

# **Double Standards of Recruitment for Men and Women? New Evidence from Moonlighting of German Parliamentarians**

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## **Abstract**

Public and scholarly debate on politicians' extra-parliamentary activities (i.e. moonlighting) often concentrates on whether such 'moonlighting' engenders neglect of parliamentary duties and/or signals financial greed. Building on extensive research on gender effects in recruitment processes, this paper takes a different perspective and analyses how politicians' gender affects their probability of having outside activities. Using a novel dataset of moonlighting by members of the German national parliament (including 614 MPs over the period 10/2005-09/2007), our analysis confirms theoretical predictions that political experience and university education are more important determinants of moonlighting for women than men, while age and children work more forcefully against women than men. These results hold particularly for outside jobs in the private sector. Overall, even when analysing a recruitment pool of higher-than-average ability working in a strongly politicised environment, gender matters strongly in recruitment processes.

**Keywords:** Moonlighting, Recruitment, Gender, Glass Ceiling, Stereotypes.

## **1. Introduction**

Elected representatives in most countries are legally permitted to carry out jobs in addition to their political mandate. Such activities are generally referred to as ‘moonlighting’ and attract significant public and political debate (e.g., Noel-Baker 1961; Hollingsworth 1991; van Aaken and Voigt 2009) as well as scientific inquiry (e.g., Parker 1992; Couch et al. 1992; Rosenson 2007; Gagliarducci et al. 2008; Eggers and Hainmueller 2009; Becker et al. 2009; Mause 2009). These discussions, however, mostly focus on whether or not moonlighting indicates financial greed, impinges on politicians’ time for political activities (Gagliarducci et al. 2008; Becker et al. 2009), leads to conflicts of interest (Stigler 1967; Couch et al. 1992; Rosenson 2007) or, in a more positive view, improves the quality of parliamentarians by bringing additional experience and qualifications (Gagliarducci et al. 2008).

Studies addressing variation in moonlighting across politicians are much less common (for exceptions, see Parker 1992; Becker et al. 2009; Mause 2009). Yet, clearly, not every politician moonlights. Moreover, some moonlighting politicians hold one outside job, while others have more than 20 such jobs. Likewise, some members of parliament (MPs) have outside jobs in the private sector of the economy, others only (or also) in the public sector (e.g., in city councils, regulatory agencies, associations or foundations). This paper applies insights from a large social science literature addressing the development of women into positions of leadership to address whether, and how, politicians’ gender influences the extent of their moonlighting behaviour. Such approach adds to the burgeoning moonlighting literature by investigating the—thus far largely neglected—determinants of moonlighting activity. It also complements extensive research documenting the effects of gender on purely political careers (both relative to individuals’ electoral choice—e.g., Rosenwasser and Seale 1988; Sanbonmatsu 2002—and intra-party candidate selection—e.g., Rahat 2007) and on private-sector hiring and promotion decisions (e.g., Morgan 1998; Cotter et al. 2001; Albrecht et al. 2003; for a recent review, see Jackson and O’Callaghan 2009).

The empirical results, using data covering 614 German MPs in the period 10/2005-09/2007 (32.1% of which are female), show considerable support for the influence of gender on moonlighting activities. Women are not only less likely than men to have many sideline jobs (especially in the private sector), but university education and political experience are significantly more important for women than men in obtaining such ancillary positions (once again especially in the private sector). The latter effect is not age-related as age tends to work against women, while it works in favour of men. Finally, while children are associated with a higher number of sideline jobs (possibly due to higher spending and, thus, income requirements associated with family expansion), this ‘child effect’ is particularly strong for male MPs while being statistically insignificant for female ones. Hence, overall, even when analysing a recruitment pool of higher-than-average ability working in a strongly politicised environment, gender retains a manifest influence on recruitment decisions.

It is important to stress at this point that, although our analysis uncovers clear and strong gender-effects in politicians’ moonlighting activities, this should not be interpreted as manifest evidence of gender discrimination or employers’ pro-male bias. While this is one possible interpretation, another is that women are simply less likely than men to seek and/or accept offers of (especially private-sector) sideline activities. The ‘double standards’ in the title thus need not imply employers’ discrimination, but could also refer to different standards male and female MPs set themselves as regards outside employment. We return to this much more extensively below.

The paper proceeds as follows. In the next section, we present our theoretical argumentation and hypotheses. Then, Section 3 introduces the dataset employed, and Section 4 describes our empirical strategy and results. Finally, Section 5 reviews our main empirical findings, and suggests some avenues for further research.

## **2. Theoretical Background and Hypotheses**

From a theoretical perspective, politicians' observed moonlighting behaviour can be interpreted as the equilibrium outcome of a recruitment market. On one side of this market, there are firms or organisations with openings in, for example, their supervisory board seeking candidates that best match their desired profile and performance expectations. Assuming a candidate with a political background is considered, they will map their preferences onto the pool of available politicians and make an offer to one or more politicians closest to their desired profile. Then, on the other side of the market, there are politicians being offered sideline positions and deciding whether or not to accept such offers. However, as politicians are not necessarily passive agents in this market, they may also actively apply or put themselves forward for consideration by sending appropriate signals. In either case, their decisions will be guided by constraints from their personal situation (family, income, and so on), their professional circumstances (e.g., potential for conflicts of interest), and the characteristics of the available or offered jobs (e.g., remuneration, intellectual challenge, time-investment, future career possibilities).

Bringing both sides together, it is clear that, when analysing moonlighting behaviour across politicians, it is important to take into account what makes politicians *a*) of interest to those offering outside jobs and *b*) likely to seek out and/or accept/reject such positions.<sup>1</sup> Building on a large social science literature addressing the development of women into positions of leadership (e.g., Morgan 1998; Cotter et al. 2001; Albrecht et al. 2003; for a recent review, see Jackson and O'Callaghan 2009), this paper specifically analyses the role of gender on the outcome of such recruitment process.

A central reason why gender affects recruitment decisions is that it functions as one of the primary means of categorizing individuals in Western societies (Brewer and Lui 1989;

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<sup>1</sup> Evidently, these elements need not overlap. For example, politicians in important political positions may be more interesting to outside employers (Parker 1992; Becker et al. 2009), but may also be more reluctant to seek out and/or accept such offers due to real or perceived conflicts of interest.

Ridgeway 1997, 2009). As soon as we establish some form of interpersonal interaction—e.g., during recruitment procedures—actors “automatically and unconsciously sex-categorize any other to whom we must relate” (Ridgeway 1997: 220; Brewer and Lui 1989; Stangor et al. 1992). Such sex categorization has important consequences for how we behave towards each other, because it easily triggers gender stereotypes (Deaux and Major 1987; Bargh 1989), which “denote not only differences in how women and men actually are [descriptive], but also norms about behaviours that are suitable for each [prescriptive]” (Heilman 2001: 659; Terborg 1977; Burgess and Borgida 1999).

Importantly, inherent in such (gender) stereotypes are status beliefs, that is, “widely held cultural beliefs that link greater social significance and general competence (...) to one category of social distinction [e.g., race, gender, age, and so on] compared to another” (Ridgeway 2001: 638). According to expectation states theory (Berger et al. 1977), such status beliefs help shape expectations concerning the performance and ability of individuals—and thus play a critical role in recruitment processes. Crucially, with respect to gender, status beliefs are well documented to associate “greater status worthiness and competence with men than women” (Ridgeway 2001: 637; Broverman et al. 1972; Deaux and Kite 1987).<sup>2</sup> The implied gender hierarchy prescribed by gender status beliefs therefore gives rise to expectations that women *ceteris paribus* are likely to underperform men (for reviews, see Swim et al. 1989; Bowen et al. 2000) and have been ‘lucky’ rather than ‘able’ to reach performance of a given quality (e.g., Deaux and Emswiller 1974; for a review, see Swim and Sanna 1996). Consequently, a ‘lack of fit’ is perceived between women’s characteristics and the requirements for success in particular jobs (Heilman 1983; 1995) and “a preference for male workers” is generated (Ridgeway 1997: 228).

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<sup>2</sup> Although perceptions of women have improved in recent decades, this basic hierarchical ordering has proven very resistant to change and remains to date (Ridgeway 2001; Leuptow et al. 1995).

Moreover, as these status beliefs and resulting expectations have important self-fulfilling effects (Harris and Rosenthal 1985; Miller and Turnbull 1986; Heilman 2001), they not only affect recruiters, but also female applicants. In fact, it undermines the latter's "assertiveness and confidence, (...) their actual performance and their influence in the situation" (Ridgeway 1997: 222). As such, women may not only be less likely to be offered certain positions, but may at the same time perceive themselves as less appropriate and thus be less likely to seek them out or accept them when offered. Hence, our first hypothesis is:

**H1a:** *Women are, compared to men, less likely to have many sideline jobs.*

The depicted effects have, unsurprisingly, been shown to hold more strongly for male gender-typed jobs (such as private-sector management positions); see, e.g., Schein (1973, 2001), Heilman et al. (1989), Dodge et al. (1995) and Watts (2009) on the perceived incompatibility between women's abilities and leadership requirements. Indeed, for such positions, the "very essence of gender stereotypes defines men as more instrumentally competent and agentic compared to women" (Biernat and Fuegen 2001: 707; Broverman et al. 1972; Smith and Midlarsky 1985), leading to a larger perceived lack of fit between female characteristics and characteristics required for success in such jobs.<sup>3</sup> This further specifies the hypothesis formulated above in the following way:

**H1b:** *The relation described under H1a is most pronounced for private-sector functions.*<sup>4</sup>

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<sup>3</sup> A similar 'self-selection' effect exists in female-dominated professions. Snyder and Green (2008: 271), for example, find evidence of a "pervasive pattern of horizontal sex segregation" in the nursing profession, whereby "male nurses tend to gravitate toward areas of nursing they perceive to be more masculine".

<sup>4</sup> As 'old-boy networks' often have important effects on hiring processes and outcomes (e.g., Saloner 1985; Burke 1997; McDonald et al. 2009), this hypothesised effect may, in the German setting analysed below, be strengthened by the fact that the German private sector is still very much male-dominated.

By predefining the way we regard candidates for a given position, gender stereotypes lead women to start off with a disadvantage. To overcome this, it is “in women’s interest to introduce additional job-relevant information that undermines [gender status beliefs’] effects on perceptions of the competence” of women (Ridgeway 1997: 227). Often, this implies that they have to show “greater evidence of competence” in order to be recruited or promoted (Lyness and Heilman 2006: 779). This is closely related to the theory about double standards for competence derived from expectation states theory (Foschi and Foddy 1988; Foschi 1996; 2000), which argues that “performance by members of lower status groups (e.g., women) [is] assessed by stricter standards than similar performance by members of higher status groups (e.g., men)” (Lyness and Heilman 2006: 779). Experimental analyses of this theory have indeed shown that women are held—by both male *and* female judges—to higher confirmatory standards (i.e. standards indicating a candidate has the ability to perform a given task). They thus need to have a higher skill level to convince judges they actually have the ability to perform the job (Biernat and Kobrynowicz 1997; Biernat and Fuegen 2001; Ridgeway 1997; 2001). Such gendered difference in standards for diagnosing competence implies that women have to work harder to achieve promotions, certain jobs, or levels of authority (for empirical evidence, see, e.g., Kramer and Lambert 2001; Lyness and Heilman 2006; Gorman and Kmec 2007).

This line of argument suggests that evidence of individual abilities is likely to be more important for women than men to prove their value—both to outside employers *and* to women themselves (in order to overcome the self-fulfilling effects of status beliefs, see above). In the setting under investigation, such evidence can derive from educational attainment and relevant professional experience (i.e. more terms in office). Although both signal competence, the underlying mechanisms differ. Indeed, educational attainment signifies a given stock of human capital, while more terms in office generates both human (i.e. knowledge of political processes) and social capital (i.e. social networks and contacts).

The latter attributes are likely to be especially valued by firms and organisations hiring politicians. This leads to our second set of hypotheses.

**H2a:** *University education is more important for women than men in obtaining sideline activities.*

**H2b:** *Political experience is more important for women than men in obtaining sideline activities.*

Gender stereotypes not only contain beliefs about the overall status of men versus women (see above), but also grant each gender particular skills (Ridgeway 2001). For example, mechanical ability and leadership skills are often associated with men, while domestic skills tend to be reserved for women (Broverman et al. 1972; Schein 1973; Heilman et al. 1989; Williams and Best 1990). This not only affects women's perceived aptitude at 'male'-designated tasks, but also creates expectations that women are more likely to take care of family and children. Moreover, family is perceived to be a "greater priority for women than career advancement" (Kramer and Lambert 2001: 113; Nieva and Gutek 1981). This family-orientation is "often viewed by promotional decision makers as lacking loyalty and commitment" (Kramer and Lambert 2001: 113; Collinson 1998), and thereby reduces the promotion potential of women (for supportive evidence, see Halford et al. 1997; Kramer and Lambert 2001; Granqvist and Persson 2005). Interestingly, the reverse pattern might hold for men. Indeed, men with children are often perceived as "more serious" and in possession of "mature leadership qualities" compared to their single or childless counterparts, which, in turn, makes them more likely candidates for promotion (Coltrane 2004: 215).

Both elements suggest that the number of children may have different effects for the ability of male and female politicians to attract outside jobs. Clearly, on the other side of the recruitment market, gender might also affect whether outside jobs are actively pursued and how obtained offers are evaluated once the family expands. Such "gendered expectations

about home-making and breadwinning” (Coltrane 2004: 214) present us with our third hypothesis.

**H3:** *Children have a stronger constraining effect of the number of sideline activities of female than male politicians.*<sup>5</sup>

Finally, while numerous studies illustrate employers’ negative stereotypical attitudes towards older employees (e.g., Taylor and Walker 1994, 1998), little research to date has regarded how gender interacts with such ageism on occupational outcomes (Duncan and Loretto 2004). Nonetheless, it has since long been argued that physical signs of aging might be a stronger disadvantage for women compared to men (e.g., Ward 1977; Sontag 1979) because women are valued, much more than men, “in accordance to the conditions ascribed to their youth” (Walker et al. 2007: 39; Bazzini et al. 1997). Such “double standard of aging” (Sontag 1979: 462) or “gendered ageism” (Itzin and Phillipson 1995) has found significant empirical support. For example, women are perceived as losing physical attractiveness to a faster and greater extent than men as they age (e.g., Berman et al. 1981; Deutsch et al. 1986), which is important since individuals perceived physically less attractive are disadvantaged in the job market (e.g., Dipoye et al. 1975; Morrow et al. 1990; Hamermesh and Biddle 1994). Moreover, in movies, often viewed as “reflections of a culture’s attitudes, beliefs and standards” (Bazzini et al. 1997: 532; Lincoln and Allen 2004), there is a “much stronger tendency to negatively portray aging women as compared to aging men”, in terms of physical attractiveness as well as intelligence, friendliness and goodness (Bazzini et al. 1997: 540). This closely relates to experimental evidence showing that older men score higher than older women when judged by experimental subjects on “intellectual competence” and “autonomy” (Canetto et al. 1995), placing older women at a clear comparative disadvantage in the job

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<sup>5</sup> While a general hypothesis, Germany’s conservative welfare state (see Esping-Anderson 1990) may strengthen this effect by institutionally promoting male breadwinners and female caregivers.

market relative to older men. Importantly, managers/employers are not immune to this stereotype, since research shows that they perceive women to ‘reach their peak’ at a younger age than men (Itzin and Phillipson 1995).

These arguments imply possibly significant negative effects of being older *and* being female on occupational outcomes. The only empirical study thus far addressing this age-gender interaction finds supportive evidence using a sample of 318 movie stars over the period 1926-1999. Specifically, older female actors are “disadvantaged both in terms of number of film roles and in terms of average star presence [i.e. importance in movie judged via rank in the film’s credits], compared to older male actors” (Lincoln and Allen 2004: 623). While age and physical appearance may be particularly important for actors, similar effects could occur in “elite professions that require a public presentation of self, such as (...) public officials” (Lincoln and Allen 2004: 613). Hence, female politicians may have a lower likelihood than male politicians of being offered outside jobs with increasing age. Moreover, as “women’s internalization of societal and institutional ageism has helped to perpetuate their marginality” (Bronstein 2001: 184; Walker et al. 2007), women may also become less likely than men to put themselves forward for consideration and/or accept such offers with increasing age. This leads to our fourth and final hypothesis.

**H4:** *Age is more disadvantageous to female than male politicians in obtaining sideline activities.*

### **3. Moonlighting in Germany: Background and Data**

Members of the German national parliament (*Bundestag*) are legally allowed to moonlight, but are subject to public disclosure rules. Specifically, they are obliged to report all professional activities—pursued in addition to their political mandate in the current

legislature or within the last two years before their election into parliament—that fall into the following categories:<sup>6</sup>

- (1) “paid activities in addition to the mandate” as self-employed or employee;
- (2) “activities as a member of the management board, supervisory board, administrative board, advisory board or another board in a private enterprise”;
- (3) corresponding activities in “local authorities or public corporations”;
- (4) corresponding activities in “clubs, associations and foundations”;
- (5) “agreements on future activities or pecuniary advantages”;
- (6) “investments in business companies” when the MP has a voting share of more than 25%.<sup>7</sup>

The declared sideline activities were published for the first time on September 17, 2007 in the Official Handbook of the German Bundestag (German Bundestag 2007), and have been continuously updated on the Bundestag website since then. Below, we employ the data published in the print version of the Official Handbook since these had a uniform due date for all parliamentarians. This provides us with information on the moonlighting activities of all 614 members of the German Bundestag in the period 10/2005 to 09/2007 (197 female and 417 male members). We thereby focus on the sideline activities that MPs exert simultaneously to their *current* political mandate (and disregard occupational activities in the two years *before* entering the German Bundestag).

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<sup>6</sup> These new disclosure rules came into effect in October 2005 (hence the start of our sample period).

<sup>7</sup> Furthermore, for each sideline job, MPs are obliged to report ancillary income exceeding 1,000 Euro (gross) per month or 10,000 Euro (gross) per year. However, these revenues are only published in the form of three income levels: 1000—3500 Euro, 3500—7000 Euro and above 7000 Euro. Moreover, MPs are free to indicate whether they reach a given income level on a monthly or yearly basis. Clearly, this introduces significant uncertainty about actual income generated from each activity. Hence, we do not employ these income data in the analysis (see Rosenson 2007, for a similar decision on similarly reported US data).

The number of ancillary activities for each MP is counted as the sum of activities in the above-mentioned categories (1) through (4) and (6).<sup>8</sup> To address the variety in ancillary activities, we separate activities MPs perform in the private sector of the economy from those performed in the public sector. The former is measured by the sum of ancillary activities in categories (2) and (6), and adding activities from category (1) when the data indicate that the activity is *not* performed for organizations in the political system (e.g., party leader, parliamentary state secretary, federal minister, and so on). Sideline activities in the public sector are measured as the sum of ancillary activities in categories (3) and (4), and adding activities from category (1) when the data indicate that the function *is* performed for an organization in the political system.

-- *Figure 1 about here* --

Figure 1 presents these data in the form of three histograms. In each, the number of MPs is given on the vertical axis, while the number of sideline jobs is on the horizontal axis. Figure 1a shows the distribution for the overall number of sideline jobs MPs hold, while Figures 1b and 1c give the number of MPs' sideline jobs in the private and public sector, respectively. From Figure 1, it is clear that 12.1 % of the 614 members of the German Bundestag (= 74 MPs) have no sideline activities requiring publication, while a few parliamentarians exhibit a large number of ancillary activities (up to a maximum of 24 activities). Sideline jobs are more common in the public sector. Indeed, 304 MPs report to have no ancillary activities in the private sector, whereas only 117 MPs report the same about secondary activities in the public

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<sup>8</sup> We exclude category (5) as we are interested in MPs' *current* sideline activities. Note, however, that only one MP declares an agreement on *future* activities or pecuniary advantages.

sector. The maximum number of sideline jobs is also higher in the public sector (24) than in the private sector (14).<sup>9</sup>

## 4. Empirical Analysis

### 4.1 Regression Model and Methodology

To empirically test our four hypotheses on gender effects in the recruitment market for moonlighting politicians, we estimate a regression model of the following form (with subscript  $i$  referring to MPs):

$$Y_i = \alpha + \beta_1 X_i + \beta_2 GENDER_i + \sum_{k=3}^6 \beta_k (GENDER_i * Z_k) + \varepsilon_i, \quad (1)$$

where  $Y_i$  represents a vector of three dependent variables: namely, total number of ancillary activities (*JOB*), ancillary activities in the private sector (*ECO*) and ancillary activities in the public sector (*POL*) (as depicted in Figure 1).  $X_i$  is a set of control variables describing the characteristics of  $MP_i$ 's political mandate (party affiliation, party-list mandate vs. direct mandate, home federal state, number of legislative periods attended), whether or not  $MP_i$  holds an important political office (i.e. minister, secretary of state, leader of his/her parliamentary fraction, committee leadership or (vice-)president of the Bundestag), and his/her socio-demographic characteristics (age, marital status, number of children, educational background, religious affiliation, occupational background). The central variable of interest is  $GENDER_i$ , an indicator variable equal to 1 for female MPs, 0 for males. This allows testing

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<sup>9</sup> Unfortunately, the data do not indicate the time MPs invest in ancillary activities, such that we implicitly have to assume all jobs take up an equal amount of time. Also, the data are self-reported and we have to assume they are complete and correct. This, however, is not unlikely since violations of the disclosure rules are punishable by a fine of up to half the MPs' annual allowance.

whether women—compared to men—have fewer sideline jobs, especially in the private sector (i.e.  $\beta_2 < 0$ ; see hypotheses H1a and H1b). Then, to test hypotheses H2a, H2b, H3 and H4, we introduce interaction terms between this gender-variable and a set of four background characteristics (contained in  $Z_i$ ; which is a subset of  $X_i$ ): number of legislative periods attended, educational background (dummy = 1 when MP obtained university degree, 0 otherwise), number of children and age. These thus evaluate whether, and how, these characteristics play a different role for men and women.<sup>10</sup> As we have clear, directional hypotheses concerning these interaction effects, we evaluate them using one-tailed significance tests (Thomas 1997).

Regarding the choice of the appropriate statistical technique, it has to be noted that our three dependent variables only take non-negative integer values, and have a highly skewed distribution. This invalidates the use of standard linear regression techniques (i.e. OLS). Moreover, tests for overdispersion in our dependent variables indicate that the variance of these variables is significantly higher than their mean ( $p < 0.01$  in all cases), implying that the Poisson distribution is likewise inappropriate. Hence, a negative binomial count model is employed to estimate equation (1).

#### 4.2 Empirical Results

Table 1 displays the main regression results with respect to our central variables of interest: namely, gender and the gender-interaction effects (results including control variables are relegated to appendix A to preserve space). Columns (1) through (3) contain results for the pure gender effect, while Columns (4) through (6) also include the interaction effects between MPs' gender and background characteristics. In both cases, we report results for the total number of outside jobs (*JOB*; Columns (1) and (4)), the amount of private-sector jobs (*ECO*;

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<sup>10</sup> Since  $Z_i$  is a subset of  $X_i$ , all constituent terms of the interaction terms are likewise included in the analysis.

This is important since their exclusion may lead to significant bias (Braumoeller 2004; Brambor et al. 2006).

Columns (2) and (5)), and the number of public-sector jobs (*POL*; Columns (3) and (6)). Before we turn to the results, a quick note on the interpretation of the coefficients in the interaction model. Essentially, in these estimations—i.e. Columns (4) through (6)—the coefficient of the ‘un-interacted variable’ refers to the effect of the background characteristic of interest (i.e. age, education, and so on) under the situation that  $gender=0$ . Hence, they represent the effect for men. The effect of these background characteristics for women equals the coefficient of the ‘un-interacted variable’ (e.g., age) *plus* the coefficient of the ‘interacted variable’ (e.g., age\*gender). The coefficient on the interaction-variable itself (e.g., age\*gender) indicates to what extent women differ from men; that is, it shows how the effect of, say, age changes when the gender-dummy switches from 0 to 1.

-- Table 1 about here --

Column (1) of Table 1 confirms that, *ceteris paribus*, women on average hold less sideline jobs than men. As shown in Columns (2) and (3), this effect is completely due to their lower number of sideline jobs in the private sector. That is, while gender has a strong and statistically significant negative relation to the number of MPs’ sideline activities in the private sector, its relation with the number of ancillary positions in the public sector remains insignificant. Although our data, unfortunately, do not allow to disentangle whether this derives from women being less likely to be offered certain positions or being more likely to reject them when offered (we return to this important interpretation issue in the conclusion), these results are in agreement with our hypotheses H1a and H1b.

There is also support for hypothesis H2a. Indeed, Columns (4) through (6) illustrate that a university degree has a negative association with the number of sideline activities for men, both in the public and private sector. While initially surprising, one potential explanation is that higher-educated parliamentarians could be more likely to decline offers for outside

activities. Given their higher education, they have a higher probability of already having amassed wealth and/or extensive social networks by the time they enter parliament, reducing their need to accept—possibly numerous—outside jobs as an insurance mechanism against the uncertainty of parliamentary life. Crucially, however, the interaction term between university education and the gender dummy (1 = female) is, in line with our theoretical predictions, consistently positively signed. Moreover, it is statistically significant at conventional levels when analysing MPs overall number of sideline activities (it, however, fails to reach significance when specifying the type of sideline activities in more detail), suggesting higher-level education is significantly more important for women than men to obtain sideline activities. Higher education—and the stock of human capital that comes with it—thus indeed appears more important for women than men to signal competence.

Turning to hypothesis H2b, we find that the number of parliamentary terms—our proxy for political experience—has an overall positive effect for men (though this effect is only statistically significant for public-sector sideline jobs). The interaction term with the gender dummy (1 = female) is once again consistently positively signed, and statistically significant for overall ancillary positions and private-sector jobs. This provides strong support for hypothesis H2b, indicating that parliamentary experience might be equally important for men and women when it comes to public-sector jobs, but is significantly more important for women when it concerns private-sector jobs. The human and, especially, social capital developed through terms in parliament thus appears to have particular value—or are particularly necessary—for women as compared to men.

Note that the findings for parliamentary experience cannot simply be explained by the higher average age of such parliamentarians. In fact, and supportive of our hypothesis H4, age is weakly positively related to the overall number of outside jobs for men, but has a weakly negative effect for women. The interaction term between gender and age is statistically significant and negative, indicating that age has a significantly different effect on men and

women. This effect is especially strong for private-sector sideline jobs, where age has a strong and statistically significant positive effect for men (0.025;  $p < 0.01$ ), but a negative (though insignificant) effect for women (i.e.  $0.025 - 0.031 = -0.006$ ;  $\text{Chi}^2 = 0.14$ ;  $p = 0.70$ ). Overall, our findings suggest that while age might well be an asset for men, it tends to be a burden (or, at least, is a significantly less valuable asset) for women. This is corroborative of a “double standard of aging” (Sontag 1979: 462) or “gendered ageism” (Itzin and Phillipson 1995).<sup>11</sup>

Finally, while the interaction term Children\*Gender is negative as predicted under hypothesis H3, it remains statistically insignificant. Nonetheless, calculating the marginal effects of children on men and women, we see that the total effect of children for women (i.e.  $0.094 - 0.049 = 0.045$ ) is statistically insignificant ( $\text{Chi}^2 = 0.65$ ;  $p = 0.42$ ) whereas it is associated with significantly more outside jobs for men (0.094;  $p < 0.01$ ). The same basic relation shows up in both Columns (5) and (6), though it is slightly stronger for private-sector outside employment. These results imply that while men have more outside jobs when they have children—possibly to generate more income to support the family—women do not (at least not significantly so). One possible explanation is that the men in our sample are more likely than the women to have a spouse at home investing significant time in the household (since highly qualified women are more likely to be part of dual career couples than highly qualified men). As such, men might simply have more time available for outside activities

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<sup>11</sup> Under the (possibly overly restrictive) assumption that employers view the hiring of politicians mainly as a way to cultivate long-term relationships with politicians, this finding might also represent a generational effect. Indeed, given the historical over-representation of men in politics, older female members of the current Bundestag may simply not have been MP at the age when recruitment for outside positions would then mainly occur. If so, the age effect we observe should disappear in the future as later generations of female MPs may have acquired and continue to hold outside positions (see Maume, 2004, for a similar argumentation when analysing wage inequality).

than women when the family expands.<sup>12</sup> Alternatively, traditional gender roles—imposing childcare on women—may provide a stronger counterweight on women than men to the need to generate more income.

## **5. Conclusion**

Normative discussions about the pros and cons of MPs' extra-parliamentary activities often overlook the fact that not all politicians engage in moonlighting. Such debates likewise frequently ignore that the ones who do, do so to a widely differing extent. This paper took a first step to address this gap in the social science literature by investigating the—thus far largely neglected—determinants of moonlighting activity. Obtaining testable hypotheses from the literatures on the glass-ceiling effect and gender stereotyping, we thereby focused on the effect of MPs' gender on observable moonlighting behaviour. As such, we also complemented extensive research documenting the effects of gender on purely political careers (e.g., Rosenwasser and Seale 1988; Sanbonmatsu 2002; Rahat 2007; Galligan and Clavero 2008) and on private-sector hiring and promotion decisions.

Our empirical results—using data on 197 female and 417 male German MPs over the period 10/2005-09/2007—indicated that gender indeed matters strongly for the number of politicians' outside jobs (controlling for party affiliation, religion, political position and other potential factors of influence). Women are, on average, less likely to have many outside jobs, especially in the private sector. Moreover, political experience and university education are more important determinants of moonlighting for women than men, while age and children work more forcefully against women than men. These findings illustrate that, even when analysing women of higher-than-average ability working in a strongly politicised

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<sup>12</sup> Interestingly, these results complement recent findings that having children “reduces the [work-related] travel activity of women, whereas there is no consistent such effect among men” (Gustafson 2006: 513).

environment, gender stereotypes continue to have a manifest influence on recruitment decisions.

It is important, however, to abstain from normative judgements based on our empirical findings. Although gender discrimination or employers' pro-male bias is one possible interpretation, another might be that women are less likely than men to seek and/or accept offers of (especially private-sector) sideline activities. That is, assuming such offers are equally distributed across gender (i.e. no gender-based discrimination), our observations would likewise come about when women are less likely than men to accept them. In such interpretation, women are not discriminated against, but rather could be seen as being more conscientious about adequately fulfilling their parliamentary work, more risk averse (when outside jobs are viewed in a negative light by the general public and therefore entail potential electoral risks), or more scrupulous in avoiding potential conflicts of interest than men.

Unfortunately, as mentioned above, our data do not allow disentangling these various interpretations because we only observe the final distribution of outside jobs across politicians, not the number of offers sought, made and/or accepted/rejected. In future work, it would therefore be of interest to complement the current analysis with a more detailed evaluation—for example, via (semi-)structured interviews—of participants at both sides of the studied recruitment market (in terms of application behaviour, job offering, acceptance/rejection decisions, and so on). Rather than focus on the market outcome (i.e. MPs' observed outside positions), this provides a possibility to look more directly into the recruitment *process* itself. As such, a more comprehensive view on how gender (consciously or subconsciously) affects this decision-making process can be reached. Additionally, such approach would also allow more direct assessment of the potential influence of 'old-boy networks' in recruitment processes concerning politicians' outside jobs (see also footnote 3 above).

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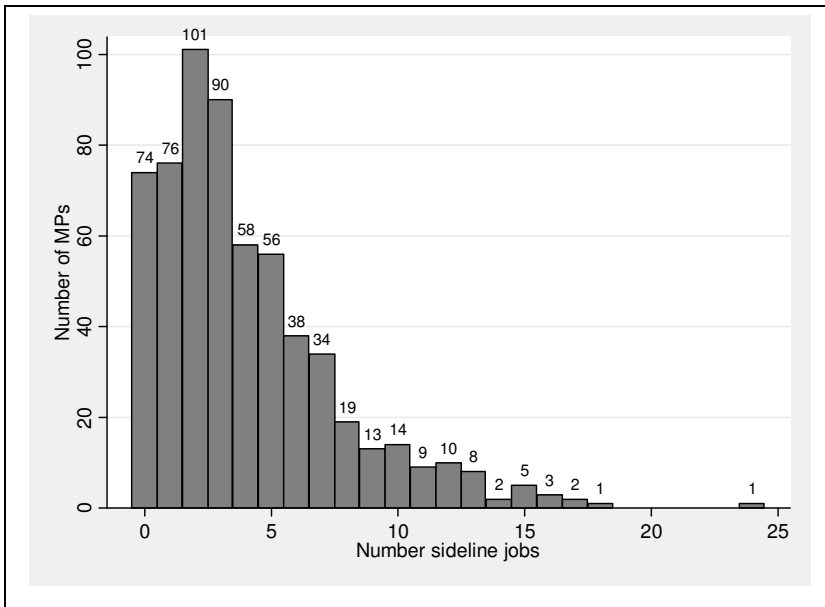
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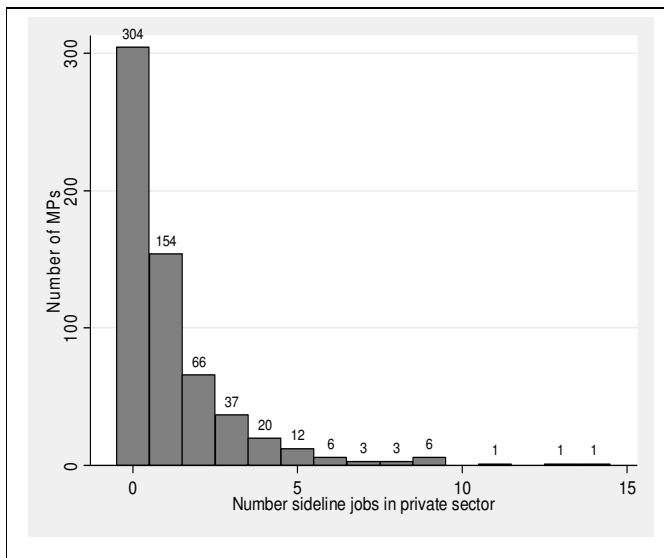
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**Figure 1: Frequency Distribution of Sideline Jobs (overall, in private and public sector)**

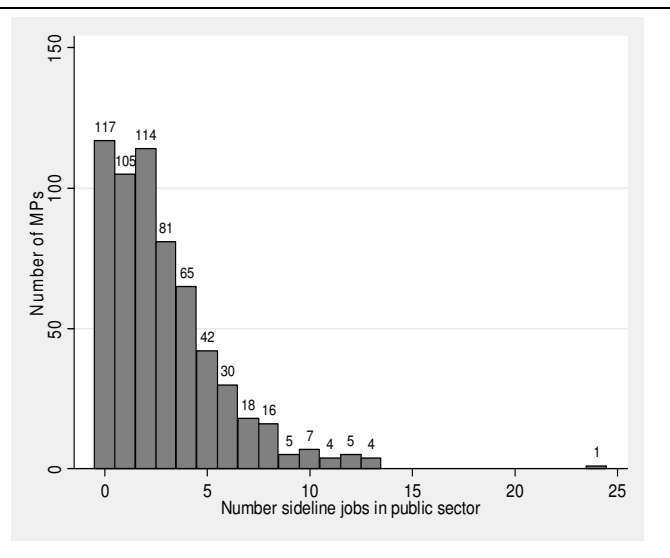
**Figure 1a (overall)**



**Figure 1b (private sector)**



**Figure 1c (public sector)**



**Table 1: Determinants of MPs' Sideline Activities**

<i>Variable</i>	<i>JOB</i> <i>(1)</i>	<i>ECO</i> <i>(2)</i>	<i>POL</i> <i>(3)</i>	<i>JOB</i> <i>(4)</i>	<i>ECO</i> <i>(5)</i>	<i>POL</i> <i>(6)</i>
Intercept	1.445 *** (6.44)	0.126 (0.34)	1.197 *** (4.80)	1.343 *** (5.50)	-0.024 (-0.06)	1.193 *** (4.35)
University degree	-0.208 *** (-2.31)	-0.367 ** (-2.40)	-0.163 * (-1.57)	-0.342 *** (-3.14)	-0.521 *** (-2.92)	-0.259 *** (-2.00)
Terms in parliament	0.051 *** (2.35)	0.021 (0.61)	0.064 *** (2.64)	0.032 (1.29)	-0.007 (-0.19)	0.055 ** (1.93)
Number of children	0.079 *** (2.83)	0.150 *** (3.13)	0.055 ** (1.72)	0.094 *** (3.05)	0.167 *** (3.24)	0.068 ** (1.83)
Age	-0.001 (-0.18)	0.018 *** (2.49)	-0.009 ** (-1.65)	0.004 (0.74)	0.025 *** (3.06)	-0.008 (-1.23)
Gender	-0.130 * (-1.63)	-0.774 *** (-5.13)	0.075 (0.87)	0.162 (0.40)	-0.115 (-0.14)	0.032 (0.08)
Gender * Children	-	-	-	-0.049 (-0.82)	-0.083 (-0.62)	-0.039 (-0.60)
Gender * Terms	-	-	-	0.090 ** (1.81)	0.197 *** (2.45)	0.036 (0.67)
Gender * Age	-	-	-	-0.152 ** (-1.68)	-0.031 ** (1.70)	-0.004 (-0.41)
Gender * University	-	-	-	0.331 ** (1.86)	0.515 * (1.40)	0.238 (1.26)
Log pseudolikelihood	-1455.94	-835.62	-1288.51	-1451.03	-830.82	-1287.14
Wald Chi <sup>2</sup> (R)	131.39 ***	160.68 ***	99.23 ***	139.67 ***	168.93 ***	105.18 ***

Note: N=614; t-values based on robust standard error between brackets: \*\*\* significant at 1%; \*\* at 5% and \* at 10% (one-tailed). Wald-test indicates joint-significance of all regressors in the model (with R equal to the number of regressors minus one). Controls for party affiliation, party-list mandate vs. direct mandate, home federal state, important political offices and MPs' marital status, religious affiliation and occupational background included in all regressions (see full regression results in appendix A).

## Appendix A

**Table A1: Full Results**

<i>Variable</i>	<i>JOB</i> <i>(1)</i>	<i>ECO</i> <i>(2)</i>	<i>POL</i> <i>(3)</i>	<i>JOB</i> <i>(4)</i>	<i>ECO</i> <i>(5)</i>	<i>POL</i> <i>(6)</i>
Intercept	1.445 *** (6.44)	0.126 (0.34)	1.197 *** (4.80)	1.343 *** (5.50)	-0.024 (-0.06)	1.193 *** (4.35)
SPD	-0.308 *** (-3.62)	-0.810 *** (-5.12)	-0.159 * (-1.73)	-0.308 *** (-3.63)	-0.833 *** (-5.39)	-0.157 ** (-1.70)
FDP	0.010 (0.08)	-0.196 (-1.06)	0.040 (0.27)	0.009 (0.07)	-0.200 (-1.08)	0.039 (0.26)
Green	-0.108 (-0.64)	-1.145 *** (-3.91)	0.134 (0.75)	-0.125 (-0.75)	-1.158 *** (-3.90)	0.123 (0.70)
Die Linke	-0.670 *** (-3.45)	-0.891 *** (-3.17)	-0.676 *** (-2.76)	-0.688 *** (-3.57)	-0.909 *** (-3.23)	-0.682 *** (-2.80)
Independent	-0.481 *** (-1.97)	-1.113 ** (-1.77)	-0.328 (-0.50)	-0.557 *** (-1.99)	-1.254 *** (-2.09)	-0.366 (-0.55)
Mandate	0.086 (1.10)	-0.044 (-0.32)	0.136 * (1.63)	0.076 (0.96)	-0.053 (-0.39)	0.127 * (1.50)
East	-0.222 * (-1.63)	-0.284 * (-1.63)	-0.201 ** (-1.73)	-0.236 *** (-2.47)	-0.301 ** (-1.74)	-0.209 ** (-1.87)
Catholic	0.017 (0.16)	-0.576 *** (-3.25)	0.188 ** (1.65)	0.013 (0.12)	-0.581 *** (-3.34)	0.186 * (1.62)
Protestant	0.191 *** (2.15)	-0.230 * (-1.42)	0.325 *** (-3.29)	0.208 ** (2.36)	-0.213 * (-1.34)	0.336 *** (-3.39)
Married	-0.141 ** (-1.65)	-0.135 (-0.92)	-0.143 * (-1.55)	-0.135 * (-1.58)	-0.132 (-0.89)	-0.139 * (-1.51)
University degree	-0.208 *** (-2.31)	-0.367 *** (-2.40)	-0.163 * (-1.57)	-0.342 *** (-3.14)	-0.521 *** (-2.92)	-0.259 *** (-2.00)
Economics background	0.080 (0.80)	0.208 (1.26)	0.020 (0.20)	0.113 (1.12)	0.253 * (1.52)	0.038 (0.37)
Legal background	0.025 (0.30)	0.225 * (1.54)	-0.036 (-0.39)	0.050 (0.58)	0.254 ** (1.71)	-0.022 (-0.23)
Teacher background	-0.041 (-0.37)	-0.415 ** (-1.72)	0.071 (0.59)	-0.044 (-0.39)	-0.425 ** (-1.80)	0.073 (0.61)
Terms in parliament	0.051 *** (2.35)	0.021 (0.61)	0.064 *** (2.64)	0.032 (1.29)	-0.007 (-0.19)	0.055 ** (1.93)
Number of children	0.079 *** (2.83)	0.150 *** (3.13)	0.055 ** (1.72)	0.094 *** (3.05)	0.167 *** (3.24)	0.068 ** (1.83)
Age	-0.001 (-0.18)	0.018 *** (2.49)	-0.009 ** (-1.65)	0.004 (0.74)	0.025 *** (3.06)	-0.008 (-1.23)
Minister	0.134 (0.96)	-0.838 *** (-2.22)	0.344 *** (-2.27)	0.142 (0.99)	-0.835 *** (-2.19)	0.342 *** (-2.23)
Secretary of state	0.238 ** (1.73)	-0.249 (-0.90)	0.398 *** (2.99)	0.221 ** (1.65)	-0.280 (-1.07)	0.390 *** (2.86)
Member party fraction leadership	-0.013 (-0.08)	-0.048 (-0.21)	0.011 (0.14)	-0.023 (-0.15)	-0.063 (-0.27)	0.016 (0.10)
Leader party fraction	0.497 *** (2.45)	0.540 * (1.56)	0.529 ** (1.92)	0.487 *** (2.35)	0.492 * (1.44)	0.519 ** (1.84)
Leader committee	0.113 * (1.62)	0.197 ** (1.69)	0.115 * (1.59)	0.105 * (1.51)	0.185 * (1.58)	0.112 * (1.53)
Substitute leader committee	0.179 ** (1.75)	-0.101 (-0.46)	0.276 *** (2.67)	0.223 *** (2.10)	-0.044 (-0.20)	0.304 *** (2.84)
Vice-president	0.643 ***	-0.648 **	0.814 ***	0.640 ***	-0.622 *	0.816 ***

Bundestag	(2.41)	(-1.64)	(2.82)	(2.30)	(-1.49)	(2.79)
Gender	-0.130 * (-1.63)	-0.774 *** (-5.13)	0.075 (0.87)	0.162 (0.40)	-0.115 (-0.14)	0.032 (0.08)
Gender * Children	-	-	-	-0.049 (-0.82)	-0.083 (-0.62)	-0.039 (-0.60)
Gender * Terms	-	-	-	0.090 ** (1.81)	0.197 *** (2.45)	0.036 (0.67)
Gender * Age	-	-	-	-0.152 ** (-1.68)	-0.031 ** (1.70)	-0.004 (-0.41)
Gender * University	-	-	-	0.331 ** (1.86)	0.515 * (1.40)	0.238 (1.26)
Log pseudolikelihood	-1455.94	-835.62	-1288.51	-1451.03	-830.82	-1287.14
Wald Chi <sup>2</sup> (R)	131.39 ***	160.68 ***	99.23 ***	139.67 ***	168.93 ***	105.18 ***

Note: N=614; t-values based on robust standard error between brackets: \*\*\* significant at 1%; \*\* at 5% and \* at 10% (one-tailed). Wald-test indicates joint-significance of all regressors in the model (with R equal to the number of regressors minus one).