levels of trade and subsequent levels of government consumption expenditure. The coefficient

of -.97 for Tradeit−1 ∗MDIit−1 in Model 4 suggests that, on average, a two standard deviation

increase in media density (109.3 points) in the long-run decreases the estimated effect that a two standard deviation increase in trade openness (85.2% of GDP) would have on government spending by .97% of GDP. Model 6 estimates that this negative conditioning effect of media density is as little as .61% of GDP. Moving from the minimum density of media (0) to the maximum in the sample (313.3) is associated, on average, with a decrease of as much as

2.78%, and as little as 1.8%, in the expected effect of an 85% increase in trade openness on government spending as a share of GDP. Although the estimated effect appears relatively slight, it should be kept in mind that the mean level of government consumption expenditure in the sample is only 15.7% of GDP. Thus, for a country that begins with no mass media and becomes as fully penetrated as the most penetrated (the United States in 1986), the roughly 1-3% of GDP by which we would expect the country to reduce its compensatory public spending in the long-run for an 85% increase in trade openness is a substantial portion of what a typical country spends.

Models 5 and 6 also tests the conditioning effect of media density against the conditioning effects of democracy on the trade-welfare relationship, which previous research has found to increase the redistributive responsiveness of domestic welfare spending to international trade (Adserà and Boix 2002). The results here suggest that media density has a robust conditioning effect on the relationship between trade and spending, while the interaction found by Adserà and Boix no longer appears statistically distinguishable from zero.

To check for the possibility that the above models are spuriously driven by some dif- ferent but unobserved process distinct from the effects of mass media, I gather additional data to test my state-level argument against a series of rival explanations. Specifically, it is argued that left parties and union density are aspects of the domestic institutional environ- ment which lead to more redistributive responses to economic liberalization (Garrett 1995, 674); that electoral systems defined by proportional representation are more redistributive

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than majoritarian systems (Iversen and Soskice 2006); and that the degree of unitarism or government centralization affects welfare spending (Crepaz 1998, 72). Finally, of particular interest in the literature relating economic globalization to the politics of welfare is the argu- ment of Iversen (2001) and Iversen and Cusack (2000) that deindustrialization rather than globalization has driven the expansion of welfare spending since the 1960s.

In Table 2, I re-estimate the error-correction model (as in Model 6 of Table 1) controlling for each of the rival explanations above.21 The interaction of trade levels and media density levels is robust to the inclusion of each potentially confounding variable, suggesting that the conditioning effect of media density on the trade-spending relationship is not a spurious cor- relation due to an omitted variable. In an additional model not displayed for lack of space, I

included the terms Tradeit−1∗GDPPerCapitait−1 and ∆Tradeit−1∗∆GDPPerCapitait−1 on

the right-hand side of the equation in the fashion of the models in Table 2; neither coefficient

is statistically significant and the main independent variable of interest, Tradeit−1 ∗MDIit−1,

remains signed as expected and statistically significant (β=-1.56, standard error=.000, panel- corrected standard error=.005).

Several additional robustness checks were conducted. With panel data, Nickell bias can lead to rejecting a true null hypothesis when the number of units is large relative to the time period (Gaibulloev, Sandler, and Sul, 2014). Given that the present panel contains a relatively large number of countries with a sampling period of more than 30 years (the period beyond which Nickell bias becomes ignorable), I re-estimate the models in Table 2 on only those countries with complete data for the 38 years between 1961 and 1999. Table 12 in Sup- plementary Information shows the results, substantially the same as the those reported here.

I also checked sensitivity to lag specification for Tradeit−1 ∗ MDIit−1 estimating the models in Table 1 using Tradeit ∗ MDIit as the independent variable instead. The coefficients and

standard errors are not substantially changed (see Table 13 in Supplementary Information). Thefore the state-level models provide additional, robust evidence that mass media shape

21See Supporting Information for more information on variable descriptions and sources.

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| Table 1: Determinants of Government Consumption Expenditure |
| Model 1 Model 2 Model 3 |
| Tradeit−1 0.14 0.15 0.32 |
| (0.21) (0.21) (0.23) |
| ∆Tradeit−1 −0.88∗∗  (0.41) |
| MDIit−1 −0.16 −0.11 0.70∗∗ |
| (0.40) (0.40) (0.33) |
| ∆MDIit−1 1.01  (1.76) |
| Democracyit−1 0.03 −0.03 −0.11  (0.15) (0.16) (0.16) |
| GDPPerCapitait−1 0.62∗∗∗ 0.61∗∗∗ |
| (0.19) (0.19) |
| Dependencyit−1 −0.22 −0.24  (0.23) (0.23) |
| LandAreait−1 13.92 20.36 |
| (191.65) (191.68) |
| ∆Democracyit−1 0.04  (0.30) |
| ∆GDPPerCapitait−1 0.64  (0.73) |
| ∆Dependencyit−1 2.59  (1.86) |
| Spendingit−1 0.72∗∗∗ 0.72∗∗∗ −0.19∗∗∗  (0.02) (0.02) (0.02) |
| Spendingit−2 0.11∗∗∗ 0.11∗∗∗ |
| (0.02) (0.02) |
| Spendingit−3 0.03  (0.02) |
| ∆Spendingit−1 −0.09∗∗∗  (0.02) |
| Tradeit−1 \* MDIit−1 −0.97∗∗∗ −0.88∗∗∗ −0.71∗∗ |
| (0.32) (0.33) (0.33) |
| ∆Tradeit−1 \* ∆MDIit−1 −0.29  (16.85) |
| Tradeit−1 \* Democracyit−1 −0.43 −0.45 |
| (0.30) (0.31) |
| ∆Tradeit−1 \* ∆Democracyit−1 −3.53∗  (2.01) |
| N 3911 3911 3785 |
| R-squared 0.67 0.67 0.11 |
| Adj. R-squared 0.64 0.64 0.10 |
| ∗∗∗p < .01; ∗∗p < .05; ∗p < .1 |

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