	**1.891** **(0.760)**	**2.663** **(1.357)**	***3.026*** ***(1.704)***
Number of Parties	**3.343** **(1.747)**	***10.647*** ***(5.400)***	**4.742** **(2.408)**
Differences			
Income	**0.004** **(0.003)**	***0.005*** ***(0.003)***	0.006 (0.004)
Population	0.005 (0.011)	-0.072 (0.076)	0.007 (0.005)
Density	-0.061 (0.523)	0.112 (0.735)	-0.193 (0.462)
Old	**16.971** **(8.994)**	**19.780** **(3.571)**	***18.947*** ***(11.290)***
Young	1.147 (1.803)	-1.863 (1.355)	1.557 (1.744)
Unemployed	**30.437** **(5.881)**	10.416 (11.880)	**32.265** **(7.344)**
Income Dispersion	0.148 (0.458)	-0.775 (0.995)	0.664 (0.961)
Grants	0.002 0.004)	0.002 (0.003)	0.002 (0.004)
Female Proportion	14.463 (13.383)	30.047 (25.950)	19.221 (18.800)
Ideology	-18.868 (6.421)	-26.976 (13.832)	-13.2518 (8.049)
Number of Parties in Government	**8.243** **(3.692)**	**5.358** **(2.079)**	***8.200*** ***(4.377)***

TABLE 2: GMM ESTIMATION OF THE ADJUSTMENT OF CURRENT EXPENDITURE (Continued)

Females in Government (Proportion)	***-45.294*** ***(25.034)***	***-136.417*** ***(88.501)***	-0.387 (0.301)
Females in Parliament (Proportion)	-0.193 (0.409)	0.215 (0.766)	-0.081 (0.623)
Female Mayor	-6.639 (12.882)	-8.742 (17.320)	-2.349 (14.320)
No Women in Present Parliament	15.404 (19.941)		
No Women Ever in Parliament until 2006	**42.704** **(22.802)**		
Spending Minister	0.144 (10.030)	0.144 (10.030)	-8.167 12.110)
Time Effects (Wald significance)	Yes 75.47	Yes 24.86	Yes 79.65
R2	0.221	0.159	0.190
SARGAN	13.772	6.861	10.006
SC 1	**-5.965**	**-3.245**	**-5.638**
SC 2	1.209	1.029	0.420
WALD OVERALL	**320.217**	**228.523**	**272.446**
STABILITY (2006)			
Wald	1.178	1.221	0.883
Hall and Sen	3.771	4.331	2.887
STABILITY (2000)			
Wald	2.443	0.448	3.115
Hall and Sen	4.196	3.992	5.663
WALD (MEN ONLY)			
Wald	0.224		
Hall and Sen	4.902		

Notes to Table 2

All estimations use the DPD package of Arellano and Bond. The results are one-step estimates, robust to heteroscedasticity with estimated standard errors in parentheses. SC(1) and SC(2) are the required tests of serial correlation with SARGAN being the test of over-identifying restrictions, see Arellano and Bond (1991) and Baltagi (1995). WALD gives the test of omission of the variables as indicated. In all cases, degrees of freedom are given in parentheses. The intercept is omitted for space but is insignificant in all cases. Numbers in bold are significant at 5% significance; number in italics at 10% significance. The WALD test of joint significance of the variables is highly significant in all cases. It is possible to omit Density with a Wald statistic of 2.022 with no discernible effect on the other results. The effect of a female mayor in Men Only municipalities is due to 3 municipalities having a female mayor from 2007.

TABLE 3: ESTIMATION OF TAX EXPENDITURE TURBULENCE

Dependent Variable	Turbulence	Turbulence	Turbulence	Turbulence	Turbulence
Estimation	OLS	OLS	OLS	OLS	OLS
Time Periods	**1999- 2008**	**1999- 2008**	**1999- 2008**	**2001, 2002**	**2007, 2008**
Ideology (Absolute)	**0.081** **(0.008)**	-0.008 (0.008)	-0.014 (0.009)	**0.117** **(0.014)**	***0.024*** ***(0.013)***
Females in Government (Proportion)	**0.054** **(0.025)**	0.021 (0.027)	0.032 (0.023)	0.054 (0.034)	0.044 (0.036)
Females in Parliament (Proportion)	**0.284** **(0.044)**	0.012 (0.043)	0.013 (0.044)	**0.216** **(0.095)**	-0.052 (0.064)
Female Mayor	0.014 (0.018)	- 0.016 (0.016)	- 0.005 (0.017)	0.005 (0.031)	-0.011 (0.027)
No Women in Present/Last parliament	***0.063*** ***(0.039)***	0.015 (0.020)	0.026 (0.022)		-0.030 (0.026)
No Women Ever in Parliament	0.042 (0.026)	0.001 (0.023)	0.000 (0.024)		0.002 (0.028)
Number of Spending Minister	**0.021** **(0.011)**	**0.020** **(0.010)**	**0.023** **(0.011)**	**0.051** **(0.026)**	***0.026*** ***(0.014)***
Number of Parties	**-0.103** **(0.032)**	**-0.108** **(0.031)**	***-0.056*** ***(0.033)***	**-0.102** **(0.056)**	***-0.068*** ***(0.034)***
Number of Parties ^{2}	**0.022** **(0.008)**	**0.027** **(0.007)**	**0.018** **(0.008)**	**0.031** **(0.013)**	***0.018*** ***(0.010)***
2001		**0.203** **(0.008)**			
2002		0.005 (0.007)			
2007		**0.028** **(0.009)**			
2008		**0.061** **(0.008)**			
One Year After Election			**0.134** **(0.006)**		
Two Years After Election			0.028 (0.004)		
INTERCEPT	**0.126** **(0.002)**	**0.103** **(0.005)**	**0.103** **(0.002)**	**0.156** **(0.006)**	**.155** **(0.006)**
R^2	0.085	0.306	0.187	**0.237**	**0.062**
Time Effects		**254.375**	**186.244**		
Restriction of Years		**104.803**			
RESET	**5.542**	0.374	**0.973**	2.665	2.117
NORMALITY	**35.772**	*8.337*	**13.221**	**27.942**	**36.302**
HETEROSCEDASTICITY	**2.441**	1.429	**2.178**	1.652	1.226

Notes: Estimated standard errors are in parentheses. Using a panel estimation of group means give the same overall implications as those presented here. All independent variables are first differences. All other year effects are insignificant, both individually and jointly; the test of restriction of years is a test of the parameters on the election years being the same in different elections. Where heteroscedasticity is identified, heteroscedastic robust standard errors are provided; in no case, does this change the inference from using OLS errors.

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TABLE A1: ESTIMATION OF CURRENT EXPENDITURE

Dependent Variable	EXP	EXP	EXP	EXP
Estimation	OLS	IVE	OLS	OLS
Time Periods	**2007, 2008**	**2007, 2008**	**1999, 2000**	**2001, 2002**
Income	3.390 1.028	3.372 (1.157)	9.749 10.854	22.177 (11.665)
Population	-0.003 (0.001)	-0.003 (0.001)	-.0002 (0.001)	-.0003 (0.001)
Density	0.024 (0.023)	0.046 (0.039)	.0137 (0.023)	-.0002 (0.023)
Old	31.538 (5.810)	23.629 (6.107)	4.932 (4.566)	15.801 (4.864)
Young	-6.911 (6.884)	3.441 (5.372)	-2.613 (4.114)	3.447 (6.102)
Unemployed	68.981 (19.068)	61.004 (22.752)	15.415 (12.704)	48.249 (16.224)
Income Dispersion	0.575 (1.218)	0.449 (1.630)	2.192 (1.416)	2.201 (1.534)
Grants	0.025 (0.011)	0.027 (0.013)	1.815 (0.340)	1.713 (0.325)
Female Proportion	34.980 (16.494)	38.980 (17.119)	83.724 (15.671)	62.580 (15.555)
Ideology	-18.571 (11.784)	-17.118 (12.493)	-43.757 (13.050	-27.967 (11.468)
Females in Government (Proportion)	-0.145 (0.722)	-0.369 (0.592)	-0.158 (0.060)	.1211 (0.770)
Females in Parliament (Proportion)	-2.042 (1.041)	-2.173 (1.239)	2.375 (1.038)	1.320 (0.976)
Female Mayor	-22.781 (28.907)	-14.781 (31.907)	-57.726 (31.749)	-11.990 (29.656)
No Women in Present/Last parliament	6.322 (28.838)	2.892 (19.308)	16.990 (25.979)	49.067 (33.116)
No Women Ever in Parliament	12.742 (32.056)	9.862 (28.002)	-38.678 (32.656)	-4.877 (29.521)
Number of Spending Minister	**20.002** **(8.421)**	**19.086** **(9.113)**	**28.886** **(9.417)**	**26.214** **(9.195)**
Number of Parties	-146.594 (69.625)	-139.661 (71.423)	-106.876 (58.727)	-135.511 (57.850)
Number of Parties 2	26.772 (16.291)	32.761 (17.981)	23.946 (15.145)	32.709 (14.606)
INTERCEPT	-2766.243 723.380	-2475.328 (811.608)	-3393.592 (702.341)	-3099.399 (704.229)
Time Effects	Yes	Yes	Yes	Yes
R2	0.485	0.527	0.518	0.526
SARGAN – instruments		7.119		
RESET	**12.442**	**7.331**	**11.630**	**9.114**
NORMALITY	31.427	28.332	29.042	18.331
HETEROSCEDASTICITY	1.622	**1.407**	**1.397**	**1.711**

Notes: Estimated standard errors are in parentheses. In the Instrumental Variable Estimation, the previous two years are used as instruments; results for the other years show similar changes. The results are mirrored if artificial instruments, following Dagenais and Dagenais (1997) are used. Using a panel estimation of group means give the same overall implications as those presented here.